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Ben Schumann

Declaration of Co-Authorship and Publications

This dissertation consists of four (working) papers. The contribution in conception, implementation, and drafting of the four chapters can be summarized as follows:

1. Crude Controversy - Energy prices and European monetary policy
   Contribution: 100%

2. What goes around comes around: How large are spillbacks from US monetary policy?
   with Max Breitenlechner and Georgios Georgiadis
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Summary

Chapter 1: In this paper, I analyze the interplay between (European) monetary policy and energy prices. Employing a Bayesian proxy structural vector autoregressive model, I establish that the ECB's decisions have material effects on global and local energy prices. This starkly contrasts the public communication and internal assumptions of the ECB. Through Lucas-critique robust counterfactuals, I demonstrate that the monetary-policy-induced changes in energy prices play a crucial role in shaping the response of inflation and inflation expectations to monetary policy. By affecting fast-moving energy prices, monetary policy transmits quickly and more strongly to consumer prices. This turns energy prices from a foe to a friend that, when managed correctly, can assist monetary policy in achieving its objective of price stability. Finally, I ask how European monetary policy should optimally respond to an energy price shock and find that, historically, it has been too accommodative. My estimates suggest that the ECB could have largely avoided the latest energy-price-driven surge in inflation of 10% at the cost of a short-run and comparably small loss in output. I argue that this favorable outcome is precisely due to the ability of the ECB to affect fast-moving energy prices.

Chapter 2: This paper, which is joint work with Max Breitenlechner and Georgios Georgiadis, deals with the domestic repercussions of the global effects of the Federal Reserve's decisions. In particular, it shows that these "spillovers" from US monetary policy entail "spillbacks" to the domestic economy. Applying counterfactual analyses in a Bayesian proxy structural vector-autoregressive model we find that spillbacks account for a non-trivial share of the slowdown in domestic real activity following a contractionary US monetary policy shock. Spillbacks materialize as a monetary policy tightening depresses foreign sales and valuations of US firms so that Tobin's q/cash flow and stock market wealth effects impinge on investment and consumption. Net trade does not contribute to spillbacks because US monetary policy affects exports and imports similarly. Geographically, spillbacks materialize through advanced rather than emerging market economies.

Chapter 3: In this paper, which is joint work with Gernot Müller and Georgios Georgiadis, we analyze the interplay of changes in global risk and the appreciation of the US dollar. We identify global risk shocks using high-frequency asset-price surprises around narratively selected events. Global risk shocks appreciate the US dollar, induce tighter global financial conditions and a synchronized contraction of world economic activity. To isolate the role played by the appreciation of the US dollar we benchmark the estimated effects of

these global risk shocks against counterfactuals in which the US dollar does not appreciate. By leveraging recent advances in sufficient statistics approaches to macroeconomic policy evaluation and building a rich two-country DSGE model we show that, in the absence of US-dollar appreciation, the contractionary impact of a global risk shock is much weaker. This holds true in the rest of the world as well as the US. For the rest of the world, contractionary financial channels thus dominate expansionary expenditure switching effects when global risk rises and the US dollar appreciates.

Chapter 4: In this paper Georgios Georgiadis and I develop the partial and asymmetric dominant-currency pricing (DCP) hypothesis and test for its empirical relevance. This hypothesis states that a large but not necessarily identical share of global export and import prices are sticky in US dollars and that this impacts the response of an economy to unexpected changes in the US dollar. We first set up a structural three-country New Keynesian dynamic stochastic general equilibrium model which nests DCP, producer-currency pricing (PCP), and local-currency pricing (LCP). Under DCP, the output spillovers from shocks that appreciate the US dollar decline with an economy's export-import US dollar pricing share differential, i.e. the difference between the share of an economy's export and import prices that are sticky in US dollar. Underlying this prediction is variation in an economy's net exports in response to US dollar appreciation that arises because the shares of export and import prices that are sticky in US dollar are different. We then document that this prediction from partial and asymmetric DCP is consistent with the data. We do so by estimating impulse responses to different shocks that appreciate the US dollar for a sample of up to 45 advanced and emerging market economies. We document that our findings are robust to considering US demand, US monetary policy, and exogenous exchange rate shocks as a trigger of US dollar appreciation, zooming in on the responses of economies' exports and imports, as well as accounting for the role of commodity trade in US dollar invoicing.

Zusammenfassung

Kapitel 1: In diesem Artikel erforsche ich die Wechselwirkung zwischen der (europäischen) Geldpolitik und den Energiepreisen. Mithilfe eines Bayesianischen, strukturellen vektorautoregressiven Modell zeige ich, dass die Entscheidungen der EZB wesentliche Auswirkungen auf die globalen und lokalen Energiepreise haben. Dies steht in starkem Gegensatz zur öffentlichen Kommunikation und den internen Annahmen der EZB. Anhand von Lucas-Kritik robusten, kontrafaktischen Simulation zeige ich, dass die geldpolitisch bedingten Veränderungen der Energiepreise eine entscheidende Rolle für die Auswirkungen der Geldpolitik auf die Inflation der Inflationserwartungen spielen. Ich dokumentiere empirisch, dass sich die Geldpolitik durch die Beeinflussung der sich schnell verändernden Energiepreise schneller und stärker auf die Verbraucherpreise überträgt. Dadurch verändern sich die Energiepreise von einer lästigen Komplikation zu einem Werkzeug, welches, wenn es richtig eingesetzt wird, der Geldpolitik helfen kann, ihr Ziel der Preisstabilität zu erreichen. Anschließend stelle ich die Frage, wie die europäische Geldpolitik optimal auf einen Energiepreisschock reagieren sollte. Ich lege empirisch dar, dass sie in der Vergangenheit nicht kontraktiv genug auf diese Schocks reagiert hat. Meine Schätzungen ergeben, dass die EZB den jüngsten energiepreisbedingten Inflationsanstieg von 10% hätte vermeiden können, hätte sie eine restriktivere Zinspolitik implementiert. Der Preis für das Erfüllen ihres Mandats wäre ein lediglich kurzfristiger und vergleichsweise geringer Produktionsverlust gewesen. Ich zeige, dass dieses günstige Ergebnis auf die Fähigkeit der EZB zurückzuführen ist, die sich schnell verändernden Energiepreise zu beeinflussen.

Kapitel 2: Dieses Papier, welches ich gemeinsam Max Breitenlechner und Georgios Georgiadis verfasst habe, befasst sich mit den inländischen Auswirkungen der globalen Effekte der Entscheidungen der Federal Reserve. Es zeigt, dass diese "Spillovers" der US-Geldpolitik "Spillbacks" auf die inländische Wirtschaft nach sich ziehen. Durch die Anwendung neuester Methoden zur Berechnung von kontrafaktischen, makroökonomischen Szenarien kommen wir zu dem Schluss, dass diese Spillbacks einen nicht-trivialen Anteil an der beobachteten Verlangsamung der inländischen Realwirtschaft nach einem kontraktiven geldpolitischen Schock in den USA ausmachen. Diese Effekte entstehen dadurch, dass eine kontraktive Geldpolitik der Federal Reserve den Auslandsabsatz von US-Firmen verringert, was sich negativ auf die inländischen Investitionen auswirkt. Zudem sinken die Aktienpreise von inländischen ausländischen Firmen, wodurch US-Konsumenten, welche diese Wertpapiere halten, ihren Konsum einschränken. Der Außenhandelt trägt nicht zu Spillbacks bei, da

die US-Geldpolitik Exporte und Importe in gleicher Weise beeinflusst. Geographisch gesehen werden Spillbacks eher durch die Auswirkungen der US-Geldpolitik auf entwickelte Volkswirtschaften ausgelöst.

Kapitel 3: In diesem Artikel, welchen ich gemeinsam mit Gernot Müller und Georgios Georgiadis geschrieben habe, analysieren wir das Zusammenspiel des Dollars und globalen Finanzmarktrisiken. Der US-Dollar ist eine Safe-Haven-Währung und wertet auf, wenn das globale Risiko zunimmt. Wir untersuchen welche Rolle diese Aufwertung des Dollars bei der Übertragung globaler Finanzmarktrisiken auf die Weltwirtschaft spielt. Wir identifizieren zunächst globale Risikoschocks mit Hilfe von hochfrequenten Veränderungen von Vermögenswerten zum Zeitpunkt von narrativ ausgewählten Ereignissen. Globale Risikoschocks werten den US-Dollar auf, führen zu einer Verschärfung der globalen Finanzbedingungen und zu einem globalen Rückgang der Wirtschaftsleistung. Wir vergleichen diese Effekte mit kontrafaktischen Szenarien, in denen der US-Dollar nicht aufwertet. Wir benutzen die jüngsten Fortschritte bei den Ansätzen der hinreichenden Statistik zur Berechnung makroökonomischer, kontrafaktischer Szenarien nutzen und ein umfangreiches Zwei-Länder-DSGE-Modell um zu zeigen, dass die kontraktiven Auswirkungen eines globalen Risikoschocks ohne die daraus hervorgehende Aufwertung des Dollars wesentlich schwächer ausfallen. Dies gilt sowohl für den Rest der Welt als auch für die USA. Für den Rest der Welt überkompensieren also bei einer Aufwertung des US-Dollars die kontraktive Finanzmarktkanäle die expansiven Handels- und Wettbewerbsfähigkeitskanäle.

Kapitel 4: In diesem Papier entwickeln Georgios Georgiadis und ich die partielle und asymmetrische dominant-currency pricing (DCP) Hypothese und testen ihre empirische Relevanz. Diese Hypothese besagt, dass ein großer, aber nicht unbedingt gleicher Teil der globalen Exporte und Importe in US-dollar gepreist werden und dass dies die Reaktion einer Volkswirtschaft auf unerwartete Veränderungen des US-Dollars beeinflusst. Zunächst stellen wir ein Neukeynesianisches, strukturelles, dynamisches, stochastisches Drei-Länder-Modell auf, welches DCP, producer-currency pricing (PCP) und local-currency pricing (LCP) beinhaltet. Wir zeigen analytisch, dass unter unserer partiellem und asymmetrischem dominant-currency pricing, die Spillover-Effekte von US-Dollar aufwertenden Schocks, von der Differenz zwischen dem Anteil der in US-Dollar gepreisten Exporte- und Importe einer Volkswirtschaft abhängen. Anschließend weisen wir empirisch nach, dass diese Vorhersage aus der partiellen und asymmetrischen DCP Hypothese mit den Daten konsistent ist. Dazu werden zunächst Impulsantworten auf verschiedene, US-dollar aufwertende Schocks für eine Stichprobe von bis zu 45 fortgeschrittenen und aufstrebenden Marktwirtschaften geschätzt. Anschließend untersuchen wir empirisch, ob die Auswirkungen dieser Schocks auf unterschiedliche Länder von dem besagten Differenzial zwischen dem Anteil der in US-Dollar in Rechnung gestellten Exporte- und Importe abhängen. Wir zeigen, dass unsere Ergebnisse robust sind, wenn man die US-Nachfrage, die US-Geldpolitik und exogene Wechselkursschocks als Auslöser für die Aufwertung des US-Dollars betrachtet und die Rolle des US-Dollars für den globalen Rohstoffhandels berücksichtigt.

Introduction and overview

Over the past centuries, globalization has not only reshaped the world and affected the daily lives of billions of people, but the unprecedented amounts of goods, services, and capital flowing across borders also intertwined the fate of economies around the world with the state of global economic conditions. As these are affected by changes in economic activity and economic policies in large open economies such as the United States and the Euro Area, the world economy now dances to the tune of policymakers in those regions (Miranda-Agrippino and Rey (2020b), Miranda-Agrippino and Nenova (2022)). But this extraordinary privilege also implies global responsibilities, which necessitates a careful analysis of (i) why and how policies in these large open economies transmit to other economies around the globe, and (ii) how this ultimately alters their effect on the domestic economy.

An answer to the first question opens the door for global policy cooperation and would allow policymakers to design a more resilient global financial system that is tailored to take into account the vulnerabilities arising from increasingly tight global interlinkages. For example, after the Global Financial Crisis, some policymakers complained that US monetary policy measures aimed at stabilizing the domestic economy elicited waves of capital flows and accentuated financial market volatility in the rest of the world (Rajan 2013). Understanding why the Federal Reserve plays such an outstanding role in the global economy and what defines if and how a foreign country is affected by changes in the Federal Reserve's policy stance is key to understanding the aforementioned vulnerabilities. Chapters three and four of this dissertation speak to these topics by analyzing the implications of the dominant role of the currency that is under the direct control of the Federal Reserve, the US dollar.

An answer to the second question contributes to making economic policy advice fit for the challenges of the 21st century where textbook examples of closed economies are largely a thing of the past, implying that the corresponding policy conclusions may have to be rewritten. This is particularly true for large open economies because economic policies in these regions spill over to the global economy, which in turn implies economic spillbacks to the domestic economy. As argued by Fischer (2014), one cannot make "sensible" policy choices without taking these effects into account. The first two chapters of this dissertation therefore empirically analyze how the domestic transmission of monetary policy in large open economies is altered by its corresponding global repercussions.

The **first chapter** focuses on the Euro Area. In particular, it questions the narrative that the ECB's decisions do not affect energy prices (see Lagarde (2022b), Christoffel et al. (2008)) and analyzes how the ability of the ECB to do so changes transmission and optimal conduct of Euro Area monetary policy. Building on Ider et al. (2023) I first estimate a Bayesian proxy structural vector autoregressive model, to establish that the ECB's decisions have material effects on global and local energy prices. Through Lucas-critique robust counterfactuals, I then isolate the crucial role of energy prices in shaping the response of domestic inflation and inflation expectations to changes in the ECB's policy stance. I show that, by affecting fast-moving energy prices, monetary policy transmits quickly and more strongly to consumer prices. I then ask how European monetary policy should optimally respond to an energy price shock and find that, historically, it has been too accommodative. My estimates suggest that the ECB could have largely avoided the latest energy-price-driven surge in inflation of 10% at the cost of a short-run and comparably small loss in output. I show that this favorable outcome can be attributed to the ability of the ECB to affect fast-moving energy prices, allowing it to tackle the root cause of the economic disruption directly. This turns energy prices from a foe to a friend that, when managed correctly, can assist monetary policy in achieving its objective of price stability.

The **second chapter**, which is joint work with Max Breitenlechner and Georgios Georgiadis, then turns the attention to the United States. In particular, we analyze how the United States' position as the center of the global economy affects the transmission of the Federal Reserve's monetary policy decisions to the domestic economy. Much empirical work, as well as prominent policy debates, suggest that US monetary policy spillovers are large and an important driver of business cycles and financial conditions in the global economy (Dedola et al. 2017; Bräuning and Sheremirov 2019; Iacoviello and Navarro 2019; Degasperi et al. 2020). Against this background, some policymakers have argued the Federal Reserve should internalize its effects on the rest of the world in the calibration of its monetary policy (Rajan 2016). The Federal Reserve has responded that it already does so implicitly: "The Fed recognizes that its own policies do have international spillovers, and, in turn, because they affect global performance, they are going to have spillbacks to US economic performance" (Yellen 2019). However, to the best of our knowledge, no rigorous analysis of spillbacks from US monetary policy exists in the literature. This paper fills this gap by applying counterfactual analyses in a Bayesian proxy structural vector-autoregressive model. We find that spillbacks account for a non-trivial share of the slowdown in domestic real activity following a contractionary US monetary policy shock. As such, by observing its effect on the domestic economy, the Federal Reserve already internalizes large parts of its global impact.

These pronounced spillback effects materialize because the decisions of the Federal Reserve entail large spillovers to the rest of the world. It has long been argued that this is due to spikes in risk in global financial markets that are caused by changes in the Federal Reserve's policy stance (Miranda-Agrippino and Rey (2020b), Miranda-Agrippino and Ricco (2021)). The **third chapter**, which is joint work with Gernot Müller and Georgios

Georgiadis, refines this view and ties together spikes in global risk and the dominance of the US dollar in the global financial system. It is well documented that the US dollar dominates safe asset supply (Krishnamurthy and Lustig (2019)), cross-border bank credit (Bruno and Shin (2015)) and global trade (Gopinath et al. (2020a)). Therefore the US dollar is a safe-haven currency and appreciates when global risk goes up. We investigate the dollar's role in the transmission of global risk to the world economy within a Bayesian proxy structural vectorautoregressive model. The identified global risk shocks appreciate the US dollar, induce tighter global financial conditions and a synchronized contraction of world economic activity. We benchmark these effects against counterfactuals in which the dollar does not appreciate either because the Federal Reserve establishes an exchange rate peg or because the US dollar loses its dominant status. In the absence of a dollar appreciation, the contractionary impact of a global risk shock is much weaker, both in the rest of the world and the US. This implies that the decisions of the Federal Reserve disproportionally impact the global economy not only because they cause changes in risk aversion in financial markets but precisely because the dollar is the dominant global currency.

The **fourth chapter**, which is joint work with Georgios Georgiadis, then analyzes what determines the exposure of individual economies to changes in the dollar with a particular focus on the role of dollar dominance in global trade, a phenomenon called dominant-currency pricing (DCP). More specifically, we test for the empirical relevance of partial and asymmetric DCP, the hypothesis that large but not necessarily identical shares of economies' export and import prices are sticky in US dollar. We first set up a structural three-country New Keynesian dynamic stochastic general equilibrium model which nests DCP, producer-currency pricing, and local-currency pricing. Under partial and asymmetric DCP, the output spillovers from any kind of shock that appreciates the US dollar decline with an economy's export-import US dollar pricing share differential, i.e. the difference between the share of an economy's export and import prices that are sticky in US dollar. Underlying this prediction is a variation in an economy's net exports in response to US dollar appreciation that arises because the shares of export and import prices that are sticky in US dollar are different. We then provide evidence that this prediction from partial and asymmetric DCP is consistent with the data. We do so by estimating impulse responses to different shocks that appreciate the US dollar for a sample of up to 45 advanced and emerging market economies for the period from 1995 to 2018. Our hypothesis is robust to considering US demand, US monetary policy, and exogenous exchange rate shocks as a trigger of US dollar appreciation, zooming in on the responses of economies' exports and imports, as well as accounting for the role of commodity trade in US dollar invoicing. Therefore a country's dollar invoicing shares constitute an additional indicator for its exposure to the US dollar, which lies at the heart of the global economic system.

Chapter 1

Crude Controversy - Energy prices and European monetary policy

Abstract

In this paper, I delve into the interplay between (European) monetary policy and energy prices and uncover their role in the transmission of monetary policy. Employing a Bayesian proxy structural vector autoregressive model, I establish that the ECB's decisions have material effects on global and local energy prices. Through Lucas-critique robust counterfactuals, I isolate the crucial role of energy prices in shaping the response of inflation and inflation expectations to monetary policy. I show that, by affecting fast-moving energy prices, monetary policy transmits quickly and more strongly to consumer prices. This turns energy prices from a foe to a friend that, when managed correctly, can assist monetary policy in achieving its objective of price stability. Finally, I ask how European monetary policy should optimally respond to an energy price shock and find that, historically, it has been too accommodative.

Keywords: inflation, energy prices, monetary policy transmission mechanism, optimal policy
JEL Codes: C22, E31, E52, Q43

1.1 Introduction

Inflation in the euro area has been steadily on the rise since 2021, reaching unprecedented levels in 2022. This surge can be primarily attributed to the sharp increase in energy prices. Despite inflation rising at a rate not seen for 40 years, there has been debate about if and how the ECB should respond. On the one hand, there were doubts about the ECB's ability to fight inflation caused by high energy prices. One argument frequently raised by policymakers was that the ECB cannot influence energy prices, which are determined by the world market (Lagarde (2022b)). On the other hand, even if the ECB could influence energy prices, there were doubts as to whether it should act because the costs in terms of output and employment could outweigh the effects on prices (Lagarde (2022a)). In this paper, I question both these conjectures by empirically analyzing the influence of European monetary. By employing a Bayesian proxy SVAR model I corroborate the findings of Ider et al. (2023) who show that contractionary ECB policies imply a fall in the prices of energy goods traded on the global market as well as energy prices faced by consumers in the euro area.

Second, I examine the importance of energy prices for the transmission of monetary policy in the euro area using a counterfactual scenario computed along the lines of McKay and Wolf (2023). In particular, I base the assessment of this importance on an empirical counterfactual, where the ECB's decisions do not affect global energy prices. I document that the HICP and inflation expectations respond much less to changes in the ECB's policy stance when its decisions do not impact global oil prices. Intuitively, by affecting fast-moving, relatively flexible energy prices, monetary policy has a tighter grip on short- and medium-term inflation dynamics.

Third, given the insight that the ECB can fight energy-price-driven surges in inflation, I ask whether euro area monetary policy should actually do so. Using the method to empirically conduct optimal-policy counterfactuals proposed by McKay and Wolf (2023), I find that if conventional monetary policy and forward guidance were used optimally to target medium-term price stability, the ECB should respond more aggressively to an oil price shock than it did historically. This curbs the response of inflation while only leading to a slightly deeper front-loaded recession.

Fourth, I document that the degree of additional tightening required for the optimal policy strategy crucially depends on the response of global energy prices to a change in the ECB's policy stance. In particular, if (dollar) prices of energy goods traded on the global market would not respond, the optimal strategy would entail further tightening especially at the longer end of the yield curve. But even in this case, the optimal policy strategy succeeds in curbing inflation without having to engineer a large recession. The reason is that the additional tightening quickly and strongly appreciates the euro, which in turn quickly transmits into lower local energy prices. Therefore, if handled correctly, energy prices can be the ECB‘s friend, not its foe.

Lastly, I ask the question of how the recent energy-price-driven inflation surge would have played out if, counterfactually, the ECB had committed to (optimally) stabilize medium-term inflation. I first show that the latest policy decisions during that period indeed contributed to the surge in energy prices and inflation, which already calls into question the proclaimed “looking through” strategy of the ECB (Schnabel (2022)). I then employ the recently developed method of Caravello et al. (2024), who extend the approach of McKay and Wolf (2023) to historical episodes, to show that the optimal “medium-term inflation targeting” strategy would have entailed a strong initial tightening and would have prevented the surge in inflation at the costs of approximately 4% lower output in the short-run.

In more detail, I first employ a slightly modified version of the Bayesian proxy SVAR (BPSVAR) model of Ider et al. (2023) to estimate the dynamic effects of euro area monetary policy shocks on energy prices, inflation, and inflation expectations. In line with their results, this analysis reveals that contractionary euro area monetary policy shocks significantly reduce prices of energy goods traded on the global market as well as energy prices faced by consumers in the euro area. Intuitively, and in line with the theoretical work by Auclert et al. (2023), contractionary monetary policy transmits to energy prices faced by euro area consumers by (i) affecting global energy prices (in dollars) due to a change in domestic and global demand and (ii) by appreciating the euro exchange rate resulting in a fall in prices for imported energy goods. Ider et al. (2023) refer to the role played by changes in energy prices for the monetary transmission as the “energy-price channel of monetary policy”.

Second, I examine the importance of energy prices for the transmission of monetary policy in the euro area using a counterfactual exercise. In particular, I base the assessment on an empirical counterfactual, where the ECB’s decisions — as frequently stated in multiple of its press conferences — do not affect global energy prices. More precisely, in the counterfactual exercise, OPEC’s policy rule is such that it aims to stabilize deviations of global oil prices from its preferred path. Consequently, in this counterfactual world, European monetary policy does not influence the global oil price. The estimation of this counterfactual is based on the method by McKay and Wolf (2023) which takes into account the anticipatory effects of such an OPEC policy rule change. I approximate the solution to the counterfactual question using the “best Lucas-critique robust approximation” of the true counterfactual in an empirical model (McKay and Wolf 2023, p. 1698). I furthermore accompany this approximation technique with two additional approaches and provide a detailed discussion of their respective advantages and disadvantages. To implement these approaches I borrow from the literature on high-frequency identification of oil supply shocks (Känzig (2021)) to jointly identify a short- and a medium-run oil supply news shock as well as a euro area monetary policy shock within the BPSVAR model. Under the counterfactual OPEC policy rule, which aims to stabilize the global (dollar) price of oil, the corresponding impulse response functions to a euro area monetary policy shock display a much weaker response of euro-denominated energy prices faced by euro area consumers. Crucially, also the harmonized index of consumer prices (HICP) and inflation expectations — arguably the ultimate objectives of the ECB — respond much less when the ECB's decisions do not impact global oil prices. Comparing the counterfactual responses to the baseline reveals that, by affecting fast-moving energy prices which are known to be relatively flexible compared to other goods in the HICP basket, monetary policy has a tighter grip on inflation dynamics especially in the short- to medium-term.

Third, I ask the question of whether euro area monetary policy should actually fight energy-price-driven surges in inflation. This question is particularly interesting for two reasons. First, the previous analysis revealed that the ECB can in principle counteract rapidly rising energy prices by increasing interest rates. Second, I document that, in line with many of their public statements, the ECB historically followed a "looking-through" strategy (Schnabel (2022)), implying that they historically did not increase interest rates in response to the inflationary energy price shock. To answer the question of which strategy should be preferred I use the McKay and Wolf (2023) framework to explore how the ECB, given its mandate, should optimally respond to an oil price shock. I implement this approach by employing multiple instruments developed by the literature on high-frequency identification of euro area monetary policy shocks to jointly identify a conventional monetary policy shock and a forward guidance shock alongside an oil price shock within the BPSVAR model (Altavilla et al. (2019)). I find that, if conventional monetary policy and forward guidance were used optimally to target medium-term price stability, the ECB should respond more strongly to an oil price shock than it did historically. This leads to a slightly deeper front-loaded contraction in output. However, it also succeeds in curbing inflation almost completely. The reason is that the more contractionary policy stance under optimal policy lowers global energy prices quickly, which then results in a smaller increase in the CPI and almost none in inflation expectations. Consequently, only a small additional output loss is necessary to minimize medium-term inflation deviations from the target, since it leads to sizeable movements in relatively more flexible energy prices.

Fourth, I empirically characterize the optimal response of the ECB to an oil price shock if, in line with the aforementioned statements by ECB officials, the ECB's decision would not exert pressure on global energy prices. I compute the answer to this thought experiment by combining the approach of McKay and Wolf (2023) to compute counterfactual impulse responses with their approach to estimating the optimal policy response. The findings suggest that the optimal response to an oil price shock would markedly differ under this counterfactual scenario. If the claims that the ECB's decisions do not affect oil prices were true, the optimal strategy would entail a substantially stronger tightening, especially at the longer end of the yield curve. Interestingly, even in this case, the optimal policy decision would not require engineering a large recession to exert pressure on the prices of domestically produced goods. It rather implies the use of some additional tightening at the longer end of the yield curve to strongly appreciate the euro, which in turn lowers fast-moving local energy prices, again highlighting the pivotal role of energy prices in the conduct of monetary policy.

Lastly, I estimate how the recent energy-price-driven inflation surge would have played out if, counterfactually, the ECB had committed to (optimally) stabilize medium-term inflation. I first show that the latest policy decisions during that period indeed contributed to the surge in energy prices and inflation, which calls into question the proclaimed "looking through" strategy of the ECB (Schnabel (2022)). I then employ the recently developed method of Caravello et al. (2024) to estimate Lucas-Critique robust historical counterfactuals, which allows me to quantify the additional tightening and output losses that would have been necessary to (optimally) stabilize inflation at the medium-term target. I document that, by switching from a "looking-through" to a committed "medium-term inflation targeting" strategy, the ECB would have been able to stabilize inflation at roughly 2% at the cost of GDP being approximately 4% in the short run. Furthermore, I show that it is possible to decompose the mitigated inflation of approximately 7.5% relative to the observed path into two components, with the first (second) one capturing the change in the propagation of shocks occurring before (after) April 2021. Regarding the first one, I estimate that the optimal strategy would have entailed a strong initial tightening at the beginning of 2021 which would have been required to tame the underlying inflationary forces resulting from the fading out of the pandemic. This additional front-loaded tightening explains roughly 3% of the mitigated inflation. The remaining 4.5% can be traced back to the fact that the switch from a "looking-through" to a committed "medium-term inflation targeting" alters the propagation of structural shocks occurring after April 2021, which in part have been shocks to energy prices. As shown in the previous sections, the ECB's decisions affect global energy prices and therefore a change in the ECB's strategy can help to alter and tame the (global) propagation of energy price shocks.

Related literature. First, the paper contributes to the extensive literature that studies the transmission channels of monetary policy (Christiano et al. (1999), Gertler and Karadi (2015), Miranda-Agrippino and Ricco (2021)). Relative to existing work I focus on a largely neglected transmission channel. In particular, I analyze how energy prices respond to a monetary policy shock and how this shapes the transmission of monetary policy to the economy. Therefore, the paper relaxes the prevalent assumption that energy prices are exogenous (see for instance Christoffel et al. (2008)) and adds to the literature that studies the effects of monetary policy on commodity prices (Degasperi et al. (2020), Miranda-Pinto et al. (2023), Ca'Zorzi et al. (2023)). In contrast to much of the existing work, I follow Ider et al. (2023) focus on the effect of European monetary policy on energy prices and document that the ECB's decisions, similar to those of the Federal Reserve, affect energy prices on the global and euro area levels.

Second, my paper speaks to the literature that studies the transmission of energy price shocks conditional on the reaction of monetary policy. There exists ample literature on the macroeconomic effects of oil price shocks (Kilian (2009), Baumeister and Hamilton (2019), Känzig (2021)) and whether the response of monetary policy exacerbates the effects of these shocks (Bernanke et al. (1997), Leduc and Sill (2004); Kilian and Lewis (2011), Miyamoto et al. (2023)). Although the existing literature provides insights about both questions, the novel contribution of this paper is the empirical analysis of the optimal monetary policy response to an oil price shock, which until recently could only be studied with theoretical models (Bodenstein et al. (2012), Natal (2012)). To the best of my knowledge, this paper is the first to empirically estimate the ECB's optimal monetary policy response to an oil price shock.

Lastly, this paper is related to the recent literature on the use of sufficient statistics approaches in macroeconomic policy evaluation (Barnichon and Mesters (2023), McKay and Wolf (2023), Barnichon and Mesters (2024), Caravello et al. (2024)). This paper is not only the first one to apply these approaches to evaluate euro area monetary policy but also does so in a fully coherent Bayesian framework. Furthermore, it demonstrates how the approach of McKay and Wolf (2023) can be used to learn about the importance of expectations in monetary transmission. It also shows how to extend the method of Caravello et al. (2024) to decompose the implications that a change in the policy rule has for the evolution of the economy into a component that can be related to the way policy alters the transmission of unexpected shocks and a component related to the expected evolution of the economy.

The paper is structured in the following way. Building on the Bayesian Proxy SVAR of Ider et al. (2023) Section 1.2 investigates how euro area monetary policy shocks affect prices of energy goods traded on the global market as well as energy prices faced by euro area consumers. Section 1.3 describes the framework on how to compute empirical counterfactuals and shows how energy prices matter in the transmission of monetary policy. In Section 1.4, the same framework is used to estimate the optimal policy response by the ECB to an energy price shock. Section 1.5 analyzes the role played by monetary policy in the recent energy price surge and estimates how the economy would have evolved if the ECB had optimally targeted medium-term inflation. The final section concludes.

1.2 The effects of euro area monetary policy on energy prices

The goal of this section is to provide a bridge between this paper and the analysis of Ider et al. (2023) who investigate if and how decisions of the ECB translate into changes in energy prices traded (in dollars) on the global market and energy prices faced by euro-area consumers (in euros). To this end, I estimate a slightly modified version of the Bayesian Proxy SVAR (BPSVAR) model of Ider et al. (2023) which serves as a starting point and foundation for the analysis in this paper.

1.2.1 The Bayesian Proxy SVAR model

This section briefly lays out the BPSVAR model with the exposition closely following Georgiadis et al. (2024) and Ider et al. (2023). Although the main goal of this section is to identify the impulse responses to a single euro area monetary policy shock, I keep the notation general, as in later sections of the paper I aim to identify k structural shocks of interest with $k \geq 1$ proxy variables.

Following the notation of Rubio-Ramirez et al. (2010), consider without loss of generality the structural VAR model with one lag and without deterministic terms

$$\boldsymbol{y}_t' \boldsymbol{A}_0 = \boldsymbol{y}_{t-1}' \boldsymbol{A}_1 + \boldsymbol{\epsilon}_t', \qquad \boldsymbol{\epsilon} \sim N(\mathbf{0}, \boldsymbol{I}_n), \tag{1.1}$$

where $\boldsymbol{y}_t$ is an $n \times 1$ vector of endogenous variables and $\boldsymbol{\epsilon}_t$ an $n \times 1$ vector of structural shocks. The BPSVAR framework builds on the following assumptions in order to identify k structural shocks of interest: There exists a $k \times 1$ vector of proxy variables $\boldsymbol{m}_t$ that are (i) correlated with the k structural shocks of interest $\boldsymbol{\epsilon}_t^*$ and (ii) orthogonal to the remaining structural shocks $\boldsymbol{\epsilon}_t^o$. Formally, the identifying assumptions are

$$E[\boldsymbol{\epsilon}_t^* \boldsymbol{m}_t'] = \underset{(k \times k)}{\boldsymbol{V}}, \tag{1.2a}$$

$$E[\boldsymbol{\epsilon}_t^o \boldsymbol{m}_t'] = \underset{((n-k) \times k)}{\mathbf{0}}, \tag{1.2b}$$

and represent the relevance and the exogeneity condition, respectively.

I estimate the BPSVAR model using the algorithm developed in Arias et al. (2021), where they estimate (1.1) augmented with equations for the proxy variables. Denote by $\tilde{\boldsymbol{y}}_t' \equiv (\boldsymbol{y}_t', \boldsymbol{m}_t')$, by $\tilde{\boldsymbol{A}}_\ell$ the corresponding $\tilde{n} \times \tilde{n}$ coefficient matrices with $\tilde{n} = n + k$, and by $\tilde{\boldsymbol{\epsilon}} \equiv (\boldsymbol{\epsilon}_t', \boldsymbol{v}_t')' \sim N(\mathbf{0}, \boldsymbol{I}_{n+k})$, where $\boldsymbol{v}_t$ is a $k \times 1$ vector of measurement errors. The augmented structural VAR model is then given by

$$\tilde{\boldsymbol{y}}_t' \tilde{\boldsymbol{A}}_0 = \tilde{\boldsymbol{y}}_{t-1}' \tilde{\boldsymbol{A}}_1 + \tilde{\boldsymbol{\epsilon}}_t'. \tag{1.3}$$

The algorithm of Arias et al. (2021) imposes the assumptions (1.2a) and (1.2b) in the estimation of (1.3) to identify the structural shocks. I relegate the details of the algorithm to Appendix A.2 and discuss here the advantages of this Bayesian approach relative to the standard frequentist two-step estimation. First, it allows me to refrain from imposing potentially contentious recursiveness assumptions between the endogenous variables when multiple structural shocks are identified using multiple proxy variables (Mertens and Ravn (2013)). Second, the single-step estimation of the BPSVAR model is more efficient than the standard two-stage least squares estimation of proxy SVAR and facilitates coherent inference. In fact, the Bayesian set-up allows exact finite sample inference and does not require an explicit theory to accommodate potentially weak instruments. Third, the BPSVAR framework allows the proxy variables to be serially correlated, predictable, and affected by measurement error. Lastly, Bayesian inference is particularly convenient in the presence of set identification, which potentially arises when identifying multiple shocks using multiple proxies.[1]

1.2.2 Data and specification

The baseline monetary VAR model for the euro area closely resembles the model of Ider et al. (2023) and contains eight variables and additionally, a high-frequency surprise series to identify an ECB monetary policy shock. Specifically, the baseline version of the model for the euro area includes the 1-year constant maturity German Bund yield as a measure of the monetary policy stance. Economic activity is measured by the euro area industrial production index (excluding construction). Financial conditions are represented by the BBB corporate bond spread. I employ the Harmonized Index of Consumer Prices (HICP) as a measure of the overall price level, and its energy component as the measure of local energy prices in the euro area. The Brent crude oil price acts as a stand-in for the global energy price and I also add the EUR-USD exchange rate, since oil and other energy goods are usually traded in US dollars. Crucially, and in contrast to Ider et al. (2023), I employ the Consensus one-year-ahead inflation expectations to analyze the interplay between energy prices and inflation expectations as it is documented that the latter heavily influences the former (Aastveit et al. (2023)). I utilize the commonly employed high-frequency changes in the 3-month OIS rate over the monetary event window from Altavilla et al. (2019) database as my preferred proxy for the euro area monetary policy shocks and apply the "poor man's approach" of Jarocinski and Karadi (2020) to purge these surprises from central bank information effects based on the sign of the corresponding equity-price surprise. All variables are measured at monthly frequency. Furthermore, all variables, except for interest rates and credit spreads, enter the SVAR in log-levels ($\times 100$) such that one can interpret impulse responses as percentage deviations. Further details on the data can be found in Appendix A.1.

The BPSVAR model is estimated on a sample from January 1999 to December 2019, thus leaving out the extraordinary volatility in the data induced by the COVID-19 pandemic.[2] The model has 12 lags and includes a constant. I use uninformative priors when estimating the BPSVAR parameters.[3] In addition, a relevance threshold is imposed to express my

[1] I fully acknowledge the concerns that in the case of set identification, the uniform prior for the rotation matrix, which is embedded in the approach of Arias et al. (2021), may even asymptotically influence the results as forcefully raised by Baumeister and Hamilton (2019) and Giacomini and Kitagawa (2021). But recent contributions by Inoue and Kilian (2021) and Arias et al. (2023) called into question the empirical relevance of this concern in applied research with tightly identified sets as is the case in these applications. Therefore I conduct standard Bayesian inference along the lines of Rubio-Ramirez et al. (2010) and the subsequent literature.

[2] I include and explicitly model this period along the lines of Cascaldi-Garcia (2022) in an extension in section 1.5. Results are robust extending the model along those dimensions as shown in Figure A.2 in the Appendix

[3] As in Born and Pfeifer (2021) and many other studies I impose the dogmatic prior that the SVAR is stable implying that, after being hit by an exogenous shock, the endogenous variables eventually converge

prior belief that the proxy is to some extent informative about the true monetary policy shocks. In particular, I impose a threshold that the identified structural monetary policy shocks account for at least 10% of the variance in the proxy.[4]

1.2.3 Dynamic effects of a monetary policy shock

Figure 1.1 presents the baseline estimates of the effects of a one standard deviation contractionary monetary policy shock for the euro area. The 1-year Bund yield increases on impact by 5 basis points and quickly reverts to zero, with an overall shape and magnitude that is very similar to Jarocinski and Karadi (2020). Industrial production falls by approximately 0.35% on impact and remains depressed for about 1.5 years. Likewise, the fall in the domestic headline consumer price level is immediate, reaching a trough of about 0.08% after about 18 months. Inflation expectations follow a very similar pattern. Turning to the exchange rate, as expected, the euro appreciates against the dollar by a little less than 1% and remains significantly elevated for a year.

So far all of the estimated dynamics for the endogenous variables described are arguably in line with the results from standard theory and corroborate previous findings in the literature (Jarocinski and Karadi (2020)). Crucially and in line with the findings in Ider et al. (2023) contractionary euro area monetary policy also leads to a sizable fall in the measures for local and global energy prices. The global oil price (in US dollars) falls strongly by approximately 3%. Moreover, the local energy price index, measured by the HICP energy component, falls by 0.5% which is much larger than the fall in the overall price of the HICP basket. In fact, given that energy prices receive a weight of roughly 10% in the overall HICP basket, a back-of-the-envelope calculation suggests that the overwhelming amount of the drop in the overall HICP price level in the short- and medium-run can be attributed to the effect that the contractionary monetary policy shock has on local energy prices.

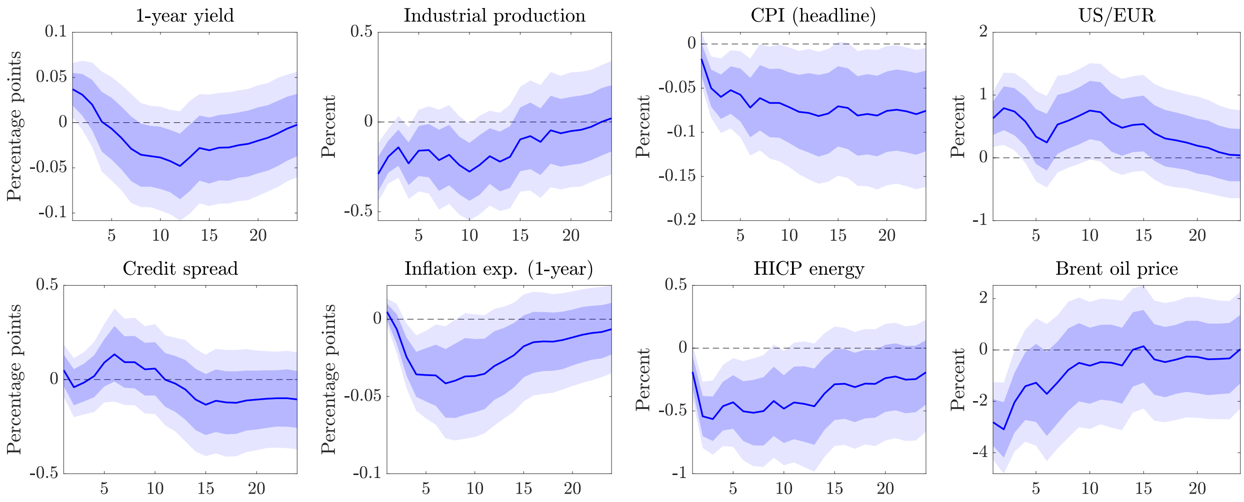
These results corroborate the findings from the micro-data literature that, not only do prices of contracts for energy goods traded on financial markets change on a daily basis, but even at the consumer level these energy goods are by far the goods with the highest frequency of price updating. For instance, when analyzing micro-data used for the computation of the Belgian HICP, Aucremanne and Dhyne (2004) show that the average price duration for energy goods is roughly 1 month which stands in stark contrast to the median price duration of all goods in the basket, which amounts to approximately 14 months. Intuitively, when viewed through the lens of a standard New-Keynesian Model, this implies that, all else equal, the Phillips Curve for energy goods has a steeper slope than the one of an average consumer good.

In summary, the BPSVAR analysis reveals that monetary policy influences not only

back to their steady state.

[4]This is a weak requirement compared to the 20% threshold of Arias et al. (2021) and the 'high-relevance' prior of Caldara and Herbst (2019). As shown in Figure A.1 in the Appendix, the results are robust to reducing the relevance condition to 0.

Figure 1.1: Baseline impulse responses to a euro area monetary policy shock



Notes: Impulse response functions to a one standard deviation monetary policy shock. Point-wise posterior means along with 68% and 90% point-wise credible sets. Horizon in months.

global but also local energy prices in the euro area. Ider et al. (2023) refer to the role played by the respective energy prices in the monetary transmission as the "global" and "local" components of the "energy price channel", which is defined as the joint contribution of these prices to the monetary transmission. Although these results may seem surprising in light of the statements made by ECB officials on multiple occasions (Lagarde (2022b), Lagarde (2022a)), it is arguably less surprising given the fact the euro area is the second largest oil importer for the majority of the sample period only trumped by the US during earlier periods of the sample and very recently by China (see Figure A.14 in the Appendix).

In light of these considerations, I next explore how monetary policy affects the economy in a counterfactual scenario where global energy prices are unaffected by a change in the euro area's monetary policy stance as frequently claimed by ECB officials. In the language of Ider et al. (2023), this implies that the global component of the energy price channel is absent from the monetary transmission.

1.3 The role of (global) energy prices in the transmission of euro area monetary policy

The results in the previous section show that euro area monetary policy has a significant impact on global energy prices and, consequently, energy prices faced by consumers in the euro area. While the impulse response functions alongside the back-of-the-envelope calculations already indicate an important role played by energy prices in monetary transmission, this is arguably only one reading of those responses. Ider et al. (2023) further corroborates this intuition using statistical approaches along the lines of Breitenlechner et al. (2022) to estimate the transmission of a euro area monetary policy shock if global energy prices would not respond. In contrast to their approach, I leverage the recent breakthrough of McKay and Wolf (2023) which allows me to complement their findings by conducting a more structural counterfactual exercise. In particular, I calculate the response to a euro area monetary policy shock under a scenario where OPEC's policy rule, counterfactually, is such that it aims to perfectly stabilize the oil price. As such the oil price does not respond to a change in the ECB's policy stance. An assumption that is not only in line with multiple statements of ECB officials but is also frequently assumed throughout the internal forecasting process (see Christoffel et al. (2008) for a discussion). Already at this stage, it's important to note that the HICP energy index of the euro area consists of a weighted average of local energy prices. Therefore movements in the (global) oil price, despite being the dominant component of this basket, only correspond to a fraction of total euro area consumer energy prices. As such, results from this exercise should be interpreted as a lower bound for the total role of energy prices in the transmission of the ECB's monetary policy decision.[5]

1.3.1 Computing structural (policy) counterfactuals

My approach to estimating impulse responses under the counterfactual OPEC policy rule builds on the recent insights of McKay and Wolf (2023, henceforth MW). In particular, MW develop an approach for constructing policy-rule counterfactuals empirically that is (i) robust to the Lucas critique and (ii) recovers the true policy-rule counterfactual for a broad range of underlying structural frameworks, including standard representative and heterogeneous-agent New Keynesian models. The key ingredients in their counterfactual analysis are impulse responses to shocks about current and future policy. In particular, they show that combining the impulse response function to the structural shock of interest —estimated under the baseline policy rule— with a particular sequence of impulse responses to policy (news) shocks, uncovers the counterfactual impulse response functions that would prevail under a counterfactual policy rule. In other words, impulse responses prove to be sufficient statistic to estimate the desired counterfactual scenario.

Formally, MW consider a linear, perfect-foresight, infinite-horizon economy in terms of deviations from the deterministic steady state for periods $t = 0, 1, 2, \ldots$. In the sequence-space notation of Auclert et al. (2021), this economy can be described by a set of equations

$$\boldsymbol{\mathcal{H}}_x \boldsymbol{x} + \boldsymbol{\mathcal{H}}_z \boldsymbol{z} + \boldsymbol{\mathcal{H}}_\epsilon \boldsymbol{\epsilon} = \boldsymbol{0}, \tag{1.4}$$

$$\boldsymbol{\mathcal{A}}_x \boldsymbol{x} + \boldsymbol{\mathcal{A}}_z \boldsymbol{z} + \boldsymbol{\nu} = \boldsymbol{0}, \tag{1.5}$$

where $\boldsymbol{x} \equiv (\boldsymbol{x}_1', \boldsymbol{x}_2', \ldots, \boldsymbol{x}_{n_x}')'$ stacks the time paths of the n_x endogenous variables over n_h

[5]Below I provide a detailed discussion of the decision to focus on global energy prices and in particular the oil price.

periods, analogously $\boldsymbol{z}$ stacks the time path of the n_z policy instruments. The matrices $\boldsymbol{\mathcal{H}}$ summarize the behavior of agents in the non-policy block, while the matrices $\boldsymbol{\mathcal{A}}$ describe the baseline policy rule of interest. $\boldsymbol{\epsilon}$ represents the n_ϵ non-policy structural shocks and $\boldsymbol{\nu}$ the n_ν policy (news) shocks; the latter are deviations from the policy rule announced at date t but implemented only in some future period $t+i$, $i \geq 0$. The key assumption reflected in Equations (1.4) and (1.5) is that $\{\boldsymbol{\mathcal{H}}_x, \boldsymbol{\mathcal{H}}_z, \boldsymbol{\mathcal{H}}_\epsilon\}$ do not depend on the coefficients of the policy rule $\{\boldsymbol{\mathcal{A}}_x, \boldsymbol{\mathcal{A}}_z\}$, so that policy affects the non-policy block's decisions only through the path of the instrument $\boldsymbol{z}$, rather than through the policy rule *per se*. As shown in McKay and Wolf (2023), this assumption holds true for a broad range of structural frameworks frequently used in counterfactual policy analysis such as standard representative and heterogeneous-agent New Keynesian models.

Under the assumption that the solution exists and is unique, the solution to Equations (1.4) and (1.5) can be written in impulse response space

$$\begin{pmatrix} \boldsymbol{x} \\ \boldsymbol{z} \end{pmatrix} = \boldsymbol{\Theta}_{\mathcal{A}} \times \begin{pmatrix} \boldsymbol{\epsilon} \\ \boldsymbol{\nu} \end{pmatrix}, \quad \boldsymbol{\Theta}_{\mathcal{A}} \equiv (\boldsymbol{\Theta}_{\epsilon,\mathcal{A}}, \boldsymbol{\Theta}_{\nu,\mathcal{A}}) \equiv \begin{pmatrix} \boldsymbol{\Theta}_{x,\epsilon,\mathcal{A}} & \boldsymbol{\Theta}_{x,\nu,\mathcal{A}} \\ \boldsymbol{\Theta}_{z,\epsilon,\mathcal{A}} & \boldsymbol{\Theta}_{z,\nu,\mathcal{A}} \end{pmatrix}. \tag{1.6}$$

where $\boldsymbol{\Theta}_{\mathcal{A}}$ collects the impulse responses of the policy instrument $\boldsymbol{z}$ and the non-policy variables $\boldsymbol{x}$ under the baseline policy rule summarized by $\boldsymbol{A}$.

In the counterfactual analysis below, the goal is to analyze impulse responses under a counterfactual policy rule. The policy block with this counterfactual policy rule is given by:

$$\tilde{\boldsymbol{\mathcal{A}}}_x \boldsymbol{x} + \tilde{\boldsymbol{\mathcal{A}}}_z \boldsymbol{z} = \boldsymbol{0}, \tag{1.7}$$

where $\tilde{\boldsymbol{\mathcal{A}}}_x$ and $\tilde{\boldsymbol{\mathcal{A}}}_z$ contain the corresponding coefficients of the counterfactual rule. MW show that knowledge of the impulse responses $\boldsymbol{\Theta}_{\mathcal{A}}$ under the baseline policy rule is sufficient to determine the impulse responses to the structural shock of interest ϵ under any counterfactual policy rule even without knowing the true underlying structural model that generates the data. In particular, they prove that

$$\boldsymbol{\Theta}_{x,\epsilon,\widetilde{\mathcal{A}}} \times \boldsymbol{\epsilon} \equiv \boldsymbol{\Theta}_{x,\epsilon,\mathcal{A}} \times \boldsymbol{\epsilon} + \boldsymbol{\Theta}_{x,\nu,\mathcal{A}} \times \widetilde{\boldsymbol{\nu}}, \qquad \boldsymbol{\Theta}_{z,\epsilon,\widetilde{\mathcal{A}}} \times \boldsymbol{\epsilon} \equiv \boldsymbol{\Theta}_{z,\epsilon,\mathcal{A}} \times \boldsymbol{\epsilon} + \boldsymbol{\Theta}_{z,\nu,\mathcal{A}} \times \widetilde{\boldsymbol{\nu}}. \tag{1.8}$$

In words, the impulse response to the structural shock ϵ under the counterfactual policy rule described by $\widetilde{\mathcal{A}}$ is exactly equivalent to a linear combination of the corresponding impulse responses under the baseline policy rule described by $\mathcal{A}$ and the impulse responses to a specific sequence of policy news shocks $\widetilde{\boldsymbol{\nu}}$. Intuitively, as long as the decisions of the non-policy block depend on the (expected) path of the policy instrument rather than the rule itself, it does not matter whether the path comes about due to the systematic conduct of policy or due to additional policy news shocks. Consequently, the policy news shocks $\widetilde{\boldsymbol{\nu}}$ are chosen such that, for the path of endogenous variables and policy instruments under the counterfactual policy rule ($\boldsymbol{x}_{\tilde{\mathcal{A}}}, \boldsymbol{z}_{\tilde{\mathcal{A}}}$), this counterfactual policy rule holds

$$\tilde{\boldsymbol{\mathcal{A}}}_x[\boldsymbol{x}_{\tilde{\mathcal{A}}}]+\tilde{\boldsymbol{\mathcal{A}}}_z[\boldsymbol{z}_{\tilde{\mathcal{A}}}] = \tilde{\boldsymbol{\mathcal{A}}}_x\left[\boldsymbol{\Theta}_{x,\epsilon,\mathcal{A}} \times \boldsymbol{\epsilon} + \boldsymbol{\Theta}_{x,\nu,\mathcal{A}} \times \tilde{\boldsymbol{\nu}}\right]+\tilde{\boldsymbol{\mathcal{A}}}_z\left[\boldsymbol{\Theta}_{z,\epsilon,\mathcal{A}} \times \boldsymbol{\epsilon} + \boldsymbol{\Theta}_{z,\nu,\mathcal{A}} \times \tilde{\boldsymbol{\nu}}\right] = \mathbf{0}. \tag{1.9}$$

What is left to determine are the expressions $\boldsymbol{\Theta}_{x,\nu,\mathcal{A}}$ and $\boldsymbol{\Theta}_{z,\nu,\mathcal{A}}$ in Equation (1.8). In theory, it would require knowledge of impulse responses to news shocks which communicate changes in the future level of the policy instrument for all possible n_h horizons. All of these are necesarry to construct the full matrix $\boldsymbol{\Theta}_{z,\nu,\mathcal{A}}$. However, in practice, it is difficult if not often impossible to estimate impulse responses to policy news shocks for all n_h periods. I tackle the problem in the following way. I start by stacking the system in Equation (1.9) across all the responses of all the $n = n_x + n_z$ endogenous variables $\boldsymbol{x}$ and the policy instrument $\boldsymbol{z}$ and all horizons $n_H = n \times n_h$ in order to arrive at:

$$\tilde{\boldsymbol{\mathcal{A}}}\boldsymbol{\Theta}_{\epsilon,A} \times \boldsymbol{\epsilon} + \tilde{\boldsymbol{\mathcal{A}}}\boldsymbol{\Theta}_{\nu,\mathcal{A}} \times \tilde{\boldsymbol{\nu}} = 0, \tag{1.10}$$

Afterwards, I approximate the solution to Equation (1.10) using three different techniques, which despite their seemingly different nature, are all nested in the stacked sequence space representation in Equation (1.10). In particular, I employ (i) the “best Lucas-critique-robust approximation” (McKay and Wolf 2023, p.5) of the problem as proposed by MW, which amounts to choosing the counterfactual policy shocks $\tilde{\boldsymbol{\nu}}$ as to minimize the squared deviations from the counterfactual policy rule, (ii) the procedure of Sims and Zha (2006), which has a long-standing tradition in the SVAR literature and which I show is nested in the representation in Equation (1.10), and (iii) a “hybrid” version of the two approaches which is briefly sketched in the Online Appendix of McKay and Wolf (2023).

Least-Squares-approximation: The first approach to computing an approximation of the solution to Equation (1.10) is the one advocated by MW and which aims to have the counterfactual policy rule hold at point 0 and for all n_h periods ahead expectations. This is reminiscent of a policy counterfactual in a fully structural (DSGE) model, where the agents are forward-looking and know exactly the structure of the economy and therefore form consistent expectations about the path of policy instrument.

MW discuss the approximation of the solution to Equation (1.9) in the situation where the researcher has only a subset of the required policy news shocks ($\boldsymbol{\nu}^{\text{LSQ}} \in \boldsymbol{\nu}$) available. This constrains the space of allocations that can be implemented via policy changes and thus necessitates an approximation to the solution of Equation 1.10, since the desired counterfactual may not lie in that space.

In such a case, which is the relevant case in practice, MW advocate choosing a linear combination of the corresponding impulse responses that solves

$$\min_{\tilde{\boldsymbol{\nu}}^{\text{LSQ}}} \; ||\tilde{\boldsymbol{\mathcal{A}}}\boldsymbol{\Theta}_{A,\epsilon} \times \boldsymbol{\epsilon} + \tilde{\boldsymbol{\mathcal{A}}}\boldsymbol{\Theta}_{\nu^{\text{LSQ}},\mathcal{A}} \times \tilde{\boldsymbol{\nu}}^{\text{LSQ}}||, \tag{1.11}$$

and thereby enforces the desired counterfactual rule of Equation (1.10) as well as possible (in a least-squares sense). This solution can be computed directly from Equation (1.10) by (i) collecting all IRFs to the $n_{\nu,\mathrm{LSQ}}$ identified policy (news) shocks in $\boldsymbol{\Theta}_{\nu^{\mathrm{LSQ}},\mathcal{A}}$, (ii) assuming that this matrix has a specific structure $\boldsymbol{\Theta}^{\mathrm{LSQ}}_{\nu,\mathcal{A}} = [\boldsymbol{\Theta}_{\nu^{\mathrm{LSQ}},\mathcal{A}}, \mathbf{0}_{(n\times n_h)\times(n_h-n_{\nu,\mathrm{LSQ}})}]$ and (iii) solving for $\tilde{\boldsymbol{\nu}}^{\mathrm{LSQ}} = -\left(\tilde{\boldsymbol{\mathcal{A}}}\boldsymbol{\Theta}^{\mathrm{LSQ}}_{\nu,\mathcal{A}}\right)^{+}\tilde{\boldsymbol{\mathcal{A}}}\boldsymbol{\Theta}_{A,\epsilon}\times\boldsymbol{\epsilon}$ with $\left(\tilde{\boldsymbol{\mathcal{A}}}\boldsymbol{\Theta}^{\mathrm{LSQ}}_{\nu,\mathcal{A}}\right)^{+}$ as the Moore-Penrose inverse of $\tilde{\boldsymbol{\mathcal{A}}}\boldsymbol{\Theta}^{\mathrm{LSQ}}_{\nu,\mathcal{A}}$.

To understand when this approximation corresponds to the true solution and in order to compare the three approximation techniques presented in this paper, the assumption regarding the structure of the matrix of impulse responses to policy (news) shocks $\boldsymbol{\Theta}_{\nu^{\mathrm{LSQ}},\mathcal{A}}$ is key. Intuitively, this proposed approximation of the true solution in Equation (1.10) corresponds or comes close to the true solution, if in the unknown underlying data generating structural model (i) the effects of the policy news shocks corresponding to the announcement about the future level of the policy instrument at horizons that are not identified are small, implying that approximating them by zero is a good approximation, or (ii) if the structure of the model is such that the (expected) path of the policy instrument that would prevail under the counterfactual policy rule can be closely replicated by the policymaker only the subset of shocks in v^{LSQ}, implying that even if zero is not a good approximation, those entries would not matter for the solution.

While the approximation is fully robust to the Lucas critique because it only leverages policy news shocks announced at period 0 and does not make any assumptions on the expectations of agents, it only minimizes the squared deviations from the counterfactual policy rule described by $\tilde{\boldsymbol{\mathcal{A}}}$ and thus does not perfectly enforce it. Although the error of approximation vanishes as $n_{\nu,\mathrm{LSQ}}$ approaches n_h, it may be large if $n_{\nu,\mathrm{LSQ}} < n_h$.

Sims-Zha-approximation: The second approximation of the solution to Equation (1.10) is based on the procedure proposed by Sims and Zha (2006, henceforth SZ) which is frequently used to compute policy counterfactuals in the SVAR literature (see for instance Bachmann and Sims (2012), Brunnermeier et al. (2021), Breitenlechner et al. (2022)). Traditionally this approximation is sketched as a hypothetical scenario where a policymaker only has access to a single (contemporaneous) policy shock (i.e. ν_0) and then, successively in each period, surprises agents by choosing the size of the policy shock such that the counterfactual policy rule holds perfectly after all shocks play out. In contrast to this hypothetical scenario, the stacked sequence space representation in Equation 1.10 features sequences of policy news shocks $\boldsymbol{v}$ about changes in the policy instrument announced at horizon 0 instead of (unexpected) contemporaneous policy shocks implemented each period. But parsing the procedure of SZ into the sequence space representation uncovers the underlying assumption on the data-generating, structural model that would need to prevail for the SZ procedure to yield a good approximation to the true solution of Equation (1.10). In particular, the assumption underlying the SZ procedure is that, for each variable $j \in [\boldsymbol{x}, \boldsymbol{z}]$, the corresponding matrix of impulse responses to policy news shocks $\boldsymbol{\Theta}^{\mathrm{SZ}}_{j,\nu,\mathcal{A}}$ is a lower triangular matrix. This would be the case if agents in the underlying structural model are fully myopic concerning announced changes in the future level of the policy instrument. Intuitively this assumption implies that there is no difference between a change in the level of the policy instrument announced at date 0 and implemented at date t and an unexpected shock to the instrument at date t.

Consequently, under this assumption, the full matrix of impulse responses to policy (news) shocks can $\mathbf{\Theta}^{\mathrm{SZ}}_{\nu,\mathcal{A}}$ can be constructed and thus the counterfactual policy rule can be perfectly enforced, using solely the knowledge of the impulse responses to a single contemporaneous policy shock $\nu_{0,t}$. In particular, each column c of matrix $\mathbf{\Theta}_{j,\nu,\mathcal{A}}$, which describes the n_h responses of variables j to an announced change in the policy instrument c periods ahead, is given by $\mathbf{\Theta}^{\mathrm{SZ}}_{j,\nu_c,\mathcal{A}} = [\mathbf{0}'_{1\times(c-1)}, \mathbf{\Theta}'_{j,\nu_{0,t},\mathcal{A},0:(n_h-c)}]'$ where the vector $\mathbf{\Theta}_{j,\nu_{0,t},\mathcal{A},0:(n_h-c)}$ describes the impulse responses of variable j to the contemporaneous policy shock $\nu_{0,t}$ from period 0 to period $n_h - c$. Stacking across all n_h policy news shocks ν yields the lower triangular matrix of impulse responses to all policy news shocks $\boldsymbol{v}$ for variable j $\mathbf{\Theta}^{\mathrm{SZ}}_{j,\nu,\mathcal{A}} = [\mathbf{\Theta}^{\mathrm{SZ}'}_{j,\nu_0,\mathcal{A}}, \mathbf{\Theta}^{\mathrm{SZ}'}_{j,\nu_1,\mathcal{A}}, ..., \mathbf{\Theta}^{\mathrm{SZ}'}_{j,\nu_{n_h},\mathcal{A}}]'$. Stacking across all variables $j_1...j_n$ then gives the full matrix of impulse responses of all variables to all policy (news) shocks $\mathbf{\Theta}^{\mathrm{SZ}}_{\nu,\mathcal{A}} = [\mathbf{\Theta}^{\mathrm{SZ}'}_{j_1,\nu,\mathcal{A}}, \mathbf{\Theta}^{\mathrm{SZ}'}_{j_2,\nu,\mathcal{A}}...\mathbf{\Theta}^{\mathrm{SZ}'}_{j_n\nu,\mathcal{A},n}]'$, which almost surely results in the matrix product $\tilde{\boldsymbol{\mathcal{A}}}\mathbf{\Theta}^{\mathrm{SZ}}_{\nu,\mathcal{A}}$ being an invertible matrix. As such, the solution to the problem of choosing the sequence of policy (news) shocks that make the counterfactual policy rule hold in Equation (1.10) is then given by $\tilde{\boldsymbol{\nu}}_{\mathrm{SZ}} = -\left(\tilde{\boldsymbol{\mathcal{A}}}\mathbf{\Theta}^{\mathrm{SZ}}_{\nu,\mathcal{A}}\right)^{-1}\tilde{\boldsymbol{\mathcal{A}}}\mathbf{\Theta}_{A,\epsilon} \times \epsilon$.

Although this approximation yields an invertible matrix implying that it perfectly enforces the counterfactual policy rule and does not seek to minimize the squared deviations, the approximation only corresponds to the true solution under certain assumptions. In particular, to understand when this approximation yields an accurate solution the assumptions on the impulse responses to policy news shocks $\mathbf{\Theta}^{\mathrm{SZ}}_{\nu,\mathcal{A}}$ are key. The assumption of lower triangular submatrices for each $\mathbf{\Theta}^{\mathrm{SZ}}_{j_1,\nu,\mathcal{A}}$ implies that the effects of the policy news shocks only materialize once the change in the policy instrument is actually enacted, which justifies the zero entries and the specific structure of each of the impulse response matrices. If agents are forward-looking and the policy communication is effective (i.e. they know about the announced change in the policy instrument period before it materializes), this approximation deteriorates. Therefore, the accuracy of the approximation depends on the way agents form their expectations about the future path of the policy instrument and thereby the degree of myopia of agents in the non-policy block of the (unknown) structural model underlying the data.

"Hybrid"-approximation: The third approximation to the solution of Equation (1.10) combines the two previously described approaches. While the first approximation tries to best enforce the counterfactual policy rule (in a least squares sense) at each point in time t and in expectations n_h for all horizons, the second approximation neglects the expectations components. It rather perfectly enforces the counterfactual policy rule ex-post (i.e. once all shocks played out), implying that, although the counterfactual policy rule holds when observing the counterfactual impulse responses, these may not correspond to the desired counterfactual from the true underlying structural model. Since both approaches are on the opposing spectrum of the degree of forward-lookingness of agents in the non-policy block of the true, unknown underlying model, a natural "hybrid" approach would be to (i) enforce the counterfactual policy rule perfectly ex-post and (ii) in n_e of the n_h ex-ante expectations with $n_e \leq n_h$.

MW in their Online Appendix briefly sketch a sequential procedure on how to obtain an approximation that perfectly enforces the policy rule ex-post as well as in n_e period ahead expectations using impulse responses to the $\boldsymbol{\nu}^{\text{hybrid}} \in \boldsymbol{\nu}$ policy shocks with $n_{\nu^{\text{hybrid}}} = n_e{+}1$. In their sequential procedure, in each period the policymaker announces that the counterfactual policy rule will hold today and can be expected to hold for the next n_e periods.[6] In the next period, agents are then surprises that she extends her promise for one more period, causing agents to (surprisingly) revise their expectations about this newly revealed path of the policy instrument for the next period and thereby leading them to reconsider their decisions for the remaining periods. This revision would imply that, despite her previous commitment, the path of the policy instrument would differ from the announced path one period ago. Therefore she issues a set of precisely specified policy news shocks that perfectly enforce the previously announced path of the policy rate and imply that the policy rule again holds at time t in the next n_e period ahead expectations. In Appendix A.4 I provide a detailed description of how this sequential approach can be parsed into the sequence space representation of Equation 1.10 and discuss the assumptions on the underlying structural model that are revealed by doing so. In a nutshell, these assumptions are threefold. The first one is that the policymaker can directly steer the next n_e time t private sector expectations of the policy instrument using explicit policy rules alongside a sequence of corresponding and distinct policy shocks. The second assumption is that she lacks a commitment device to further guide private sector expectations for all remaining periods. Thirdly, this is paired with an assumption on the information set of agents that are assumed to only observe policy rules and corresponding policy shocks that are explicitly spelled out at time t, implying that only their t to $t + n_e$ are formed under the counterfactual policy rule. As in the previous cases, the assumptions behind this approach imply that the matrix $\boldsymbol{\Theta}_{\nu,\mathcal{A}}^{\text{hybrid}}$ (i) has a specific structure (ii) can be fully recovered using only the impulse responses to the $n_{\nu^{\text{hybrid}}} = n_e + 1$ policy shocks and (iii) is invertible. This implies that the hybrid solution to the problem in Equation 1.10 is given by $\tilde{\boldsymbol{\nu}}_{\text{hybrid}} = -\left(\tilde{\boldsymbol{\mathcal{A}}}\boldsymbol{\Theta}_{\nu,\mathcal{A}}^{\text{hybrid}}\right)^{-1} \tilde{\boldsymbol{\mathcal{A}}}\boldsymbol{\Theta}_{A,\epsilon} \times \epsilon$. Intuitively, given the assumed structure of $\boldsymbol{\Theta}_{\nu,\mathcal{A}}^{\text{hybrid}}$ implied by the "hybrid approximation" it becomes apparent that the hybrid solution will correspond to the true solution in Equation (1.10) if expectations about the policy instrument more than n_e periods ahead expectations are irrelevant for the decision problem of agents in the underlying data generating structural model. This

[6] This is reminiscent of an optimal policy problem, where the policy maker lacks a commitment device to convince agents of changes in the policy stance more n_e periods ahead.

counterfactual approach could, for instance, be a good approximation if, in the underlying true model, agents follow some form of exponential discounting, so that the importance of the expectations decays quickly with the horizon.[7] As in the least-squares approximation, the approximation error of the hybrid approximation vanishes as n_e approaches n_h implying that the number of identified shocks $n_{\nu^{\text{hybrid}}}$ approaches n_ν.

1.3.2 Counterfactual: What if OPEC stabilizes the oil price?

I employ the described approach to constructing empirical policy rule counterfactuals to gauge the role played by energy prices for the monetary transmission in the euro area. In particular, I assume that OPEC — as stated in its statutes— aims to stabilize the global oil price.[8] This implies the counterfactual OPEC policy rule is $E_t[\hat{p}^{oil}_{t+s}] = 0 \; \forall t, s$, which can readily be embedded into the system in Equation (1.10). In fact, $\tilde{\boldsymbol{\mathcal{A}}}$ simply becomes a selection matrix that selects the entries in $\boldsymbol{\Theta}_{\epsilon,A}$ and $\boldsymbol{\Theta}_{\nu,A}$ corresponding to the oil price for all n_h horizons.

To implement the three different approaches to the solution of Equation (1.10) I identify OPEC-related oil supply news shocks using the proxy variables constructed by Känzig (2021). In particular, Känzig (2021) proposes that the high-frequency changes of oil price futures around OPEC meetings provide a valid instrument for OPEC's oil supply news shocks. As explained in the previous sections, the accuracy of the approximations depends on the number of identified policy news shocks. Therefore I depart from Känzig (2021), who only uses the first principal component of changes in the oil price futures at many horizons to identify a single oil supply news shock. Instead I use high frequency changes in the 1 month ($m^{oil}_{t,1m}$) and 12 month ($m^{oil}_{t,12m}$) futures to identify a short-term ($\nu^{oil}_{t,short}$) and a medium-term ($\nu^{oil}_{t,medium}$) oil supply news shock. For the baseline specification, I directly follow the procedure of McKay and Wolf (2023) and condition on the point-estimate impulse responses for separately estimated initial shock $\boldsymbol{\epsilon}$. This is in keeping with standard practice in the policy counterfactual literature, which tends to take as given initial point estimates (see e.g Rotemberg and Woodford (1997), Eberly et al. (2020), Wolf (2023)). For this exercise, this corresponds to conditioning on the impulse responses to a monetary policy shock in Figure 1.1. For reasons of consistency and robustness, when identifying the two oil supply news shocks I nevertheless also identify— and thereby account for the presence of— a euro area monetary policy shock. I report the resulting impulse responses in Figure A.5 of the Appendix which shows that results for the effects of a euro area monetary policy shock are robust to identifying the monetary policy shock jointly alongside the oil supply

[7] Even for a full information rational expectations HANK model where the counterfactual policy rule is only enforced at point time t and in $n_e = 1$ period ahead expectations, this approximation already comes very close to the true counterfactual obtained from the underlying model (see McKay and Wolf (2022b)).

[8] As nicely summarized in (Känzig 2021, p.6), "According to the statutes, OPEC's mission is to stabilize global oil markets to secure an efficient, economic and regular supply of petroleum to consumers, a steady income to producers and a fair return on capital for those investing in the petroleum industry".

news shocks.[9]

In the notation of Equation (1.2a) and (1.2b) identifying the oil supply news shocks alongside the euro area monetary policy shock requieres extending the vector of proxies $\boldsymbol{m_t} = [m^{oil'}_{t,1m}, m^{oil'}_{t,12m}, m^{ir'}_{t}]$ and the vector of structural shocks of interest $\boldsymbol{\epsilon^\star_t} = [\nu^{oil'}_{t,short}, \nu^{oil'}_{t,medium}, \epsilon^{MP'}_t]$. To disentangle the three structural shocks identified using the three instruments I postulate

$$\mathbf{E}[\boldsymbol{m_t}, \boldsymbol{\epsilon_t}^{o'}] = \mathbf{0}, \quad \mathbf{E}[\boldsymbol{m_t}, \epsilon^{\star'}_t] = \begin{bmatrix} \upsilon_{1,1} & \upsilon_{1,2} & \upsilon_{1,3} \\ \upsilon_{2,1} & \upsilon_{2,2} & \upsilon_{2,3} \\ 0 & 0 & \upsilon_{3,3} \end{bmatrix}, \quad \upsilon_{1,1} > \upsilon_{1,2}, \quad \upsilon_{2,2} > \upsilon_{2,1} \tag{1.12}$$

Thus, I assume that (i) all three instruments are uncorrelated with the remaining structural shocks $\boldsymbol{\epsilon_t}^o$, (ii) the high-frequency changes in the interest rate m^{ir}_t are unaffected by OPECs oil supply news shocks and (iii) the medium-term (short-term) oil supply news shock has a larger effect on the high-frequency change in 12 months (1 month) oil price futures.[10] I furthermore require that a contractionary medium-term (short-term) oil supply news shock raises the price of oil at horizon 12 (1). The impulse responses to the two oil supply news shocks are depicted in Figures A.3 and A.4 of the Appendix. In a nutshell, both shocks raise the price of oil, the energy component of the HICP and lead to a contraction in output in the euro area. For the short-term oil supply news shocks the response of the oil price is strong and immediate, while the medium-term oil supply shock rather results in an increase in the price of oil at longer horizons.

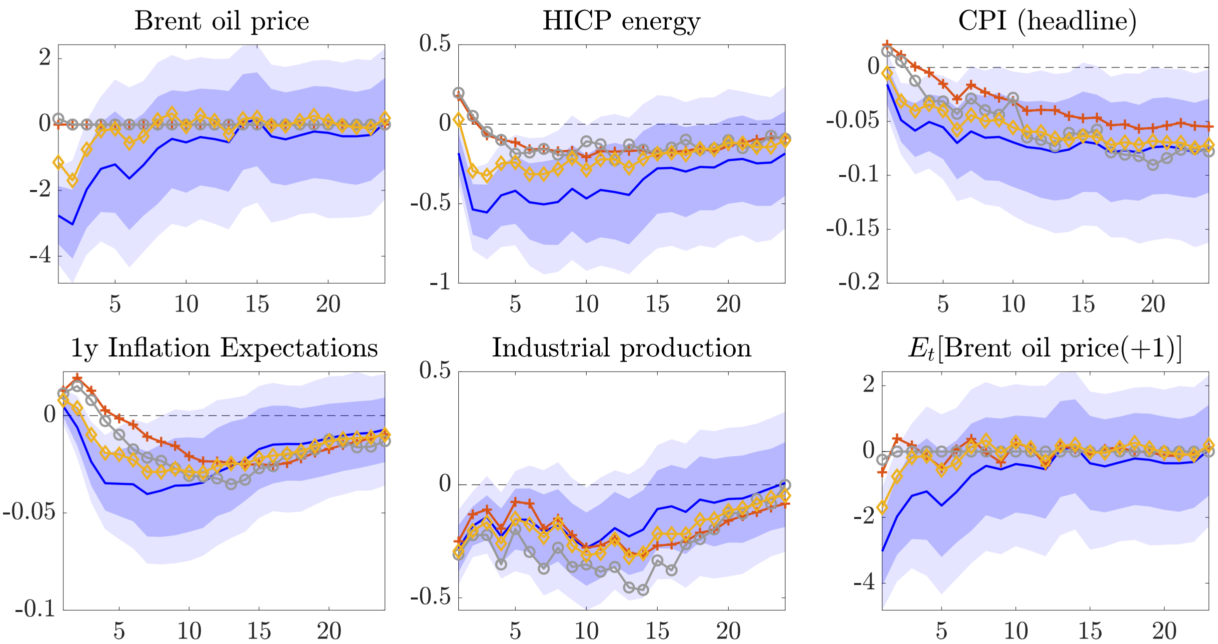
I then use the estimated impulse responses to these shocks to compute the three approximations to the solution of Equation (1.10) which characterizes the true counterfactual, in which OPEC stabilizes the oil price. The results from this exercise are shown in Figure 1.2. In particular, the golden line in Figure 1.2 corresponds to the least-squares approximation, which MW advocate as the “best Lucas-critique-robust approximation” to the solution of Equation (1.10). Under the counterfactual policy rule, where OPEC aims to perfectly stabilize the oil price, the reduced responsiveness of the oil price to an ECB monetary policy shock directly translates into a notably smaller response of local energy prices and thereby inflation and inflation expectations, especially at shorter horizons. This is true even though the transmission to economic activity is hardly altered, which further corroborates the intuition that the reduction in the monetary-policy-induced fall of CPI in the counterfactual is not due to changes in prices of domestically produced goods.

It is important to note that in this application the number of identified OPEC oil supply

[9] While the impulse responses look similar and asymptotically should perfectly coincide with the ones in Figure 1.1, they are not identical due to the finite sample nature of the estimation. In that sense, one can think of this two-step approach as an asymptotic approximation.

[10] A natural additional assumption would be that the high-frequency changes in the oil price futures around OPEC announcements are unaffected by monetary policy shocks (i.e. $\upsilon_{1,3}$, $\upsilon_{2,3} = 0$). This would imply an overidentified system that can not be handled by the algorithm of Arias et al. (2021). I therefore propose this less restrictive set of assumptions. It turns out that the estimated coefficients for the two parameters are close to zero even without these assumptions.

Figure 1.2: What if EA monetary policy shocks do not affect oil prices?



Notes: Impulse response functions to a one standard deviation monetary policy shock showing the point-wise posterior means along with 68% and 90% point-wise credible sets in blue. Horizon in months. The bronze line shows the Point-wise posterior means from the Sims-Zha approximation to the policy rule counterfactual, where OPEC stabilizes the oil prices. The golden line corresponds to the counterfactual obtained using the least-squares approximation. The silver line corresponds to the hybrid approximation. For all approaches, I directly follow the procedure of McKay and Wolf (2023) which necessitates conditioning on the estimated baseline impulse response to the monetary policy shock depicted by the blue lines in Figure 1.1 when computing the counterfactuals. Furthermore, I only retain draws from the posterior of impulse responses for which the resulting counterfactual response does not show explosive dynamics.

news shocks is smaller than the envisioned horizon $n_h = 24$, therefore the oil price is not perfectly stable under the "least squares approximation" as depicted by the golden line in the top left panel of Figure 4. In particular, it still falls after an EA monetary tightening, although much less than in the baseline. Therefore this approximation may underestimate the importance of global energy prices as proxied for by the response of the oil prices in the monetary transmission. A way out could be to resort to the Sims-Zha approximation depicted in bronze. Recall that under this approximation the policy rule is enforced perfectly at the cost of assuming that agents in the underlying model are myopic, which implies that the expectations component about the future path of the oil price that enters agents' decisions in the non-policy block is neglected.[11] Intuitively, this assumption implies that what matters for the behavior of agents in the non-OPEC block of the system is the spot price of oil (i.e. approximately the price at the gas station) and not their expectations about the future path of the oil price. As shown in Figure 1.2 the counterfactual fall in headline HICP (i.e. the ECB's primary objective) is even further reduced when the solution

[11]In line with the discussion in the previous section I treat the short-term oil supply news shock as the contemporaneous policy shock $\nu_{0,t}$ necessary for the construction of the Sims-Zha approximation.

to the counterfactual is approximated using the Sims-Zha approximation. In other words, the bronze line lies above the golden line, which indicates that the importance of the global component of the energy price channel for the domestic transmission of EA monetary policy is estimated to be larger under the Sims-Zha approximation than under its least squares counterpart.

This could be the case for two reasons. The first potential explanation is that, as shown in the top left panel of Figure 1.2, the oil price is perfectly stabilized under the Sims-Zha approximation. This implies that local energy prices and thereby ultimately local inflation fall by less compared to the least-squares approximation. The second possible explanation is related to oil price expectations, which are embedded in the least-squares approximation but neglected in the Sims-Zha approximation. In particular, because the importance of oil prices in the monetary transmission is estimated to be lower in the least-squares approximation, one could also deduce that expectations about OPEC stabilizing the oil price not only at time t but also in the future play an important role for the total effect. The hybrid approximation depicted in silver allows for a partial answer to this question. In particular, in the hybrid approximations agents are assumed to know that OPEC will stabilize the oil price not only in period t, as it is the case in the Sims-Zha approximation, but also in $n_e = 1$ period ahead expectations. Therefore deviations between the bronze and silver line in Figure 1.2 can be attributed to the role of one period ahead oil price expectations. By inspecting the panels in Figure 1.2 it becomes apparent that at least the one period ahead expectation about the oil price is not able to explain a large fraction of the differences between the least squares solution in gold, which incorporates the role of expectations, and the SZ approximation in bronze, which neglects the expectations component. One way to rationalize why the estimated counterfactual impulse responses using the hybrid approach, which explicitly incorporates the fact that agents know that the counterfactual policy rule also holds in one period ahead expectations, closely track the ones from SZ approximation for many variables, is the notion that oil prices are a perceived as a random walk. Thus, once the oil price is observed to be at a certain level, the expectation of the oil price remains at that level. Therefore, once the oil price is stabilized, it does not seem to matter a lot for the decisions of agents that policy communicates that it will continue to stabilize in the next period, as agents in the data seem to expect the oil price to remain at that level anyway.

In the bottom right panel of Figure 1.2 I illustrate this insight by plotting for each period t the model-consistent one period ahead oil price expectations, which can be recovered by combining the procedure of MW with the impulse responses estimated from the BPSVAR. By inspecting the bronze line which depicts the solution under the SZ approximation it becomes clear that the estimation suggests that, once the oil price is stabilized ex-post, agents expect it to remain close to that level for the remaining periods. Therefore the announcement that OPEC will also continue to stabilize the oil price in the next period, which is embedded in the hybrid approach in silver and neglected by the SZ approach in bronze, does not significantly alter agents' decision on their time t allocation once they observe the oil price to be stabilized at time t. This rationalizes why the Sims-Zha approximation and the hybrid approximation closely coincide.

For this reason, I tentatively lean towards the first explanation for the differences between the Sims-Zha approximation and the least-squares approximation, which implies that a larger share of the difference between the two estimates can be attributed to the fact that the least-squares approximation does not fully enforce OPECs counterfactual policy rule instead of attributing it to the difference regarding the expectations of agents. Regardless of the approximation technique and therefore regardless of the assumptions on the underlying model structure, I conclude that (i) all approximations to the policy rule counterfactual agree that global energy prices as measured by global oil prices, matter for the transmission of euro area monetary policy to inflation and inflation expectations (ii) the least squares approximation (Sims-Zha approximation) seems to provide a lower (upper) bound of the importance of this channel.

Lastly, it is important to note that, irrespective of the approximation technique, this analysis should be seen as a lower bound for the overall importance of energy prices in the transmission of euro area monetary policy. Recall that the HICP energy index of the euro area consists of a weighted average of all local energy prices (denominated in euros). Therefore movements in the dollar price of oil, despite oil being the dominant component of this basket, only correspond to a fraction of total price movements in the index of euro area consumer energy prices. As such the fully empirical counterfactual analysis does not exactly correspond to a case where the ECB's decisions do not impact all components of the euro area's energy basket, which would allow for fully quantifying the role of energy prices in the monetary transmission. I conjecture that if the ECB's decisions would not move global *and* local energy prices, the short- and medium-run effects of a change in the ECB's policy stance on euro area inflation would be further mitigated. But this counterfactual would arguably be hard to envision as being caused by a plausible change in some policy rule. As I aim to leverage the generality of the policy rule approach by MW to circumvent having to commit to a specific model when computing the counterfactual impulse responses, I solely focus on the role of global energy prices, thereby stacking the deck against finding an important role of energy prices in the monetary transmission.

1.4 Optimal monetary policy in light of the energy price channel

Previous sections of the paper explored the role played by the endogenous response of energy prices in the transmission of exogenous monetary policy shocks. In this section, I flip the question and ask: What role does endogenous monetary policy, as characterized by the policy rule, play in the transmission of exogenous energy price shocks? In particular, I ask the question of how the ECB —given its mandate— should optimally react to a surge in energy prices, which adds a normative dimension to the largely positive analysis in the previous sections.

Although the question regarding the role played by the endogenous monetary policy reaction for the transmission of energy price shocks was raised already in 1997 by Bernanke et al. (1997), it could not be rigorously addressed in a fully empirical model and instead has been answered in different theoretical approaches (see e.g. Bodenstein et al. (2012)).[12] The recent developments of sufficient statistics approaches to the evaluation of optimal policy (Barnichon and Mesters (2023), McKay and Wolf (2023)) enable addressing this question in an empirical framework without having to rely on a theoretical model. To do so, I leverage the approach of McKay and Wolf (2023) discussed in Section 1.3.1 to empirically analyze the optimal conduct of ECB monetary policy when being confronted with an exogenous oil price shock, which acts as a stand-in for an exogenous increase in energy prices. Afterward, I interpret the results in light of previous findings on the role of energy prices in the monetary transmission mechanism.

1.4.1 Computing optimal policy counterfactuals

The approach of MW to estimate Lucas-critique robust policy-rule counterfactuals in SVAR models discussed in Section 1.3.1 readily extends to computing optimal policy counterfactuals. In line with Barnichon and Mesters (2023) MW define the optimal policy response as the response that implements an allocation that allows the policymaker to achieve its mandate optimally. Note that this definition of optimality differs from the standard textbook definition under which the policymaker aims to maximize a measure of welfare whose exact nature is inherently tied to a specific model and calibration (see (Galí 2015, Ch.5)). Throughout this section, I define optimality of policy along the lines of McKay and Wolf (2022a) and Barnichon and Mesters (2023) because (i) I aim to connect to the recent surge in sufficient statistics, approaches to macroeconomics (McKay and Wolf (2023), Barnichon and Mesters (2023), Barnichon and Mesters (2024)) and (ii) because, as argued by (McKay and Wolf 2022a, p.3) “it is arguably the relevant objective function for real-world central banks”. Furthermore, this definition comes with the advantage that the loss function does not depend on specific features of a particular theoretical model.

In particular, suppose the mandate of the central bank can be described by a loss function and the central bank is expected to minimize a quadratic loss function of the form

$$\mathcal{L} = \frac{1}{2}\sum_{i=1}^{n_x} \lambda_i \boldsymbol{x}_i' \boldsymbol{W} \boldsymbol{x}_i' = \frac{1}{2}\boldsymbol{x}'(\Lambda \otimes W)\boldsymbol{x} \tag{1.13}$$

where the $\boldsymbol{x}_i$ contains the time path of the endogenous variable i, λ_i describes the policy

[12] For instance Bernanke et al. (1997) humbly argue that their VAR-based method and its application to this question only constitutes “some modest [...] first steps toward sorting out the effects of systematic monetary policy on the economy [...]” ((Bernanke et al. 1997, p.92)).

weights attached to that variable with $\Lambda = diag(\lambda_1, \lambda_2 .... \lambda_{n_x})$. Finally, the matrix $\boldsymbol{W}$ summarizes the effects of time discounting in the policymaker's preferences potentially parameterized using a single discount factor β. The textbook solution to the optimal policy problem would amount to minimizing Equation (1.13) subject to the constraints embedded in Equations (1.4) in order to choose the optimal path of the policy instrument and thereby the optimal allocation (see (Galí 2015, Ch.5)).

MW show that the optimal policy problem can analogously be stated in impulse-response space. This implies minimizing the loss function subject to Equation (1.6) instead of Equation (1.4). While MW provide formal proof for this result and establish the equivalence of the resulting optimal policy rule and more classical results from the optimal policy literature as in Svensson (1997), and Giannoni and Woodford (2002), I provide a more heuristic explanation below. In particular, the approach utilizes the observation that the implementable space of allocations for the endogenous variables $\boldsymbol{x}$ and for the policy instrument $\boldsymbol{z}$ is fully characterized by the impulse responses $\Theta_{\nu,\mathcal{A}}$ to the sequence of policy (news) shocks $\boldsymbol{\nu}$:

$$\begin{bmatrix} \boldsymbol{x} \\ \boldsymbol{z} \end{bmatrix} = \Theta_{\nu,\mathcal{A}} \times \boldsymbol{\nu}. \tag{1.14}$$

Focusing on a single variable $\boldsymbol{x}_i$, Equation (1.14) implies that the space of possible allocations that the policymaker can achieve for this variable is given by

$$\boldsymbol{x}_i = \sum_{j=1}^{n_\nu} \Theta_{x_i,\nu_j,\mathcal{A}} \times \nu_j. \tag{1.15}$$

Plugging this expression for each $\boldsymbol{x}_i$ into Equation (1.13) and taking the first-order conditions with respect to each ν_j, one arrives at the condition

$$\sum_{i=1}^{n_x} \lambda_i \Theta'_{x_1,\nu,\mathcal{A}} W \times \boldsymbol{x_i} = \boldsymbol{0}. \tag{1.16}$$

For each $\boldsymbol{x}_i$ the term following the summation sign describes how a change in the policy instruments $\boldsymbol{\nu}$ would translate into a change in the endogenous variable $\boldsymbol{x}_i$ and weights these changes over time using the time discount matrix W. All the implied changes are then summed over all n_x variables x_i using the corresponding policy weight λ_i, which translates them into changes in the loss function of Equation (1.13). This rule then implies that the (weighted) sum of changes in the objective function resulting from a change in the policy instruments $\boldsymbol{\nu}$ has to equal zero. In other words, the gradient of the objective function with respect to the policy instruments has to be set to zero at the optimum.

This condition can be encapsulated into the matrices $\boldsymbol{\mathcal{A}_x}, \boldsymbol{\mathcal{A}_z}$ of the sequence-space representation to the of the model in Equation (1.5) by noting that the optimality condition in Equation (1.16) can be written in matrix form as

$$\sum_{i=1}^{n_x} \lambda_i \Theta'_{x_i,\nu,\mathcal{A}} W \boldsymbol{x_i} = \mathcal{A}^*_{\boldsymbol{x}} \boldsymbol{x} = \mathbf{0}. \tag{1.17}$$

Thus to embed the optimality condition into the policy block of Equation (1.16) it suffices to set

$$\begin{aligned} \mathcal{A}^\star_{\boldsymbol{x}} &= (\lambda_1 \Theta'_{x_1,\nu,\mathcal{A}} W, \lambda_2 \Theta'_{x_2,\nu,\mathcal{A}} W, \ldots, \lambda_{n_x} \Theta'_{x_{n_x},\nu,\mathcal{A}} W), \\ \mathcal{A}^*_{\boldsymbol{z}} &= \mathbf{0}, \end{aligned} \tag{1.18}$$

with $\Theta_{x_i,\nu,\mathcal{A}}$ as the matrix of impulse responses of variable i to all shocks in $\boldsymbol{\nu}$ under the baseline policy rule characterized by $\boldsymbol{\mathcal{A}}$.

By definition the policy rule described by $\mathcal{A}^\star_{\boldsymbol{x}}$ and $\mathcal{A}^\star_{\boldsymbol{z}}$ in Equation (1.18) denotes the optimal policy rule as it embeds the optimality condition directly into the model. In contrast to previous approaches in the literature (see Svensson (1997)), the implied optimal policy rule is fully characterized by impulse responses to policy (news) shocks, which can theoretically all be estimated from the data. Therefore the optimal policy counterfactual is a special case of approach to empirically estimated counterfactual impulse responses of Section 1.3.1, which implies setting the counterfactual policy rule as described in Equation (1.18). Again impulse response functions prove to be sufficient statistics, even for an optimal policy counterfactual.

1.4.2 Optimal euro area monetary policy response to an oil supply shock

The framework in Section 1.4.1 requires three key ingredients. First, one needs to take a stance on the ECB's relevant loss function. I follow McKay and Wolf (2022a) and Barnichon and Mesters (2023) and derive the loss function from the primary mandate of the central bank and define the optimal policy response as the response that implements an allocation that allows the policy maker to optimally achieve this mandate. The ECB's primary mandate is to maintain price stability, which it defines as an inflation target of approximately 2 percent over the medium term. Therefore I specify the loss function below such that the central bank aims to minimize the deviations of contemporaneous and future levels of Harmonised Index of Consumer Prices (HICP) inflation with no weight on output or unemployment. As I focus on counterfactual impulse responses, I assume that, in steady-state, euro area inflation will be at the ECB's target. Under this assumption, the impulse responses of HICP inflation then characterize those deviations that the central bank seeks to minimize. Importantly, I assume that when observing deviation from its target, the central bank attaches a higher weight to the inflation deviations which are closer to the medium-term horizon. Although the ECB's mandate does not precisely define what the medium-term horizon exactly means, there is good evidence based on the ECB's own projections that, in practice, the relevant horizon corresponds to approximately 6-8 quarters (Paloviita et al. (2021)).[13]

Given these considerations, the loss function takes the following form:

$$\mathcal{L} = \lambda_\pi \boldsymbol{\pi}' W \boldsymbol{\pi}, \tag{1.19}$$

with $\lambda_\pi = 1$ and the matrix W is defined as $W = (\text{diag}(\beta^{24}, \ldots \beta^2, \beta, 1))$ to account for the fact that I am interested in the first 24 months following the shock and that the ECB gives a higher weight to deviations from target that occur in the medium term.[14] The discount factor β is calibrated such that, in a standard New Keynesian model, the corresponding annualized real interest rate would be 2%. Furthermore, the term $\boldsymbol{\pi} = \mathcal{D}\boldsymbol{P}^{\text{HICP}}$ represents the transformed impulse responses of the (log) level of the Harmonized Index of Consumer Prices (HICP), denoted as $\boldsymbol{P}^{\text{HICP}}$. The operator $\mathcal{D}$ appropriately converts these impulse responses to changes in year-on-year inflation rates. This transformation aligns with the mandate of the European Central Bank (ECB), as defined by its governing council in 1998.

A second key component of this analysis is the choice and identification of the shock whose propagation is to be compared under the empirically identified baseline policy rule described by $\boldsymbol{\mathcal{A}}$ and the optimal policy rule described by $\boldsymbol{\mathcal{A}}^\star$. As in the previous section, I choose to identify a short-run oil supply shock as a proxy for a general energy price shock. Specifically, I identify a contemporaneous oil supply news shock using the high-frequency change in 1-month oil price futures around the relevant OPEC meetings identified by Känzig (2021). For the sake of brevity, I will refer to this shock interchangeably as "oil supply shock" or "oil price shock".

Third, one needs to identify several euro area monetary policy shocks in order to approximate the propagation of these oil price shocks under optimal policy rule. Using a strategy similar to that used for the OPEC-related oil supply news shocks in section 1.3.2, I identify the different dimensions of euro area monetary policy using high-frequency changes in interest rates at different maturities. In particular, I retain the high-frequency changes in the 3-month OIS yield as my preferred proxy for conventional monetary policy (CMP) shocks, while using changes in the 2-year rate as a proxy for forward guidance (FG)

[13](Paloviita et al. 2021, p.132) note that "[...]one can plausibly argue that the projected inflation rates at the end of the forecast horizon give the public a good guideline for inflation which the ECB considers consistent with its mandate. This is supported by the fact that inflation forecasts have typically converged to the proximity of "close but below two" already after about six quarters.".

[14]Note that, given the focus on a single objective π, the weighting matrix should not matter in theory, if the ECB would operate in the fully unconstrained space of implementable allocations of Equation 1.14. Intuitively, if the ECB in the application would have perfect knowledge of and perfect access to all 24 instruments (shocks) in $\boldsymbol{\nu}$ it could perfectly stabilize the 24 targets, that have a positive weight in the loss function. Given that I do not fully identify the entire menu of policy shocks $\boldsymbol{v}$, the set of possible allocations that can be implemented to the space of empirically identified policy shock paths is restricted, which implies that the weighting matrix matters because in this application the central bank lacks the tools to perfectly stabilize inflation at all horizons.

shocks, which is thought to contain news about the future level of the policy instrument.[15] Furthermore, following the literature that uses the one-year yield as an indicator for the stance of short-run monetary policy, I add the 5-year Bund yield as a measure of the longer-run monetary policy stance into the baseline SVAR model. In the notation of Equation (1.2a) and (1.2b) the aforementioned changes BPSVAR model imply that I extend the vector of proxies $\boldsymbol{m_t} = [m^{oil'}_{t,1m}, m^{ir'}_{t,3m}, m^{ir'}_{t,2y}]$ and the vector of structural shocks of interest $\boldsymbol{\epsilon^{\star}_t} = [\epsilon^{oil'}_{t,short}, \nu^{mp'}_{t,cmp}, \nu^{mp'}_{t,fg}]$. As in section 1.3.2 to disentangle the three structural shocks identified using the three instruments I postulate

$$\mathbf{E}[\boldsymbol{m_t}, \boldsymbol{\epsilon_t}^{o'}] = \mathbf{0}, \quad \mathbf{E}[\boldsymbol{m_t}, \epsilon^{\star'}_t] = \begin{bmatrix} \upsilon_{1,1} & \upsilon_{1,2} & \upsilon_{1,3} \\ 0 & \upsilon_{2,2} & \upsilon_{2,3} \\ 0 & \upsilon_{3,2} & \upsilon_{3,3} \end{bmatrix}, \quad \upsilon_{2,2} > \upsilon_{3,2}, \quad \upsilon_{3,3} > \upsilon_{2,3} \tag{1.20}$$

Firstly, I assume that all three proxies, i.e. the high-frequency changes in the short-run oil price as well as the two interest rate futures, are uncorrelated with the remaining structural shocks. Secondly, I assume that the high-frequency changes in interest rate futures around ECB meetings remain unaffected by the oil price shock, as this shock is rather related to changes in OPEC's strategy. Thirdly, I posit that the conventional monetary policy shock has a stronger influence on high-frequency changes in the 3-month interest rate than the forward guidance shock. Fourthly, I assume that for the high-frequency changes in the 2-year interest rate futures, the forward guidance shock exerts more pronounced effects than the conventional monetary policy shock.[16]

As such, in this application, only two monetary policy shocks are available to the policymaker, rather than the full menu of policy shocks. Consequently, the set of hypothetical, feasible allocations that the monetary authority can implement in this application is given by

$$\begin{bmatrix} \boldsymbol{x} \\ \boldsymbol{z} \end{bmatrix} = \boldsymbol{y} = \Theta_{\nu,\mathcal{A}} \times \boldsymbol{\nu}^{ident} \tag{1.21}$$

By following the rule described in Equation (1.18) the monetary policy authority selects the allocation of $\boldsymbol{y}$ that minimizes the loss function in Equation (1.13) - within that empirically identified space of Equation (1.21).

Before analyzing the impulse responses to an oil price shock and the results from the optimal policy exercise, it is worth briefly discussing the impulse responses to the two identified monetary policy shocks, which characterize the set of feasible allocations that

[15] At the onset of the euro area's inception in 1999, the press releases communicating the ECB Governing Council decisions were not regularly followed by a press conference. However, since November 2001, each policy decision is regularly accompanied by a press conference. This change in the ECB's operational framework is an important consideration when identifying forward guidance shocks using high-frequency changes in OIS rates as a proxy. Altavilla et al. (2019) find that press releases do not contain forward guidance surprises while press conferences do. Therefore, I estimate the model for this specification using data starting in 2002.

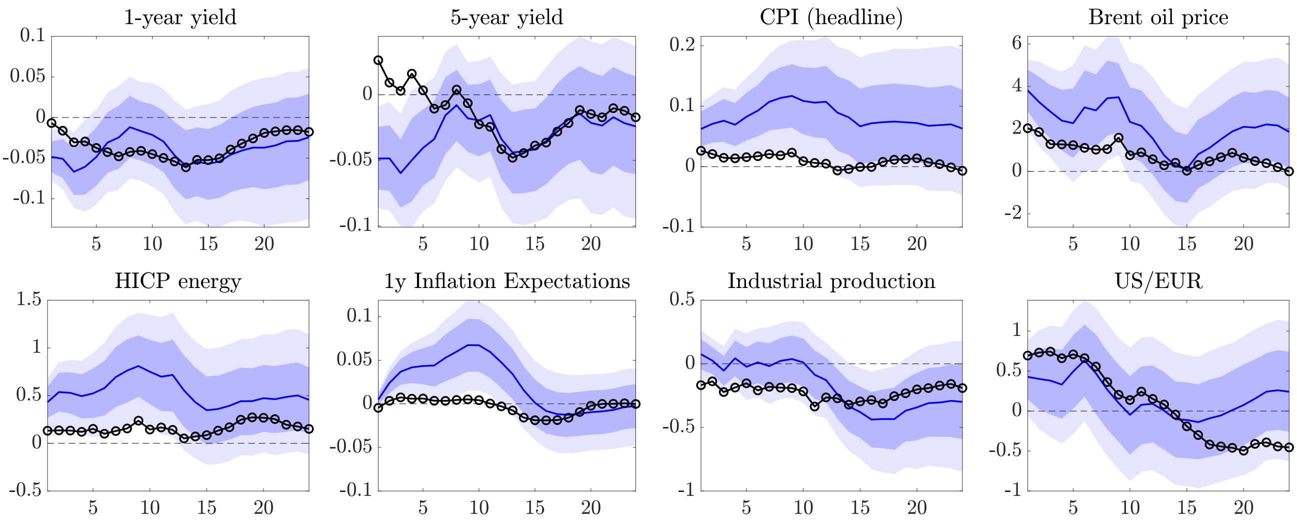
[16] For a detailed description of the implementation and estimation is provided in Appendix A.6.1.

the monetary authority can implement. The corresponding impulse response functions are shown in Figure 1.6 and in more detail in Figures A.6 and A.7 of the Appendix. In line with the results presented in Figure 1.1, a one standard deviation contractionary conventional monetary policy shock raises the 1-year yield and causes a fall in production, energy prices, and ultimately the ECB's primary objective, HICP inflation. Notably, this shock leaves the 5-year yield largely unchanged for the first 6 months. A one standard deviation forward guidance not only leads to a fall in output and inflation but also causes a steepening of the yield curve as it increases the 5-year bund yield by more than the 1-year yield.

The impulse responses to the identified oil supply shock under the empirical (baseline) policy rule are depicted by the blue lines in Figure 1.3 while the black-dotted lines describe the propagation of an oil supply shock under the optimal response of the ECB. The figure reveals stark differences between the empirical response and the one that is optimal according to the ECB‘s mandate. Under the baseline policy rule an oil supply shock leads to an immediate escalation in the Brent oil price, resulting in higher consumer energy prices, inflation, and inflation expectations within the euro area. Remarkably, this surge in energy costs and inflation rates is not a transient phenomenon; it persists into the medium term. Additionally, there is a delayed yet substantial economic contraction. Under the baseline policy rule, monetary policy does not appear to fight the inflation surge as interest rates even fall slightly. This aligns with the conventional wisdom that “in the past, central banks have typically looked through energy shocks” (Schnabel (2022)). Therefore, in response to oil supply shocks, the ECB seems to not only tolerate the implied increase in energy prices and inflation but even slightly lowers interest rates, possibly as a measure to mitigate the economic downturn. As the findings in the previous sections suggest that the ECB's decisions affect global and local energy prices, this expansionary response to the energy price shock potentially amplifies the endogenous rise in energy costs after the initial exogenous shock.

Was this optimal to do for a central bank that aims to stabilize (medium-term) inflation? And if not, what would it take to optimally fulfill the mandate within the identified space of feasible allocations? The estimated answer to this question is depicted by the black circled lines in Figure 1.3. As shown in the first row the estimates suggest that, to stabilize medium-term inflation, the policy stance should have been substantially tighter in the short run, when compared to the baseline. The prescribed additional tightening is front-loaded and slightly stronger at the longer end of the yield curve. It is important to note that the estimates do not suggest that the ECB should excessively increase interest rates after an oil supply shock to optimally fulfill its mandate. Instead, it seems like historically it has been too accommodative (relative to the optimal allocation) following the oil supply shock. The estimates imply that only slight immediate tightening at the longer end of the yield curve combined with a less expansionary “looking through” strategy at the short end of the yield curve, could go a long way in stabilizing inflation. The cost of this additional monetary policy tightening relative to the baseline strategy is arguably a mild front-loaded recession.

Figure 1.3: Impulse responses to an oil price shock under the empirical baseline and optimal monetary policy rule



Notes: Impulse response functions to a one standard deviation monetary policy shock showing the point-wise posterior means along with 68% and 90% point-wise credible sets in Blue. The black circled lines show the least-squares approximation of the responses of the endogenous variables under optimal policy with a loss function described in Equation 1.19. I directly follow the procedure of McKay and Wolf (2023) which implies that I condition on point-wise estimates of the separately estimated impulse responses to the oil supply shock.

But this induced initial contraction in output is comparatively small and offset by slightly higher output in the medium-term.

1.4.3 Optimal monetary policy in the absence of the global component of the energy price channel

I attribute the observation, that already a small additional tightening can help to quickly stabilize inflation and inflation expectations, to the energy price channel of monetary policy. In particular, I postulate that if the ECB would not affect (global) energy prices, the tightening would have to be substantially stronger to fulfill the mandate. Intuitively, the results in Section 1.3.1 indicate that global energy prices play a major role in the transmission of euro area monetary policy to domestic consumer prices especially in the short- and medium-term. The intuition again is that, in line with conventional wisdom, energy prices are fast-moving if not almost flexible prices that react much more quickly and more strongly to changes in demand than other, domestically produced, goods in the HICP basket. Therefore the ECB does not need to excessively and persistently tighten (and possibly engineer a large recession) to fulfill its mandate in the face of energy price shocks because a large part of the adjustment is borne by relatively flexible global energy prices. If, in line with the ECB's public communication, this channel is absent, the additional tightening would have to be stronger to fight the energy-price-driven surge in inflation.

To further illustrate this point I conduct a thought experiment and ask the question: What would the optimal monetary policy response to an oil price shock be if, counterfactually, the ECB's decision would not affect the Brent oil price? In other words, the thought experiment asks if the optimal allocation $\boldsymbol{y}$ would be different if the empirically identified, implementable space of possible allocations would not be described by Equation (1.21) but rather by

$$\begin{bmatrix} \boldsymbol{x} \\ \boldsymbol{z} \end{bmatrix} = \boldsymbol{y} = \Theta_{\nu,\widetilde{\mathcal{A}}} \times \boldsymbol{\nu}^{ident} \tag{1.22}$$

where the subscript $\widetilde{\mathcal{A}}$ indicates that the space is now characterized by counterfactual impulse responses. The special feature of these counterfactual responses is that, as is the case for the analysis in Section 1.3.2 and in line with the assumptions at the ECB, the identified monetary policy shocks $\boldsymbol{\nu}^{ident}$ do not impact global energy prices. Intuitively, if the policymakers at the ECB truly believe that their decisions do not impact global energy prices, the allocation resulting from this application approximates their optimal strategy in the face of an oil price shock.

A detailed step-by-step summary of the approach to estimating the optimal monetary policy response to an oil price shock under the assumption that the ECB's decision do not affect global oil prices can be found in Appendix A.6.2. In a nutshell, to estimate the results of this thought experiment I first compute the counterfactual impulse responses by applying the methodology sketched in Section 1.3.1 to the identified impulse responses for the contemporaneous monetary policy and forward guidance shocks. By approximating the solution using the least-squares approximation, I stack the deck against finding a strong implication of the global component of the energy price channel. The reason is that, as shown Section 1.3.2, this approximation results in comparatively smaller implications of energy prices in the transmission of monetary policy. As a next step, given the same loss function of Equation (1.19) and the same impulse responses to the oil price shock ϵ, the monetary authority then chooses the optimal allocation within the counterfactual space of feasible allocation described in Equation (1.22). As shown by McKay and Wolf (2023) this amounts to setting the implied policy rule to

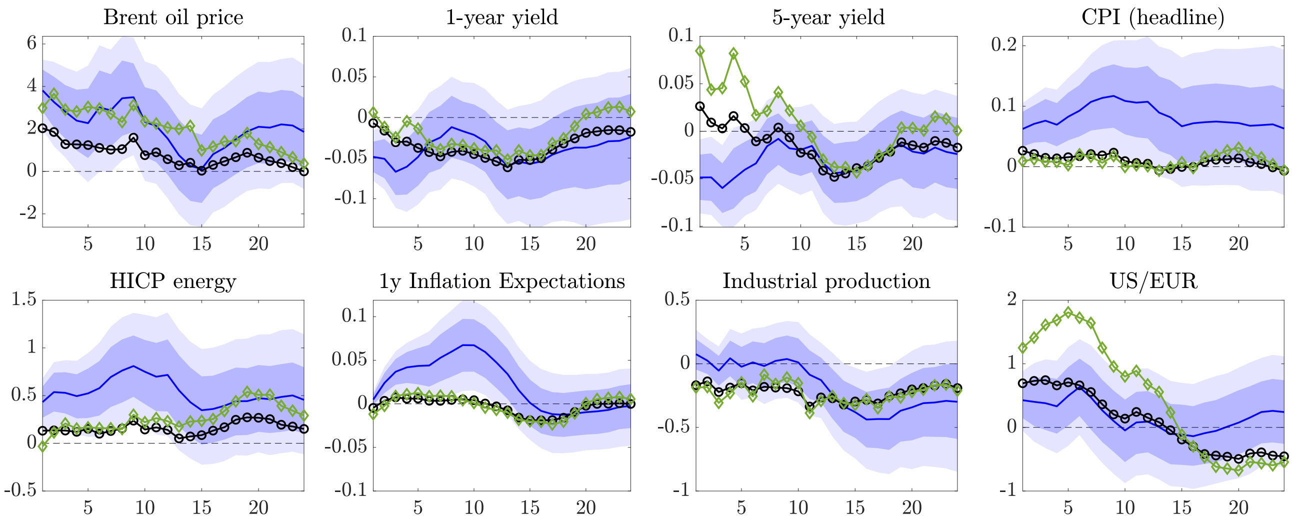
$$\mathcal{A}^*_{\boldsymbol{x}} = (\lambda_\pi \Theta'_{\pi,\nu,\widetilde{\mathcal{A}}} W), \tag{1.23}$$

$$\mathcal{A}^*_{\boldsymbol{z}} = \mathbf{0}., \tag{1.24}$$

which is now characterized by the counterfactual impulse responses of inflation to the identified monetary policy shocks $\Theta_{\pi,\nu,\widetilde{\mathcal{A}}}$ instead of their estimated empirical counterparts $\Theta_{\pi,\nu,\mathcal{A}}$. The results from this exercise are depicted by the green lines in Figure 1.4.

By inspecting the first panel of the first row it becomes apparent that, by construction, the change in the policy rule from the empirical baseline rule (as depicted by the blue lines) to the counterfactual optimal policy rule (as depicted by the green lines) hardly alters the response of the oil price. This is because, under the counterfactual assumptions, the ECB's

Figure 1.4: Impulse responses to an oil supply shock under optimal monetary policy when euro area monetary policy does not affect the Brent oil price



Notes: Impulse response functions to a one standard deviation monetary policy shock showing the point-wise posterior means along with 68% and 90% point-wise credible sets in Blue. The black circled lines show the least-squares approximation of the responses of the endogenous variables under optimal policy with a loss function described in Equation 1.19. The green triangled lines show the least-squares approximation of the responses of the endogenous variables under the optimal problem computed using the impulse responses to a conventional monetary policy and forward guidance shock, which, counterfactually, do not impact the Brent oil price due to a change in the OPEC policy rule.

decisions do not affect the global oil price. In fact, the differences between the blue and green lines for the Brent oil price solely arise from the approximation error of the least-squares approximation. More importantly, when comparing the black circled lines, which depict the optimal policy strategy in the scenario where the ECB's decision do impact the global oil price, to the counterfactual optimal policy, it becomes apparent that to achieve its mandate and thereby implement the optimal allocation for medium-term inflation, monetary policy needs to tighten much more when it does not affect the Brent oil price. This is particularly true at the longer end of the yield curve, as depicted by the response of the 5-year yield under the counterfactual optimal policy scenario in green.

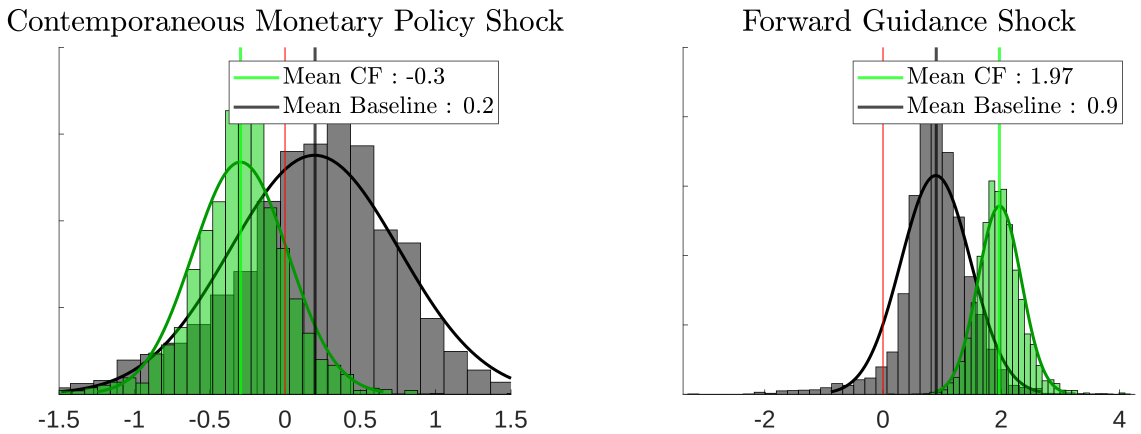
But even under the counterfactual scenario where monetary policy does not affect global energy prices, it still has the tools to almost perfectly stabilize medium-term inflation. In fact, under the optimal counterfactual policy as depicted by the green lines, inflation and inflation expectations are largely stabilized and the estimates reveal that, although it needs to tighten significantly more, this does not mean that monetary policy has to engineer a significantly stronger recession to bring inflation back to target. While puzzling at first sight, this can again be attributed to (local) energy prices being fairly flexible prices and therefore quickly responding to changes in the monetary policy stance. Intuitively, even when monetary policy is not able to affect energy prices on a global level as the Brent oil price (denominated in dollars) is not affected by the counterfactual impulse responses, it can still move local energy prices (denominated in euro), which constitute a large part of the overall HICP basket. A straightforward way to do so would be to engineer an appreciation of the euro as the euro area largely imports its energy goods. The estimates reveal that even if the monetary authority does not impact global energy prices, the optimal strategy in the face of an oil price shock would be to follow this approach. To stabilize inflation the ECB engineers a large appreciation of the euro vis-a-vis the dollar as depicted in the last panel of the second row of Figure 1.4. The induced appreciation then dampens the response of local energy prices despite the surge in global energy prices. This stabilizes HICP inflation and inflation expectations which are shown to be particularly sensitive to changes in local, euro area energy prices (Aastveit et al. (2023), Wehrhöfer (2023)). In other words, in the absence of the global component of the energy price channel, monetary policy would find it optimal to resort to its local counterpart to achieve its objective.

1.4.4 Implications of the global energy price response for the term structure of optimal policy

As shown in Figure 1.4, when the ECB's decision do not affect global oil prices, the optimal response of the ECB to an oil price shock would entail a stronger tightening relative to the baseline optimal policy response where the euro area monetary policy shock do impact global oil prices. This is particularly pronounced at the longer end of the yield curve, implying that the optimal strategy under this scenario would entail a significant amount of forward guidance policies. I can further corroborate this intuition by leveraging the result of McKay and Wolf (2023) who, for a large class of models such as representative and heterogeneous agents New Keynesian models, prove the equivalence between the change from the baseline to the counterfactual (optimal) monetary policy rule and a sequence of policy (news) shocks announced at date 0. While so far the differences between the baseline and counterfactuals have been described as being caused by changes in the policy rule, I now leverage this duality to describe them in terms of the sequence of shocks necessary to mimic this policy rule. This provides additional insights into how the response of (global) energy prices changes the optimal conduct of monetary policy.

Figure 1.5 plots the posterior distribution of the estimated sequence of policy shocks that characterize the optimal policy rule under the baseline impulse responses (black) and the counterfactual impulse responses (green), where the ECB's decisions do not impact the global oil prices. As shown by the black bars, when the ECB's decisions indeed impact oil prices, the shift from the baseline to the optimal policy rule can be characterized on average by a combination of a 0.2 standard deviation contractionary contemporaneous monetary policy shock, which on impact raise the 1-year yield while leaving the longer run yields largely unchanged (see Figure A.6), and a 0.9 standard deviation forward guidance shock, which on impact strongly raises longer run yields and to a lesser degree the 1-year yield (see Figure A.7). This policy mix changes substantially under the counterfactual

Figure 1.5: Baseline and counterfactual shocks that characterize the optimal policy rule

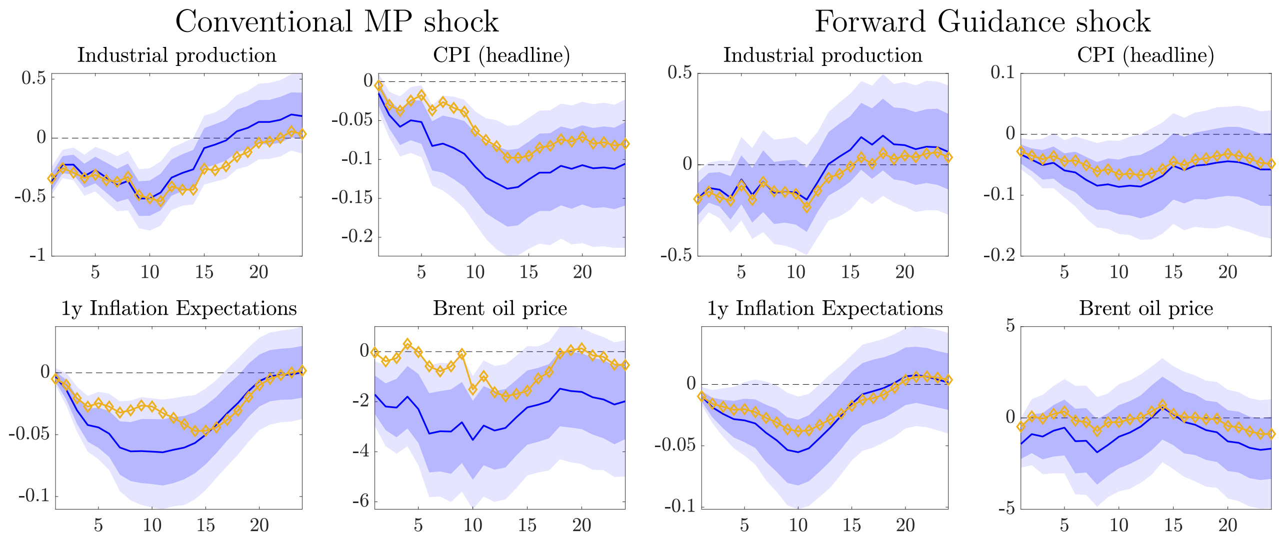


Notes: The figure plots the posterior distribution of the identified monetary policy shocks that characterize the baseline optimal policy rule (black) and counterfactual (CF) optimal policy rule (green), which is computed under the assumption that OPEC stabilizes the Brent oil price, causing euro area monetary policy to not transmit via global oil prices. Shocks are measured in standard deviations

optimal policy, where the ECB's decisions do not impact the global oil price. In this scenario, the optimal policy rule is on average equivalent to a combination of a 0.3 standard deviation expansionary monetary policy shock alongside a whopping 2 standard deviation contractionary forward guidance shock. Therefore the fact that the ECB's decision do impact global energy prices has major implications for the "term structure of monetary policy".

The reason for this shift in strategy lies in the different implications of the global component of the energy price channel for respective policy shocks. On the one hand, as shown on the left-hand side of Figure 1.6, the contemporaneous monetary policy shock significantly decreases the Brent oil prices and this strongly affects the transmission of the contemporaneous monetary policy shock to the HICP. On the other hand, the forward guidance shock is estimated to have a strong effect on the exchange rate, local energy prices as well as output and the HICP, while only slightly depressing global energy prices as measured by the Brent oil price (see Figure A.7). Therefore, as illustrated in Figure 1.6, the importance of the global component of the energy price channel for the transmission of the forward guidance shock to the ECB's objective is estimated to be smaller compared to its contemporary monetary policy shock counterpart. The shift towards forward guidance policy in the ECB's optimal strategy under the assumption that the ECB's decisions do not impact global oil prices can therefore be interpreted along the lines of the results of MW who show that under optimal monetary policy, "the policymaker will rely most heavily on the tools [...] that are best suited to offset the perturbation to its targets" ((McKay and Wolf 2022a, p.9)). When the ECB's decisions do not affect global oil prices, the optimal strategy implies a change in the policy rule that is equivalent to putting more weight on forward guidance shocks. This is because shutting down the global component of the energy

Figure 1.6: IRFs and counterfactual IRFs to a Conventional EA Monetary policy and a Forward Guidance shock



Notes: Impulse response functions to a one standard deviation conventional monetary policy shock showing the point-wise posterior means along with 68% and 90% point-wise credible sets in blue. The golden lines depict the point-wise posterior means under the assumption that OPEC, counterfactually, stabilizes the Brent oil price. Because I added the 5-year yield to the SVAR model and I jointly identify a contemporaneous monetary policy and a forward guidance shock, the estimated IRFs differ somewhat from the ones presented in Figure 1.1, where I only identify a generic monetary policy shock.

price channel affects their transmission to HICP by less, making them more "suited to offset the perturbation to its targets" in the counterfactual scenario.

Given that the ECB in the counterfactual scenario heavily relies on forward guidance policies to fight the oil price shock, the slightly expansionary contemporaneous monetary policy shock is then used to smooth out any overshooting from the target that would arise if the central bank were to purely rely on forward guidance. An arguably unintended but probably highly welcome benefit of this strategy is, that because the optimal policy rule is now characterized by a combination of an expansionary contemporaneous and a contractionary forward guidance shock, output is roughly stabilized at the level that would prevail in the baseline optimal policy case despite the significant tightening at the longer end of the yield curve.

I conjecture that only if the ECB's decisions do not move both global and local energy prices, it would be necessary to engineer a large recession to (optimally) stabilize inflation in the face of an energy price shock. This counterfactual would be arguably very difficult to phrase as an empirical policy rule counterfactual, its implementation would require a fully-fledged structural model that (i) models the interplay of monetary policy and energy prices (ii) matches their empirically identified transmission. I therefore leave the estimation of this counterfactual for future research as such a model is beyond the scope of this paper.

1.5 Application to the most recent energy price surge

In this section, I assess the role of euro area monetary policy in the recent energy-driven inflation episode in light of the findings in the previous sections. In particular, as a first step, I use the BPSVAR model to compute the historical decomposition of the underlying time series to estimate how much of the (energy price) inflation and the subsequent disinflation can be explained by the cumulative effects of exogenous monetary policy shocks. Second, I estimate how the euro area economy would have evolved during this inflationary episode if the ECB, counterfactually, had followed a different endogenous monetary policy rule, which optimally achieves its medium-term price stability mandate. To that end, I apply the methods in Caravello et al. (2024) who show how to extend the counterfactual methods of MW, that are primarily concerned with impulse responses, to estimate the counterfactual evolution of the economy in historical episodes.

1.5.1 The role of exogenous monetary policy shocks

First, I examine the role played by deviations from the policy rule (i.e. exogenous monetary policy shocks) in the recent inflation and energy price surge. To do so, as a first step, I reestimate the BPSVAR model introduced in section 1.2.2 using data up until October 2023.[17] To explicitly model the extraordinary volatility in the data induced by the COVID-19 pandemic I incorporate the "Pandemic Prior" approach of Cascaldi-Garcia (2022) into the BPSVAR model in section 1.2.2. Effectively this approach introduces a series of exogenous variables and corresponding parameters into the BPSVAR model to capture the effects of the pandemic in 2020.[18] I then compute the historical decomposition of the extended BPSVAR model to estimate the contributions of exogenous monetary policy shocks to the rise and fall of (energy price) inflation. In particular, under the assumption of stationarity and invertibility (see Fernández-Villaverde et al. (2007) for a discussion), the SVAR in Equation (1.1) can be used to compute the structural vector moving average (SVMA) representation of the variables

$$\boldsymbol{y}_t = \underbrace{\sum_{\ell=0}^{\infty} \Theta_{\ell,\epsilon,\mathcal{A}}\, \eta_{t-\ell}}_{\text{Current and past shocks}} + \underbrace{\sum_{\ell=0}^{\infty} \Theta_{\ell,\mathcal{A},D}\, \boldsymbol{D}_{t-\ell}}_{\text{Contribution of exo. variables}} \tag{1.25}$$

where $\Theta_{\ell,\eta,\mathcal{A}}$ is a matrix of all impulse responses of the variables in $\boldsymbol{y}$, computed under the baseline policy rule $\mathcal{A}$, to the structural policy and non-policy shocks $\boldsymbol{\eta} = (\boldsymbol{\nu}', \boldsymbol{\epsilon}')'$ which occurred ℓ periods ago. The matrix $\Theta_{\ell,\mathcal{A},D}$ furthermore captures the impulse responses to the exogenous variables in $\boldsymbol{D}$. Under the assumption of invertibility of the SVAR model, this infinite order moving average representation can be truncated at the first observation

[17] As shown in Figures A.10 and A.11 the impulse responses to the monetary policy shocks are robust to extending the data to include the pandemic.

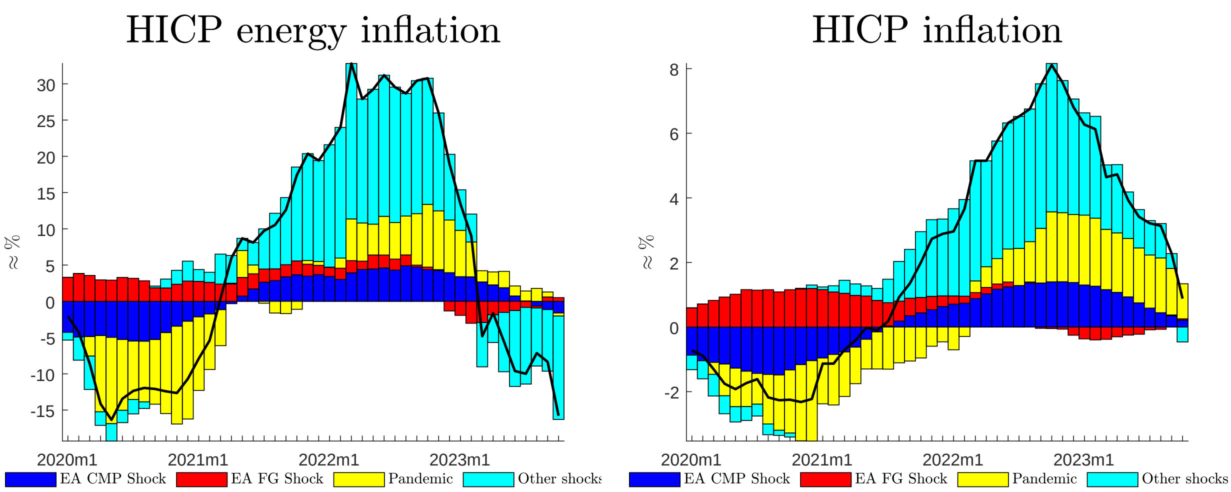
[18] Furthermore, to preserve the stationarity of the specification I replace all price levels with year-on-year inflation rates. I also include the 5-year bund yield as I aim to identify forward guidance shocks along the lines of section 1.4.2.

$\boldsymbol{y}_0$ without an approximation error. As discussed by CMW the reason for this is that, under the assumption of invertibility, the information contained in shocks occurring before $\boldsymbol{y}_0$ are already embedded into $\boldsymbol{y}_0$ and therefore a projection starting in $\boldsymbol{y}_0$ that leverages the SVAR representation perfectly captures their joint effect of these shocks on future values of the endogenous variables $\boldsymbol{y}_{1:t}$. In particular, denoting $\boldsymbol{y}^{\mathcal{A}}_{t,0} = E^{\mathcal{A}}_0[\boldsymbol{y}_t]$ as the expected value of $\boldsymbol{y}_t$ from the point of view of an agent observing data until $\boldsymbol{y}_0$ and knowing the propagation mechanisms of the model including the baseline policy rule $\mathcal{A}$, one can write

$$\boldsymbol{y}_t = \underbrace{\boldsymbol{y}^{\mathcal{A}}_{t,0}}_{\text{initial condition}} + \underbrace{\sum_{\ell=0}^{t-1} \Theta_{\ell,\eta,\mathcal{A}}\, \eta_{t-\ell}}_{\text{Current and past shocks}} + \underbrace{\sum_{\ell=0}^{t-1} \Theta_{\ell,D,\mathcal{A}}\, \boldsymbol{D}_{t-\ell}}_{\text{Contribution of exo. variables}} , \qquad (1.26)$$

where, again under the assumption of invertibility, the expectations embedded in the contributions of the initial conditions can be consistently estimated by an SVAR-based forecast. In summary, the state of the economy (i.e. the level of all variables in $\boldsymbol{y}$) can be decomposed into the effects of exogenous variables, the initial conditions, the cumulative effects of the (current and past) policy and non-policy shocks.

Figure 1.7: Exogenous monetary Policy shocks as drivers of the recent energy price inflation surge



Notes: Historical decomposition excluding the contribution of the constant and initial conditions (black line) alongside the contribution of current and past EA conventional monetary policy shocks (blue bars), forward guidance shocks (red bars), effects of the 2020 COVID pandemic (yellow bars) and the *sum of the contributions* of all current and past other (non-EA MP) shocks in the system (turquoise bars). I obtained the estimates by computing the historical decomposition using the SVMA representation for each draw. I plot the resulting point-wise means for each shock and variable.

Figure 1.7 presents the results for the historical decomposition for the period from 2020 until the end of the sample in October 2023. In particular, the black solid lines depict the evolution of headline inflation and energy price inflation alongside the contribution of exogenous ECB monetary policy shocks, which are depicted by the blue and red bars.[19] Furthermore, the "Pandemic priors" approach of Cascaldi-Garcia (2022) allows for estimating the effects of the onset and fading out of the pandemic as depicted by the yellow bars. During the year 2020 and early 2021, the fall in energy prices and HICP inflation can largely be attributed to the pandemic. Through the lens of the BPSVAR, this fall in energy prices and inflation in 2020 and early 2021 was in part amplified (dampened) by the ripple effects of a sequence of contractionary contemporaneous policy shocks (forward guidance shocks) occurring in late 2019. However, the reluctance of the ECB to make the potential switch to a more contractionary conventional and forward guidance policy in mid-2021 as energy prices rapidly rose, led to a positive contribution of monetary policy to the surge in energy price and headline inflation from mid-2021 onwards. In fact, this expansionary policy stance is estimated to correspond to a sequence of expansionary conventional monetary policy and forward shocks. The effects of these shocks can be seen by observing that the cumulative contribution of monetary policy, depicted by the sum of the red dark blue bars, quickly accelerates into large positive territory during 2021 and 2022. At its peak monetary policy shocks are estimated to have, potentially inadvertently, contributed to an increase of HICP inflation by roughly 2% and energy price inflation by roughly 5%. The contribution of these expansionary shocks in the face of rapidly increasing inflation then slowly fades away during 2023 as the policy stance starts to tighten again. Lastly and for later reference, it is important to note a sizable component of the surge in energy prices and inflation in 2022 can be attributed the yellow bars which depict the effects of the fading away of the pandemic and the corresponding reopening of the economy. Given the estimated correlation structure in the data, these effects would have been predictable through the lens of the model already in 2021.

1.5.2 The role of endogenous monetary policy in the recent inflation surge

Figure 1.7 showed that, when viewed through the lens of the BPSVAR model, exogenous expansionary monetary policy shocks contributed to the initial rise in energy prices and HICP inflation since early 2021 with the effects receding only in the beginning of 2023. In other words, the exogenous deviations from the policy rule, potentially inadvertently contributed to the rise and fall of energy prices and inflation. However, this is not the full picture because the role of endogenous monetary policy may be very different. In this subsection, I therefore again switch the question and ask, how the economy would have evolved if the endogenous component of monetary policy — the policy rule — had been optimal in light of the energy price surge. Traditionally the answer to such a question has been addressed by first, setting up a theoretical (DSGE) model of the economy and

[19] For the sake of brevity I plot the evolution of these variables net of the contribution of the constant and initial conditions, whose contribution is roughly constant and for the initial conditions is small.

specifying the entire menu of stochastic shocks that drive this economy, second, estimating this model using full-information methods as described for instance in Herbst and Schorfheide (2016) to back out the structural shocks of the economy and third, altering the policy block and simulating the counterfactual economy using the counterfactual SVMA representation alongside the structural shocks from the second step (see Christiano et al. (2015) for an example).

This paper instead takes a different route. In particular, to compute the answer to this question I again leverage a recent breakthrough in the sufficient statistics approach to optimal policy. Specifically, I employ the approach of Caravello et al. (2024, henceforth CMW) who extend the MW impulse response counterfactuals to historical episodes. The key insight of CMW is that combining the approach of MW with the frequently made assumption of invertibility of the SVAR model (see Plagborg-Møller and Wolf (2021) for a discussion), allows the researcher to compute policy rule counterfactuals going beyond impulse responses.[20] First, in line with the DSGE-model-based approach sketched above, they observe that the counterfactual evolution of the economy can, in theory, be retrieved from a counterfactual structural vector moving average (SVMA) representation of the economy, where counterfactual impulse responses to all structural shocks are computed under the counterfactual policy rule. Second, they prove that identifying all the true structural shocks and the corresponding impulse responses is not a necessary requirement for the estimation of the counterfactual SVMA representation as the reduced form representation proves sufficient. Third, they show how to coherently manipulate the initial conditions of the counterfactual SVMA representation to ensure that, in the underlying data-generating structural model, agents' expectations incorporate the counterfactual policy rule. As in the case of the counterfactual impulse response, this method allows for answering the question at hand, without having to precisely specify the structure model that has generated the data.

More formally, the aim is to simulate the evolution of the euro area economy under the assumption that, starting at point $t^\star$, the ECB would have conducted the policy according to the optimal rule embedded in $\mathcal{A}_x^\star$ and $\mathcal{A}_z^\star$ in Equation (1.18). Therefore the object of interest is the counterfactual SVMA representation of the economy truncated at point $t^\star$.

$$\boldsymbol{y}_t = \underbrace{\sum_{\ell=0}^{t-t^\star} \Theta_{\ell,\eta,\mathcal{A}^\star}\, \eta_{t-\ell}}_{\text{contrib. of new shocks after } t^\star} + \underbrace{\boldsymbol{y}_{t,t^\star}^{\mathcal{A}^\star}}_{\text{contrib. of initial condition at } t^\star} \tag{1.27}$$

The first term in Equation (1.27) captures the effects of shocks $\eta_{t-\ell}$, which occur after the change in the policy at $t^\star$, on the economy. The propagation of these shocks is altered from $\Theta_{\ell,\eta,\mathcal{A}}$ to $\Theta_{\ell,\eta,\mathcal{A}^\star}$ because euro area monetary policy no longer follows the baseline rule

[20] In a simulation study using the Smets and Wouters (2007) model, CMW show that their approach approximates the true counterfactual very well, even if the conditions for invertibility of the SVAR model do not hold.

embedded in $\boldsymbol{\mathcal{A}}$ but the optimal rule, which minimizes the loss function in line with the mandate and which is embedded in $\boldsymbol{\mathcal{A}}^\star$ in Equation (1.18). The second term in Equation (1.27) on the other hand represents the contribution of the initial conditions starting at point $t^\star$.[21] In particular, denoting $\boldsymbol{y}_{t,t^\star}^{\mathcal{A}^\star} = E_{t^\star}^{\mathcal{A}^\star}[\boldsymbol{y}_t]$ as the expected value of $\boldsymbol{y}_t$ from the point of view of an agent observing data until $\boldsymbol{y}_{t^\star}$ and knowing the propagation mechanisms of the model including the optimal policy rule described by $\boldsymbol{\mathcal{A}}^\star$. CMW show how to estimate these two terms using an SVAR model. Below I provide the basic intuition for two of two of their identification results which are then leveraged to estimate these components.

Capturing the counterfactual effects of new shocks hitting the economy after the change in the policy rule at $t^\star$ is conceptually straightforward. It would suffice to first identify all structural shocks $\boldsymbol{\eta}$ in the economy and their corresponding impulse responses under the baseline policy rule $\Theta_{\ell,\eta,\mathcal{A}}$ and then treat the impulse responses to each shock individually along the lines of the procedure sketched in Section 1.3.1 to produce $\Theta_{\ell,\eta,\mathcal{A}^\star}$. The first identification result of MW is that, under the assumption of invertibility, one does not need to identify all structural shocks. Instead, it suffices to apply the procedure to each of the reduced form impulse responses $\Psi_{l,u,\boldsymbol{\mathcal{A}}}$ corresponding to the reduced form errors of the SVAR u_t to arrive at $\Psi_{l,u,\boldsymbol{\mathcal{A}}^\star}$. While CMW provide formal proof of this result, the intuition for this result is arguably straightforward. Regardless of whether the evolution of the economy is described by its SVMA representation or by the reduced form Wold-representation, the implied path of the endogenous variables is identical. In particular, under invertibility, the reduced form residuals (reduced form impulse responses) have a one-to-one function of the structural shocks (structural impulse responses)

$$\Theta_{l,u,\boldsymbol{\mathcal{A}}} = \Psi_{l,u,\boldsymbol{\mathcal{A}}}\mathcal{P}^{-1}. \qquad \eta_t = \mathcal{P}^{-1}u_t \tag{1.28}$$

Substituting these definitions into the first component of Equation (1.27)

$$\boldsymbol{y}_t = \sum_{\ell=0}^{t-t^\star} \Theta_{\ell,\eta,\mathcal{A}^\star}\, \eta_{t-\ell} + \boldsymbol{y}_{t,t^\star}^{\mathcal{A}^\star} = \sum_{\ell=0}^{t-t^\star} \Psi_{\ell,\eta,\mathcal{A}^\star}\mathcal{P}^{-1}\eta_{t-\ell} + \boldsymbol{y}_{t,t^\star}^{\mathcal{A}^\star} = \sum_{\ell=0}^{t-t^\star} \Psi_{\ell,\eta,\mathcal{A}^\star}\, u_{t-l} + \boldsymbol{y}_{t,t^\star}^{\mathcal{A}^\star}, \tag{1.29}$$

it becomes apparent that knowledge of the counterfactual reduced form impulse responses $\Psi_{\ell,\eta,\mathcal{A}^\star}$ and reduced form innovations $u_{t-\ell}$ is sufficient to evaluate the first component.[22]

The second identification result of CMW deals with the initial conditions component $\boldsymbol{y}_{t,t^\star}^{\mathcal{A}^\star} = E_{t^\star}^{\mathcal{A}^\star}[\boldsymbol{y}_t]$. Intuitively, when simulating the economy using the counterfactual SVMA representation, this term ensures that the chosen policy path $\boldsymbol{z}_{t^\star:T}^{\mathcal{A}^\star}$ is such that the new policy rule is also embedded into the expectations of the private sector. In other words, the

[21] I leave aside the treatment of the exogenous variables because in this application the initial effects of all exogenous variables arising from the "Pandemic Priors" of Cascaldi-Garcia (2022) approach materialize before my choice of $t^\star$. As such their effect on all variables after $t^\star$ ($\boldsymbol{y}_{t^\star:T}$) is embedded into the initial conditions term $\boldsymbol{y}_t^{\mathcal{A}^\star}$.

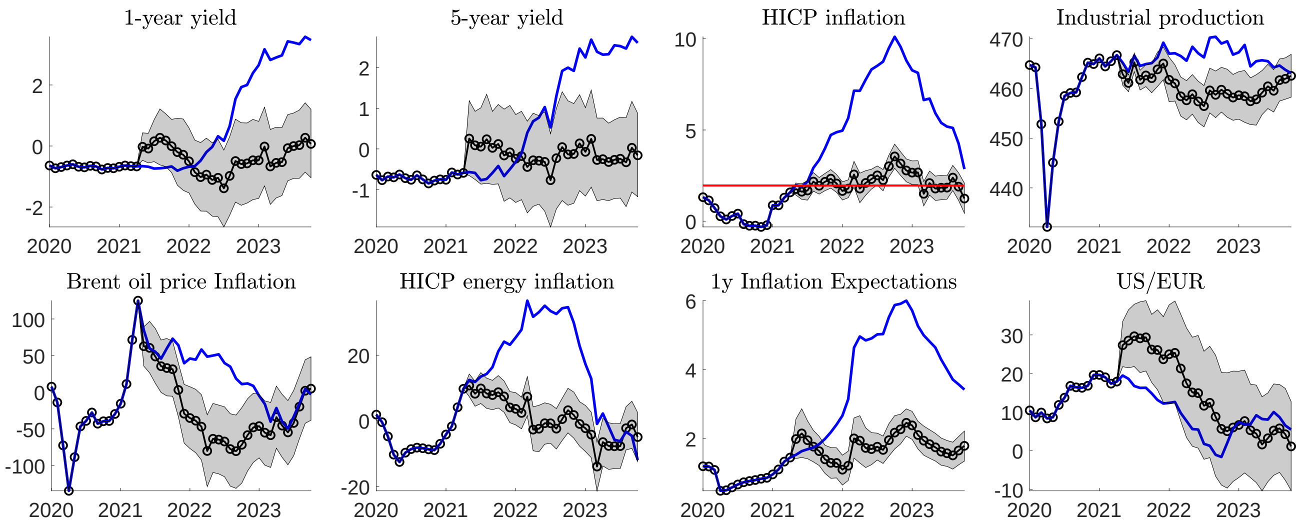
[22] CMW show that the mapping between the reduced and structural form also applies to the counterfactual impulse responses $\Theta_{l,u,\boldsymbol{\mathcal{A}}^\star} = \Psi_{l,u,\boldsymbol{\mathcal{A}}^\star}\mathcal{P}^{-1}$.

policymaker revises the planned policy path in $\boldsymbol{z}^{\mathcal{A}^\star}_{t^\star:T}$ to ensure that current and expected values of the non-policy variables $\boldsymbol{x}^{\mathcal{A}^\star}_{t^\star:T}$ are such that they are consistent with the new optimal policy rule. CMW observe that, under invertibility, the econometrician can recover the baseline version of these expectations (i.e. the forecast formed under the policy rule described by $\boldsymbol{\mathcal{A}}$) by computing the unconditional SVAR forecast starting from $y_{t^\star}$ to the final period T, which yields $\boldsymbol{y}^{\mathcal{A}}_{t^\star:T} = (\boldsymbol{z}^{\mathcal{A}'}_{t^\star:T}, \boldsymbol{x}^{\mathcal{A}'}_{t^\star:T})'$. They prove that treating these forecasts analogous to impulse responses, which implies substituting out $(\boldsymbol{\Theta}_{x,\epsilon,\mathcal{A}} \times \boldsymbol{\epsilon})$ and $(\boldsymbol{\Theta}_{z,\epsilon,\mathcal{A}} \times \boldsymbol{\epsilon})$ in Equation (1.9) for their unconditional forecast analogs $(\boldsymbol{z}^{\mathcal{A}'}_{t^\star:T}, \boldsymbol{x}^{\mathcal{A}'}_{t^\star:T})'$, allows the researcher to recover the correct expectations and path of the variables in $\boldsymbol{x}$ and $\boldsymbol{z}$ that would prevail under the optimal policy rule. Intuitively, impulse responses are a particular form of a conditional forecast and the same proof that MW use to derive the counterfactual impulse responses also applies to an unconditional forecast.

I utilize the approach of CMW in order to compute the counterfactual evolution of the euro area economy under the assumption that ECB conducted optimal monetary policy from April 2021 onwards. I choose April 2021 as the starting point for this exercise as it marks the period just before headline inflation in the euro area exceeded the 2% target of the ECB. Furthermore, in that month the Euro area economy witnessed a whopping 120% increase in the oil price in year-on-year terms. As in section 1.4.1, optimal monetary policy is defined as the policy rule that yields an allocation that optimally achieves the mandate of medium-term inflation targeting. Figure 1.8 plots the counterfactual evolution of the economy under the optimal policy rule (black dotted lines) alongside the actual evolution (blue solid lines). Three interesting conclusions can be drawn from this analysis. First, the rapid rise in energy prices and the ensuing headline inflation could have been curtailed to a large degree if the ECB had credibly committed to optimal medium term inflation targeting at the onset of the rise in inflation.

In the counterfactual scenario, inflation expectations and energy prices experience a substantially smaller rise, and the headline inflation rate largely remains at the ECB's target as defined by the mandate. Second, the larger tightening at the beginning would have resulted in a in a more severe contraction in output, as proxied by industrial production, which would have later recovered to its current actual level. The counterfactual fall in industrial production relative to the baseline is around 10% at its trough point. A back-of-the-envelope calculation taking into account the relative volatilities of GDP and industrial production suggests that, at its lowest point, GDP would have been roughly 4% lower if the ECB had followed the optimal strategy (see Georgiadis et al. (2024) for a similar way to convert changes in industrial production into changes of GDP). Third, the counterfactual evolution of 1- and 5-year interest rates suggests that the optimal policy response to the recent energy-driven inflation episode would consist of a combination of swift interest rate hikes and a commitment to maintaining higher rates until 2022. This rise in interest rates would would have persistently appreciated the euro vis-a-vis the dollar and have also directly helped to bring global energy prices. Both of these effects ultimately curtail local energy

Figure 1.8: Counterfactual evolution of the endogenous variables under optimal ECB policy



Notes: The figure plots the time series of the endogenous variables (blue solid line) alongside the estimated counterfactual evolution (black circled line) of these variables under the assumption that, from April 2021 onwards, the ECB would have conducted optimal monetary policy as described in Equation 1.19 and thereby minimizes the weighted squared deviations of "medium-term" inflation from the target. Similar to Christoffel et al. (2008) I assume an inflation target of 1.95% (red solid line) which is "below but close to 2%". Black shaded areas correspond to 68% point-wise credible sets.

price inflation, which was arguably the main driver of HICP inflation during the sample period. Importantly, a comparison of the counterfactual and actual paths of the 5-year interest rate implies that the ECB would not have needed to raise interest rates to its current historical levels if it had swiftly committed to tighter policy at the onset of rapidly rising energy prices. This is especially interesting as one of the main criticisms against the ECB has been the delayed response to utilize forward guidance in the presence of accelerating inflation.

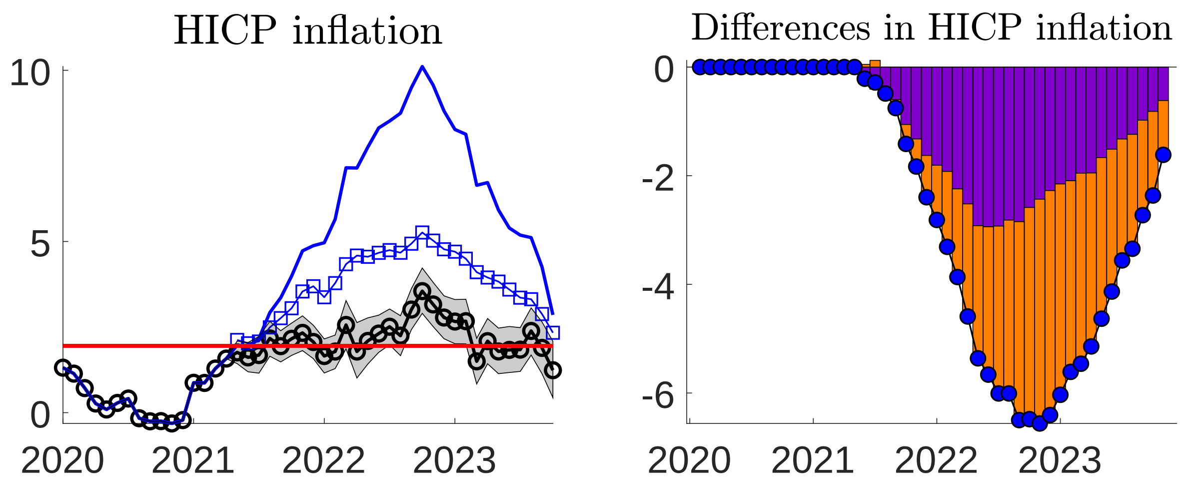
In light of this discussion of the results in Figure 1.8, at least two open questions remain. First, it may be puzzling why the estimates suggest that the ECB should have committed to the swift tightening already in 2021 although inflation was not yet at its peak and just crossed the threshold of 2%. Second, it may at first glance seem surprising that this change in strategy and the corresponding initial rate hike would stabilize inflation at its medium-term target even though interest rates in the counterfactual scenario only exceed their baseline counterparts until 2022.

Regarding the first question, the reason for the initial rapid increase in interest rates can be traced back to the initial conditions in April 2021. In particular, the blue squared line in the left-hand side panel of Figure 1.9 shows the expected path for HICP inflation of an agent observing data until April 2021 and knowing the propagation of the model under the baseline policy rate described by $\boldsymbol{\mathcal{A}}$.[23] It becomes apparent that, even in the

[23] I provide the estimates of the component relating to propagation of the initial conditions for the remaining variables in Figure A.12 in the Appendix.

absence of further structural shocks, the inflation rate would have been projected to peak at roughly 5% in the medium term given the initial conditions. By inspecting the yellow bars in Figure 1.7, it becomes clear that this projected increase in inflation is almost entirely explained by the fading effects of the 2020 pandemic, which initially decreased inflation substantially in 2020 but then caused it to overshoot. Through the lens of the model, these effects were predictable given the data until 2021 April. To realign the expectations with the commitment of the ECB under optimal policy to stabilize medium-term inflation at roughly 2%, the optimal strategy implies that the ECB needs to commit to a substantial early rate hike.

Figure 1.9: The role of the (optimal) policy rule for the evolution of euro area



Notes: The left-hand side figure plots the time series of the endogenous variables (blue solid line) alongside the estimated counterfactual evolution (black circled line) of these variables under the assumption that, from April 2021 onwards, the ECB would have conducted optimal monetary policy as described in Equation 1.19 and thereby minimizes the weighted squared deviations of "medium-term" inflation from the target. Furthermore, the blue solid line with diamonds represents the expected evolution (forecast) of the economy under the assumption that the ECB follows the baseline policy rule also from April 2021 onwards.
The right-hand side decomposes the difference between the observed evolution of the economy (blue lines in Figure 1.8) and the counterfactual evolution of the economy (black lines in Figure 1.8). These differences arise because of the change in the policy rule from the baseline to the optimal policy rule, which is assumed to happen in April 2021. They can be decomposed into differences in deterministic components (i.e. the expected path of the endogenous variables without any further shocks) which are depicted in purple and differences in the stochastic component (i.e. the way that new, unexpected shocks propagate) which are depicted in orange.

To address the question of how the ECB's strategy change stabilized inflation despite interest rates not consistently exceeding those of the baseline policy, I propose decomposing the discrepancies between counterfactual and actual economic trajectories. In particular, these differences can be decomposed into a component that is related to the change in the propagation of the initial conditions (i.e. the expectations about the future evolution of the economy) and a component that is related to the change in the propagation of incoming structural shocks. Intuitively, to decompose the difference between the counterfactual and the observed evolution of the economy into these two parts, I estimate Equation 1.27 under the baseline policy rule $\mathcal{A}$ and the counterfactual policy rule $\mathcal{A}^{\star}$ and take the difference between the respective contributions of the components.

The right-hand side panel of Figure 1.9 depicts the contribution of these two components to the difference in HICP.[24] It becomes apparent that altering the initial conditions by rapidly increasing interest rates in the early parts of 2021 already goes a long way in curbing the surge in inflation as depicted by the purple bars. But ultimately altering the transmission of structural shocks, such as the oil price shock in Section 1.3.2, allows the ECB to achieve its mandate in the medium term as depicted by the contribution of the orange bars. Similar to simulations in theoretical models that feature optimal inflation targeting policy, consumer price inflation and inflation expectations react much less to newly arriving structural shocks when the central bank credibly commits to following an inflation targeting strategy. Intuitively, as agents know that monetary policy is going to curb inflation, they do not expect inflation to rise in the first place, which would have prevented euro area monetary policy from having to excessively tighten in 2023 to counteract galloping inflation.

To summarize, this analysis reveals that the ECB could have stabilized inflation in light of the energy price surge. This would have come at the cost of arguably slightly lower output and economic activity in 2022 and early 2023. But already at the end of 2023 economic activity under the counterfactual policy rule is estimated to be at the same level as its baseline counterpart. I attribute the fact that the output losses necessary to fight a surge of inflation of up to 10% are rather mild, to the ability of the ECB to directly impact prices of energy goods traded at the global market and prices of energy goods faced by euro area consumers. Therefore it does not need to create an excessively large recession to optimally fight a surge in inflation that is due to rising energy prices.

Naturally, the applied criterion for household welfare defines to which extent the estimated trade-off is optimal in a welfare theoretic sense. As this analysis solely focuses on the ECB's primary mandate the approach is silent concerning this question. Estimating the appropriate weight on employment and output and other possible objectives in the loss function is left to future research.

1.6 Conclusion

This paper provides evidence that decisions of the ECB affect energy prices globally and that this has strong implications for the conduct of monetary policy. Building on Ider et al. (2023) I document that energy prices are materially affected and move by more than the headline consumer price index in response to monetary policy shocks. This is in line with the intuition that energy prices, like the global oil price, are much less sticky than other prices in the consumer basket. Second, using the methods to estimate Lucas-critique robust counterfactuals put forward in McKay and Wolf (2023), I establish that a sizeable

[24] I provide the results of this decomposition for the remaining variables in Figure A.13 of the Appendix.

fraction of the effects of monetary policy on the headline consumer prices is mediated via the energy price component of the HICP. In addition, short-run inflation expectations seem to be influenced by the ECB's monetary policy decisions primarily through its effect on energy prices. As a result, energy prices assist the central bank in achieving its mandate of price stability by moving quickly and strongly in response to monetary policy. Third, I analyze the more normative question of how monetary policy should optimally respond to an energy price shock, which is proxied for by an oil price shock. I show that historically, the ECB "looked through" the persistent increase in prices caused by this shock. However, to optimally achieve its mandate of medium-term price stability, the estimates suggest that the monetary policy stance should have been more restrictive, especially in the short term. Nevertheless, the analysis presented in this paper reveals that only a small additional, front-loaded tightening is necessary to substantially stabilize CPI inflation and inflation expectations. I show that this can be attributed to the fact that energy prices are indeed very sensitive to changes in ECB's policy stance. Lastly, I document that, by committing to (optimally) stabilize medium-term inflation, the ECB could have prevented the recent energy-price-driven inflation surge of 10% at the cost of increasing interest rates already in 2021 and, at its lowest point, an approximately 4% lower level of GDP. The results of this paper indicate that this arguably mild trade-off between output and inflation can be traced back to the ability of the ECB to influence energy prices. Therefore energy prices, when managed correctly, can be the ECB's friend, not its foe.

Chapter 2

What goes around comes around: How large are spillbacks from US monetary policy?

with Max Breitenlechner and Georgios Georgiadis

Abstract

Spillovers from US monetary policy entail spillbacks to the domestic economy. Applying counterfactual analyses in a Bayesian proxy structural vector-autoregressive model we find that spillbacks account for a non-trivial share of the slowdown in domestic real activity following a contractionary US monetary policy shock. Spillbacks materialise as a monetary policy tightening depresses foreign sales and valuations of US firms so that Tobin's q/cash flow and stock market wealth effects impinge on investment and consumption. Net trade does not contribute to spillbacks because US monetary policy affects exports and imports similarly. Geographically, spillbacks materialise through advanced rather than emerging market economies.

Keywords: US monetary policy, spillovers, spillbacks, Bayesian proxy structural VAR models.
JEL-Classification: F42, E52, C50.

2.1 Introduction

Much empirical work, as well as prominent policy debates, suggest that US monetary policy spillovers are large and an important driver of business cycles and financial conditions in the global economy (Banerjee et al. 2016; Dedola et al. 2017; Bräuning and Sheremirov 2019; Iacoviello and Navarro 2019; Vicondoa 2019; Degasperi et al. 2020). At the same time, it has been argued that the Federal Reserve has exhibited "benign neglect" regarding its international effects (Eichengreen 2013, p. 87). For example, after the Global Financial Crisis some policymakers complained that US monetary policy measures aimed at stabilising the domestic economy elicited waves of capital flows and accentuated financial market volatility in the rest of the world (Rajan 2013).

Against this background, some policymakers have argued the Federal Reserve should internalise its effects on the rest of the world (RoW) in the calibration of its monetary policy (Rajan 2016). The Federal Reserve has responded that it already does so implicitly: "Actions taken by the Federal Reserve influence economic conditions abroad. Because these international effects in turn spill back on the evolution of the US economy, we cannot make sensible monetary policy choices without taking them into account" (Fischer 2014); similarly: "The Fed recognizes that its own policies do have international spillovers, and, in turn because they affect global performance, they are going to have spillbacks to US economic performance" (Yellen 2019). Similar statements have been made by Shin (2015) and Carney (2019). However, to the best of our knowledge, no rigorous analysis of spillbacks from US monetary policy exists in the literature. This paper fills this gap.

We define spillbacks as the difference between the actual domestic effects of US monetary policy and those in a counterfactual in which the spillovers to RoW real activity are nil. Our analysis suggests that spillbacks are large: When spillovers to the RoW in response to a contractionary US monetary policy shock are counterfactually absent, the slowdown in real activity and the drop in consumer prices in the US are substantially smaller.

Regarding economic transmission channels our analysis suggests that spillbacks materialise through Tobin's q/cash flow and stock market wealth effects. In particular, contractionary US monetary policy depresses US firms' valuations and cash flows as revenues from sales to the RoW decline, inducing them to cut back investment; that cash flow contributes to the domestic effects of monetary policy is consistent with recent work on the determinants of firm investment (Cao et al. 2019; Lian and Ma 2021; Drechsel 2023). Similarly, as contractionary US monetary policy depresses US and foreign equity prices, the value of US households' portfolios is reduced, which triggers a drop in consumption through stock market wealth effects; that wealth effects contribute to the domestic effects of monetary policy is consistent with recent work on heterogeneous-agent New Keynesian models (Kaplan et al. 2018; Alves et al. 2020). In contrast to consumption and investment, net trade does not contribute to spillbacks because US monetary policy affects exports and imports to a similar degree. Finally, spillbacks to consumer prices materialise primarily as a US monetary policy contraction puts downward pressure on global commodity prices and thereby reduces US import prices.

Regarding geographic transmission channels our analysis suggests that spillbacks materialise more through advanced economies (AEs) than through emerging market economies (EMEs). Specifically, we can replicate the counterfactual in which US monetary policy spillovers to the entire RoW are nil by precluding spillovers only to AEs. In contrast, when only spillovers to EMEs are precluded we replicate the baseline in which spillovers to the RoW as a whole are unconstrained. This finding is consistent with the relative exposure of US firms' foreign sales and US holdings of foreign equity across AEs and EMEs. An important caveat to this finding is that it is based on the entire sample period from 1990 to 2019. As the importance of EMEs has been growing over time, their role for spillbacks from US monetary policy may be larger at the current juncture. The analysis of time variation in spillbacks is left to future research.

That spillbacks from US monetary policy materialise primarily through AEs and not EMEs suggests there may be a case for international monetary policy coordination (Engel 2016). In particular, we find that while real activity spillovers from US monetary policy are contractionary in both AEs and EMEs, consumer prices fall in AEs but rise in EMEs. Thus, while US monetary policy spillovers do not induce trade-offs between output and inflation stabilisation in AEs, they do so in EMEs. This implies that global welfare could benefit if US monetary policy internalised spillovers to EMEs independently of spillbacks.

These findings are obtained from counterfactual analyses in two-country vector-autoregressive (VAR) models for the US and the RoW. We adopt the Bayesian proxy structural VAR framework of Arias et al. (2018,0) and identify a US monetary policy shock using the high-frequency interest rate surprises on FOMC meeting dates as in Gertler and Karadi (2015) as well as Caldara and Herbst (2019) but cleansed from central bank information effects as in Jarocinski and Karadi (2020). We consider two complementary approaches to construct counterfactual impulse responses in which the real activity spillovers from US monetary policy to the RoW are nil: (i) Structural scenario analysis (SSA), and (ii) minimum relative entropy (MRE).

In SSA we identify two additional shocks that represent a convolution but together capture the universe of RoW structural shocks by combinations of zero, sign and magnitude restrictions. We use these two RoW shocks to offset the spillovers from the US monetary policy shock in the counterfactual. Technically, SSA indicates how US variables would evolve if a series of RoW shocks materialised along the impulse response horizon that offset the effect of the US monetary policy shock on RoW real activity. SSA is a point of contact with existing literature and provides a natural methodological benchmark (Kilian and Lewis 2011; Bachmann and Sims 2012; Wong 2015; Epstein et al. 2019; Rostagno et al. 2021). We also consider a more general version of SSA in which the set of shocks that offset the spillovers from a US monetary policy shock in the counterfactual is not restricted (Antolin-Diaz et al. 2021).

Instead, MRE determines the minimum 'tilt' of the posterior of the impulse responses to a US monetary policy shock that satisfies the counterfactual constraint that the mean real activity spillovers are nil. Intuitively, MRE indicates how US variables would evolve in a counterfactual world in which the spillovers from US monetary policy are nil but which is otherwise minimally different from the actual world in an information-theoretic sense (Cogley et al. 2005; Robertson et al. 2005; Giacomini and Ragusa 2014).

The rest of the paper is organized as follows. Section 2.2 provides a description of the Bayesian proxy structural VAR model and lays out our empirical specification. Section 2.3 explains how the counterfactuals are constructed and presents our results. Section 2.4 concludes.

2.2 Empirical framework

This section provides a description of the Bayesian proxy structural VAR (BPSVAR) framework of Arias et al. (2021) and then discusses our model specification and identifying assumptions.

2.2.1 The Bayesian proxy SVAR model

Because we identify a global uncertainty shock in addition to a US monetary policy shock, we lay out the BPSVAR model for the general case with k proxy variables.

Following the notation of Rubio-Ramirez et al. (2010), consider without loss of generality the structural VAR model with one lag and without deterministic terms

$$\boldsymbol{y}_t'\boldsymbol{A}_0 = \boldsymbol{y}_{t-1}'\boldsymbol{A}_1 + \boldsymbol{\epsilon}_t', \qquad \boldsymbol{\epsilon} \sim N(\boldsymbol{0}, \boldsymbol{I}_n), \tag{2.1}$$

where $\boldsymbol{y}_t$ is an $n \times 1$ vector of endogenous variables and $\boldsymbol{\epsilon}_t$ an $n \times 1$ vector of structural shocks. The BPSVAR framework builds on the following assumptions in order to identify k structural shocks of interest: There exists a $k \times 1$ vector of proxy variables $\boldsymbol{m}_t$ that are (i) correlated with the k structural shocks of interest $\boldsymbol{\epsilon}_t^*$ and (ii) orthogonal to the remaining structural shocks $\boldsymbol{\epsilon}_t^o$. Formally, the identifying assumptions are

$$E[\boldsymbol{\epsilon}_t^*\boldsymbol{m}_t'] = \underset{(k\times k)}{\boldsymbol{V}}, \tag{2.2a}$$

$$E[\boldsymbol{\epsilon}_t^o\boldsymbol{m}_t'] = \underset{((n-k)\times k)}{\boldsymbol{0}}, \tag{2.2b}$$

and represent the relevance and the exogeneity condition, respectively.

Denote by $\tilde{\boldsymbol{y}}_t' \equiv (\boldsymbol{y}_t', \boldsymbol{m}_t')$, by $\tilde{\boldsymbol{A}}_\ell$ the corresponding $\tilde{n} \times \tilde{n}$ coefficient matrices with $\tilde{n} = n + k$, by $\tilde{\boldsymbol{\epsilon}} \equiv (\boldsymbol{\epsilon}_t', \boldsymbol{v}_t')' \sim N(\boldsymbol{0}, \boldsymbol{I}_{n+k})$, where $\boldsymbol{v}_t$ is a $k \times 1$ vector of measurement errors (see below). The augmented structural VAR model is then given by

$$\tilde{\boldsymbol{y}}_t'\tilde{\boldsymbol{A}}_0 = \tilde{\boldsymbol{y}}_{t-1}'\tilde{\boldsymbol{A}}_1 + \tilde{\boldsymbol{\epsilon}}_t'. \tag{2.3}$$

To ensure that the augmentation with equations for the proxy variables does not affect the dynamics of the endogenous variables, the coefficient matrices $\tilde{\boldsymbol{A}}_\ell$ are specified as

$$\tilde{\boldsymbol{A}}_\ell = \begin{pmatrix} \underset{(n\times n)}{\boldsymbol{A}_\ell} & \underset{(n\times k)}{\boldsymbol{\Gamma}_{\ell,1}} \\ \underset{(k\times n)}{\boldsymbol{0}} & \underset{(k\times k)}{\boldsymbol{\Gamma}_{\ell,2}} \end{pmatrix}, \qquad \ell = 0, 1. \tag{2.4}$$

The zero restrictions on the lower left-hand side block imply that the proxy variables do not enter the equations of the endogenous variables. The reduced form of the model is

$$\tilde{\boldsymbol{y}}_t' = \tilde{\boldsymbol{y}}_{t-1}'\tilde{\boldsymbol{A}}_1\tilde{\boldsymbol{A}}_0^{-1} + \tilde{\boldsymbol{\epsilon}}_t{}'\tilde{\boldsymbol{A}}_0^{-1}. \tag{2.5}$$

Because the inverse of $\tilde{\boldsymbol{A}}_0$ in Equation (2.4) is given by

$$\tilde{\boldsymbol{A}}_0^{-1} = \begin{pmatrix} \boldsymbol{A}_0^{-1} & -\boldsymbol{A}_0^{-1}\boldsymbol{\Gamma}_{0,1}\boldsymbol{\Gamma}_{0,2}^{-1} \\ 0 & \boldsymbol{\Gamma}_{0,2}^{-1} \end{pmatrix}, \tag{2.6}$$

the last k equations of the reduced form of the VAR model in Equation (2.5) read as

$$\boldsymbol{m}_t' = \tilde{\boldsymbol{y}}_{t-1}'\tilde{\boldsymbol{A}}_1 \begin{pmatrix} -\boldsymbol{A}_0^{-1}\boldsymbol{\Gamma}_{0,1}\boldsymbol{\Gamma}_{0,2}^{-1} \\ \boldsymbol{\Gamma}_{0,2}^{-1} \end{pmatrix} - \boldsymbol{\epsilon}_t'\boldsymbol{A}_0^{-1}\boldsymbol{\Gamma}_{0,1}\boldsymbol{\Gamma}_{0,2}^{-1} + \boldsymbol{v}_t'\boldsymbol{\Gamma}_{0,2}^{-1}, \tag{2.7}$$

which shows that in the BPSVAR framework the proxy variables may be serially correlated and affected by past values of the endogenous variables and measurement error.

Ordering the structural shocks so that $\boldsymbol{\epsilon}_t = (\boldsymbol{\epsilon}_t^{o\prime}, \boldsymbol{\epsilon}_t^{*\prime})'$ yields

$$E\left[\boldsymbol{\epsilon}_t\boldsymbol{m}_t'\right] = -\boldsymbol{A}_0^{-1}\boldsymbol{\Gamma}_{0,1}\boldsymbol{\Gamma}_{0,2}^{-1} = \begin{pmatrix} \underset{((n-k)\times k)}{\boldsymbol{0}} \\ \underset{(k\times k)}{\boldsymbol{V}} \end{pmatrix}. \tag{2.8}$$

The first equality is obtained using Equation (2.7) and because the structural shocks $\boldsymbol{\epsilon}_t$ are by assumption orthogonal to $\boldsymbol{y}_{t-1}$ and $\boldsymbol{v}_t$. The second equality is due to the exogeneity and relevance conditions in Equations (2.2a) and (2.2b). Equation (2.8) shows that the identifying assumptions imply restrictions on the last k columns of the contemporaneous structural impact coefficients in $\tilde{\boldsymbol{A}}_0^{-1}$. In particular, if the exogeneity condition in Equation (2.2b) holds, the first $n-k$ rows of the upper right-hand side sub-matrix $\boldsymbol{A}_0^{-1}\boldsymbol{\Gamma}_{0,1}\boldsymbol{\Gamma}_{0,2}^{-1}$ of $\tilde{\boldsymbol{A}}_0^{-1}$ in Equation (2.6) are zero. From the reduced form in Equation (2.5) it can be seen that this implies that the first $n-k$ structural shocks do not impact contemporaneously the proxy variables. In turn, if the relevance condition in Equation (2.2a) holds, the last k rows of the upper right-hand side sub-matrix $\boldsymbol{A}_0^{-1}\boldsymbol{\Gamma}_{0,1}\boldsymbol{\Gamma}_{0,2}^{-1}$ of $\tilde{\boldsymbol{A}}_0^{-1}$ are different from zero. From the reduced form in Equation (2.5) it can be seen that this implies that the last k structural shocks impact the proxy variables contemporaneously. The Bayesian estimation algorithm of Arias et al. (2021) determines the estimates of $\boldsymbol{A}_0$ and $\boldsymbol{\Gamma}_{0,\ell}$ such that the restrictions on $\tilde{\boldsymbol{A}}_0^{-1}$ implied by Equations (2.2a) and (2.2b) as well as on $\tilde{\boldsymbol{A}}_\ell$ in Equation (2.4) are simultaneously satisfied, and hence the estimation identifies the structural shocks $\boldsymbol{\epsilon}_t^*$.

The BPSVAR framework of Arias et al. (2021) has numerous advantages over the traditional frequentist external instruments structural VAR framework, which are discussed in detail in Appendix B.1. In short: First, the BPSVAR framework allows us to refrain from imposing potentially contentious recursiveness assumptions between the endogenous variables when multiple structural shocks are point-identified—as done below—with multiple proxy variables. Second, the single-step estimation of the BPSVAR model is more efficient and facilitates coherent inference; in fact, the Bayesian set-up allows exact finite sample inference, and does not require an explicit theory to accommodate weak instruments. Third, the BPSVAR framework is relatively flexible in that Equation (2.7) allows the proxy variables to be serially correlated and be affected by measurement error.

2.2.2 VAR model specification

Our point of departure is the US VAR model in Gertler and Karadi (2015) for monthly (log) US industrial production (IP), the (log) US consumer-price index (CPI), the excess bond premium (EBP) and the one-year US Treasury Bill (TB) rate as a monetary policy indicator. We add the VXO, (log) non-US, RoW real industrial production, and the (log) US nominal effective exchange rate (NEER). Variable descriptions and data sources are provided in Table B.1 in the Appendix. The sample spans 2/1990-6/2019.

2.2.3 Identifying assumptions

We identify a US monetary policy shock using a proxy variable and—for the purpose of SSA counterfactuals—two RoW shocks using sign, magnitude and zero restrictions. In addition, we use a second proxy variable to identify a global uncertainty shock to preclude that the RoW shocks are contaminated by common global shocks.[25]

US monetary policy and global uncertainty shocks

For the US monetary policy shock we use the intra-daily interest rate surprises around narrow time windows on FOMC meeting days as proxy variable as in Gertler and Karadi (2015) as well as Caldara and Herbst (2019). The surprises are cleansed from central bank

[25] Ideally we would identify a larger set of global shocks. The reason for identifying only a global uncertainty shock is primarily the scarcity of instruments. However, as our interest does not lie in the effects of a global uncertainty shock *per se*, the more global—such as news—shocks our identification turns out to lump together, the more we rule out that RoW shocks are conflated with other common, global shocks. Below we explore robustness checks in which a global oil supply shock is identified as a second global shock and in which no global shock is identified at all.

information effects using the 'poor-man's' approach of Jarocinski and Karadi (2020): Only interest rate surprises that are accompanied by equity price surprises that have the opposite sign are considered.

For the global uncertainty shock we use the intra-daily gold price surprises of Piffer and Podstawski (2018) on narratively selected days as proxy variable. Piffer and Podstawski (2018) first extend the list of dates selected by Bloom (2009) on which the VXO increased arguably due to exogenous uncertainty shocks. Then they calculate the change in the price of gold between the last auction before and the first auction after the news about the uncertainty event became available to markets.[26]

Define $\boldsymbol{\epsilon}_t^* \equiv (\epsilon_t^{mp}, \epsilon_t^u)'$ and $\boldsymbol{m}_t \equiv (p_t^{\epsilon,mp}, p_t^{\epsilon,u})'$ as the vectors containing the US monetary policy and the global uncertainty shocks and the corresponding proxy variables, respectively. Analogous to Equations (2.2a) and (2.2b), our identifying assumptions are

$$E[\boldsymbol{\epsilon}_t^* \boldsymbol{m}_t'] = \begin{pmatrix} E[p_t^{\epsilon,mp}\epsilon_t^{mp}] & E[p_t^{\epsilon,u}\epsilon_t^{mp}] \\ E[p_t^{\epsilon,mp}\epsilon_t^{u}] & E[p_t^{\epsilon,u}\epsilon_t^{u}] \end{pmatrix} = \boldsymbol{V} \tag{2.9a}$$

$$E[\boldsymbol{\epsilon}_t^o \boldsymbol{m}_t'] = \begin{pmatrix} E[p_t^{\epsilon,mp}\boldsymbol{\epsilon}_t^{o}] & E[p_t^{\epsilon,u}\boldsymbol{\epsilon}_t^{o}] \end{pmatrix} = \mathbf{0}. \tag{2.9b}$$

First, as is standard in the literature (Gertler and Karadi 2015; Caldara and Herbst 2019), in the relevance condition in Equation (2.9a) we assume that US monetary policy shocks drive the interest rate surprises on FOMC meeting days, $E[p_t^{\epsilon,mp}\epsilon_t^{mp}] \neq 0$. The exogeneity condition $E[p_t^{\epsilon,mp}\boldsymbol{\epsilon}_t^{o}] = 0$ in Equation (2.9b) cannot be tested as none of the other structural shocks $\boldsymbol{\epsilon}_t^o$ is observed. However, it seems plausible that in a narrow time window around FOMC meetings monetary policy shocks are the only systematic drivers of interest rate surprises purged from central bank information effects.

Second, in the relevance condition in Equation (2.9a) we assume that global uncertainty shocks drive the gold price surprises on the narratively selected dates, $E[p_t^{\epsilon,u}\epsilon_t^{u}] \neq 0$. Intuitively, as gold is widely seen as a safe haven asset, demand increases when uncertainty rises (Baur and McDermott 2010). Piffer and Podstawski (2018) provide evidence that gold price surprises are relevant instruments for uncertainty shocks based on F-tests and Granger causality tests with the VXO and the macroeconomic uncertainty measure constructed in Jurado et al. (2015). Ludvigson et al. (2021) also use gold price changes as a proxy variable for uncertainty shocks. Regarding the exogeneity condition $E[p_t^{\epsilon,u}\boldsymbol{\epsilon}_t^{o}] = 0$ in Equation (2.9b), Piffer and Podstawski (2018) document that gold price surprises are uncorrelated with a range of non-uncertainty shocks.

When multiple proxy variables are used to identify multiple structural shocks, the relevance and exogeneity conditions as stated in Equations (2.9a) and (2.9b) only achieve set-identification. To achieve sharper point-identification, additional restrictions need to be imposed on the contemporaneous relationships between the endogenous variables $\boldsymbol{y}_t$

[26]The data of Piffer and Podstawski (2018) are available only until 2015; we use the update of Bobasu et al. (2021) until 2019.

reflected by $\boldsymbol{A}_0^{-1}$ or on the relationship between unobserved structural shocks and proxy variables reflected by $\boldsymbol{V}$. The traditional frequentist external instruments VAR approach relies on the former (Mertens and Ravn 2013). Instead, the BPSVAR framework allows us to do the latter, which is less restrictive.

A natural idea is to assume that $\boldsymbol{V}$ is a diagonal matrix. However, these additional restrictions imply over-identification, which cannot be implemented by the algorithm of Arias et al. (2021). We therefore impose a weaker set of additional identifying assumptions on $\boldsymbol{V}$, namely only that US interest rate surprises on FOMC meeting days are not driven by global uncertainty shocks, $E[p_t^{\epsilon,mp}\epsilon_t^u] = 0$.[27] This assumption is routinely made—at least implicitly—in the literature. In fact, if this assumption was not satisfied then the analyses of Gertler and Karadi (2015), Caldara and Herbst (2019) as well as Jarocinski and Karadi (2020) would be invalid, as the identified US monetary policy shocks would be contaminated by global uncertainty shocks.

Rest-of-the-world shocks

Existing literature using SSA has considered offsetting shocks that are intuitively as 'close' as possible to the transmission channel examined (Kilian and Lewis 2011; Bachmann and Sims 2012; Wong 2015; Epstein et al. 2019; Rostagno et al. 2021). To establish a point of contact with this literature, we use RoW shocks to offset the real activity spillovers from a US monetary policy shock. In particular, we 'identify' 'RoW depreciating' and 'RoW appreciating' shocks. For example, 'RoW depreciating' shocks include contractionary demand and 'RoW appreciating' shocks contractionary monetary policy shocks. While our identification implies these shocks only have a reduced-form interpretation, this is sufficient for our purposes as together they span the universe of RoW structural shocks.

Table 2.1 presents the sign and relative magnitude restrictions used to identify the two RoW shocks. Both shocks are normalised so that RoW real activity slows down on impact. We assume real activity is impacted more at home than abroad in order to distinguish US from RoW shocks. For the 'RoW depreciating' shock we assume that it appreciates the US NEER and that it slows down RoW and US real activity. We assume US real activity slows down as expenditure reducing and expenditure switching effects in the US point in the same direction. For the 'RoW appreciating' shock we assume that it depreciates the US NEER. We do not assume that US real activity slows down, as expenditure reducing and expenditure switching effects move in opposite directions.[28]

While using shocks that are as 'close' as possible to the transmission channel of interest for the offsetting may be appealing intuitively and a useful point of contact with existing literature, it is not compelling conceptually. Below we therefore also consider a more general

[27] When two proxy variables are used to identify two structural shocks one additional restriction on $\boldsymbol{V}$ is sufficient for point-identification (Giacomini et al. 2022).

[28] While we do not take a stand on the response of the exchange rate to RoW productivity, financial and fiscal policy shocks, it is clear that they are captured by one of the two RoW reduced-form shocks.

Table 2.1: Identification restrictions of the RoW shocks

Variable / Shock	RoW 'depreciating' shock	RoW 'appreciating' shock
US 1-year T-Bill rate		
US industrial production	$< 0^{\triangle}$	$\Diamond$
US CPI		
US excess bond premium		
US dollar NEER	> 0	< 0
VXO		
RoW industrial production	< 0 & $<^{\triangle}$	< 0 & $<^{\Diamond}$

Notes: The table presents the sign and magnitude restrictions we impose in order to identify the RoW shocks. We additionally impose the exogeneity restrictions $E[p_t^{\epsilon,mp}\boldsymbol{\epsilon}_t^o] = 0$ in Equation (2.9b) that the proxy variables are not driven by the RoW shocks.

version of SSA in which all structural shocks are used for the offsetting.

2.2.4 Priors

We use flat priors for the VAR parameters. We follow Caldara and Herbst (2019) and Arias et al. (2021) and impose a relevance threshold to express a prior belief that the proxy variables are relevant instruments: We require that at least a share $\gamma = 0.1$ of the proxy variable variance is accounted for by US monetary policy and global uncertainty shocks; this is weaker than the relevance threshold of $\gamma = 0.2$ in Arias et al. (2021) and—although not directly comparable—the 'high-relevance prior' of Caldara and Herbst (2019).

2.2.5 Baseline impulse responses

The first column in Figure 2.1 shows that consistent with the literature a contractionary US monetary policy shock is estimated to raise the one-year Treasury Bill rate, the EBP and the VXO, appreciates the US NEER, temporarily reduces US industrial production and persistently US consumer prices (Gertler and Karadi 2015; Caldara and Herbst 2019). Consistent with the literature on the spillovers from US monetary policy RoW real activity slows down considerably (see, e.g., Banerjee et al. 2016; Georgiadis 2016; Dedola et al. 2017; Iacoviello and Navarro 2019; Vicondoa 2019; Degasperi et al. 2020).

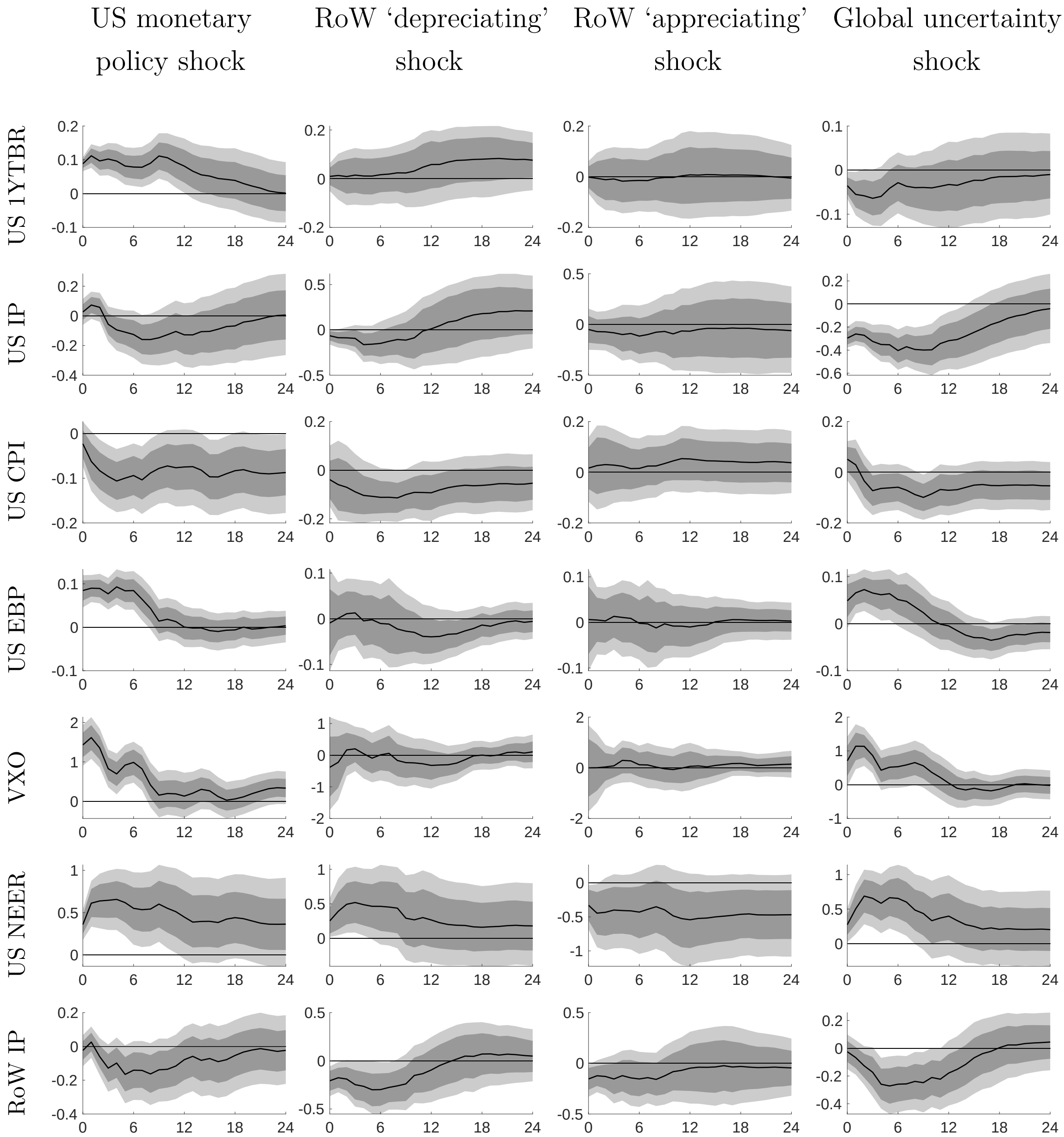
The second and third columns show the effects of the RoW shocks. As each of them lumps together different structural RoW shocks, interpreting their impulse responses is not straightforward.

The last column shows that consistent with the literature on the role of the dollar and US assets in the world economy the global uncertainty shock appreciates the US NEER, raises the VXO and the EBP, causes a slowdown in global real activity, lowers US consumer prices and the one-year Treasury Bill rate (Epstein et al. 2019; Jiang et al. 2021).[29]

[29] Figure B.1 in the Appendix documents that our baseline results are robust to several alternative specifications: (i) we drop interest rate surprises on non-scheduled FOMC 'intermeetings' as suggested

by Caldara and Herbst (2019); (ii) instead of the approach of Gertler and Karadi (2015) for temporal aggregation of the interest rate and gold price surprises from daily to monthly frequency we take the simple average in a given month as in Jarocinski and Karadi (2020); (iii) we follow Miranda-Agrippino and Ricco (2021) and purge the surprises from information contained in Greenbook projections instead of the 'poor-man's' approach of Jarocinski and Karadi (2020); (iv) we additionally include RoW consumer prices and policy rates, US exports and imports as well as global equity prices in the VAR model; (v) we additionally identify US demand and supply shocks using sign restrictions; (vi) we additionally identify an oil supply shock as a second global shock that is common to the US and the RoW using the proxy variable constructed by Känzig (2021); (vii) we do not impose a 'relevance threshold' for the proxy variables.

Figure 2.1: Baseline impulse responses to US monetary policy, RoW ‘appreciating’ and ‘depreciating’, and global uncertainty shocks



Notes: The figure shows the point-wise posterior means of the impulse responses (black solid lines) together with 68% and 90% centered point-wise probability bands (grey shaded areas) obtained from the BPSVAR model. ‘1YTBR’ stands for the one-year Treasury Bill rate, ‘IP’ for industrial production, ‘CPI’ for the consumer-price index, ‘EBP’ for excess bond premium, ‘VXO’ for the S&P 500 stock market volatility index, and ‘NEER’ for the nominal effective exchange rate.

2.3 Spillbacks from US monetary policy

The VAR model in Equation (2.1) can be iterated forward and re-written as

$$\boldsymbol{y}_{T+1,T+h} = \boldsymbol{b}_{T+1,T+h} + \boldsymbol{M}'\boldsymbol{\epsilon}_{T+1,T+h}, \tag{2.10}$$

where the $nh \times 1$ vector $\boldsymbol{y}_{T+1,T+h} \equiv [\boldsymbol{y}'_{T+1}, \boldsymbol{y}'_{T+2}, \ldots, \boldsymbol{y}'_{T+h}]'$ denotes future values of the endogenous variables, $\boldsymbol{b}_{T+1,T+h}$ an autoregressive component that is due to initial conditions as of period T, and the $nh \times 1$ vector $\boldsymbol{\epsilon}_{T+1,T+h} \equiv [\boldsymbol{\epsilon}'_{T+1}, \boldsymbol{\epsilon}'_{T+2}, \ldots, \boldsymbol{\epsilon}'_{T+h}]'$ future values of the structural shocks. The $nh \times nh$ matrix $\boldsymbol{M}$ reflects the impulse responses and is a function of the structural VAR parameters $\boldsymbol{\psi} \equiv vec(\boldsymbol{A}_0, \boldsymbol{A}_1)$.

Assume for simplicity of exposition but without loss of generality that the VAR model in Equation (2.1)—which does not have deterministic components—is stationary and in steady state in period T so that $\boldsymbol{b}_{T+1,T+h} = \boldsymbol{0}$. In this setting, an impulse response to the i-th structural shock over a horizon of h periods coincides with the forecast $\boldsymbol{y}_{T+1,T+h}$ conditional on $\boldsymbol{\epsilon}_{T+1,T+h} = [\boldsymbol{e}'_i, \boldsymbol{0}_{1\times n(h-1)}]'$, where $\boldsymbol{e}_i$ is an $n \times 1$ vector of zeros with unity at the i-th position. For example, for the impulse response to a US monetary policy shock we have $\epsilon^{mp}_{T+1} = 1$, $\epsilon^{mp}_{T+s} = 0$ for $s > 1$ and $\epsilon^{\ell}_{T+s} = 0$ for $s > 0$, $\ell \neq mp$.

We define spillbacks as the difference between the impulse responses of domestic variables to a US monetary policy shock in the baseline denoted by $\boldsymbol{y}_{T+1,T+h}$ and in a counterfactual denoted by $\widetilde{\boldsymbol{y}}_{T+1,T+h}$. The defining feature of the counterfactual is that the spillovers from a US monetary policy shock to RoW real activity are nil. We consider two approaches for constructing the counterfactual impulse response $\widetilde{\boldsymbol{y}}_{T+1,T+h}$: SSA and MRE.

2.3.1 SSA counterfactuals: Conceptual considerations

In SSA the VAR model is unchanged in the counterfactual in terms of the structural parameters $\boldsymbol{\psi}$ and hence $\boldsymbol{M}$ in Equation (2.10). Therefore, in order for the impulse response $\widetilde{\boldsymbol{y}}_{T+1,T+h}$ to satisfy counterfactual constraints additional shocks in $\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h}$ must be allowed to materialise over horizons $T+1, T+2, \ldots, T+h$. In our application of SSA the values of these additional shocks are chosen such that their effect is to offset the spillovers from a US monetary policy shock to RoW real activity.

Building on Waggoner and Zha (1999), Antolin-Diaz et al. (2021) show how to obtain $\widetilde{\boldsymbol{y}}_{T+1,T+h}$ subject to constraints on the paths of a subset of the endogenous variables

$$\overline{\boldsymbol{C}}\widetilde{\boldsymbol{y}}_{T+1,T+h} = \overline{\boldsymbol{C}}\boldsymbol{M}'\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h} \sim N(\overline{\boldsymbol{f}}_{T+1,T+h}, \overline{\boldsymbol{\Omega}}_f), \tag{2.11}$$

where $\overline{\boldsymbol{C}}$ is a $k_o \times nh$ selection matrix, $\overline{\boldsymbol{f}}_{T+1,T+h}$ is a $k_o \times 1$ vector and $\overline{\boldsymbol{\Omega}}_f$ a $k_o \times k_o$ matrix, and subject to constraints on the structural shocks given by

$$\Xi\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h} \sim N(\boldsymbol{g}_{T+1,T+h}, \boldsymbol{\Omega}_g), \tag{2.12}$$

where $\boldsymbol{\Xi}$ is a $k_s \times nh$ selection matrix, $\boldsymbol{g}_{T+1,T+h}$ a $k_s \times 1$ vector, and $\boldsymbol{\Omega}_g$ a $k_s \times k_s$ matrix. In our context, Equation (2.11) imposes the counterfactual constraint that the real activity spillovers from US monetary are nil, and Equation (2.12) the constraint that some structural shocks may not be in the set of offsetting shocks that materialise along the impulse response horizon to enforce the counterfactual constraint. Antolin-Diaz et al. (2021) show how to obtain the SSA solution in terms of $\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h}$ which satisfies the counterfactual constraints in Equations (2.11) and (2.12). The SSA counterfactual impulse response is then given by $\widetilde{\boldsymbol{y}}_{T+1,T+h} = \boldsymbol{M}'\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h}$.[30,31]

2.3.2 SSA counterfactuals: Results

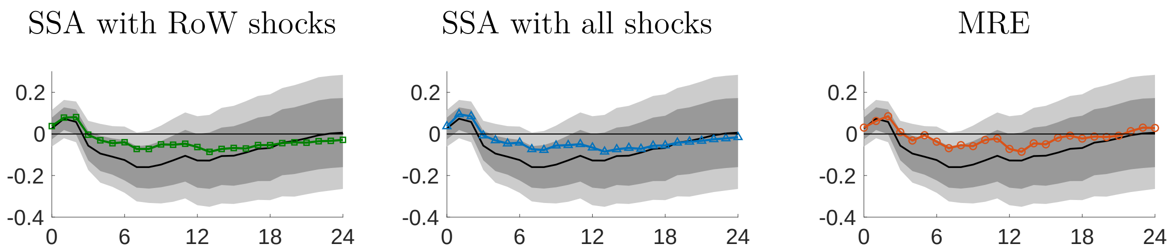
The left-hand side panel in Figure 2.2 presents the baseline impulse response of domestic industrial production to the US monetary policy shock from Figure 2.1 (black solid line) and the SSA counterfactual in which RoW shocks materialise so that real activity spillovers are offset (green line with squares). In the counterfactual the drop in US industrial production is reduced substantially compared to the baseline. This implies that spillbacks amplify the domestic effects of US monetary policy in the data. Quantitatively, spillbacks account for almost 50% of the overall domestic effect of US monetary policy on industrial production.[32]

[30] Because at every horizon two RoW shocks can be used to impose the constraint that RoW real activity does not respond to a US monetary policy shock, there is a multiplicity of solutions. Antolin-Diaz et al. (2021) show that the SSA solution minimises the Frobenius norm of the deviation of the distribution of the structural shocks under the counterfactual from the baseline. Intuitively, this means the counterfactual shocks chosen are those that are minimally different in terms of mean and variance from the baseline.

[31] See Appendix B.4 for further technical details and the specification of $\overline{\boldsymbol{C}}$, $\overline{\boldsymbol{f}}_{T+1,T+h}$, $\boldsymbol{\Xi}$, $\overline{\boldsymbol{g}}_{T+1,T+h}$, $\boldsymbol{\Omega}_g$ and $\overline{\boldsymbol{\Omega}}_f$ in the baseline and the counterfactual conditional forecast in our application.

[32] The Appendix documents that our results are similar for several alternative specifications of the 'SSA RoW shocks' counterfactual. First, Figure B.2 presents results for specifications in which we identify RoW shocks assuming they have no effect on the US contemporaneously and imposing the sign restrictions for the spillovers from Table 2.1 only with a one-month lag. Second, Figure B.3 presents results for specifications in which we (i) identify a global oil supply shock as a second global shock, and in which we (ii) do not identify any global shock. Third, Figure B.4 documents that spillovers to other RoW variables such as monetary policy, inflation, spreads and trade are mostly close to zero in the counterfactual even if they are not constrained to be so, and Figure B.5 shows that results for spillbacks are very similar if the responses of these variables are constrained to be zero in the counterfactual.

Figure 2.2: Baseline and counterfactual impulse responses of US industrial production to a US monetary policy shock



Notes: The black solid lines depict the baseline impulse responses of US industrial production to a US monetary policy shock and the coloured solid lines with symbol markers the counterfactual impulse responses based on point-wise posterior mean SSA with RoW shocks (left column, green lines with squares), based on point-wise posterior mean SSA with all shocks (middle column, blue lines with triangles), and based on point-wise posterior mean MRE (right column, orange lines with circles). The grey shaded areas represent 68% and 90% centered point-wise probability bands for the baseline impulse responses.

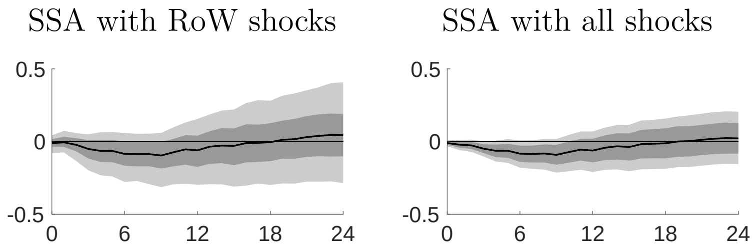
Considering RoW shocks to offset real activity spillovers from US monetary policy shocks is appealing intuitively and in line with earlier literature (Kilian and Lewis 2011; Bachmann and Sims 2012; Wong 2015; Epstein et al. 2019; Rostagno et al. 2021). At the same time, it is not compelling conceptually that there should be any constraint on the set of offsetting shocks. Therefore, the middle panel in Figure 2.2 presents results for a more general SSA specification in which all shocks in the VAR are used for the offsetting. The results are very similar to those based only on RoW shocks.[33]

Figure 2.3 presents the posterior distribution of the difference between the response of domestic industrial production to a US monetary policy shock in the baseline and the SSA counterfactuals from Figure 2.2. Two observations stand out. First, the estimation assigns a high posterior probability to spillbacks being contractionary. Second, spillbacks are estimated more precisely if the set of offsetting shocks is not constrained. This is because in this case only the US monetary policy shock needs to be identified, and hence there is no uncertainty due to the sign and relative magnitude restrictions that only set-identify the two RoW shocks.[34]

[33]Figure B.6 in the Appendix documents that our results are similar if for the offsetting we use (i) all shocks except the US monetary policy shock, or (ii) all shocks except US monetary policy and global uncertainty shocks.

[34]Results for the 'modesty statistic' of Leeper and Zha (2003) in Figure B.7 in the Appendix indicate that the offsetting shocks are not unusually large or persistent. Similarly, results for the q-divergence proposed by Antolin-Diaz et al. (2021) indicate that the distribution of shocks in the counterfactual is not very different from the baseline. Therefore, our SSA counterfactuals are unlikely to be subject to the Lucas critique.

Figure 2.3: Distribution of SSA spillback estimates



Notes: The figure presents the point-wise posterior mean of the differences between the baseline and the counterfactual effects of US monetary policy on US industrial production together with 68% and 90% centered point-wise probability bands.

2.3.3 MRE counterfactuals: Conceptual considerations

In the existing literature MRE is used to incorporate restrictions derived from economic theory in order to improve a forecast. For example, Robertson et al. (2005) improve their forecasts of the Federal Funds rate, US inflation and the output gap by imposing the constraint that the mean three-year-ahead inflation forecast must equal 2.5% through MRE.[35] Similar in spirit, we use MRE to generate a counterfactual conditional forecast based on our baseline conditional forecast in Equation (2.10) that represents the impulse responses to a US monetary policy shock.

Again conceive of an impulse response as the conditional forecast $\boldsymbol{y}_{T+1,T+h}$, where we have for $\boldsymbol{\epsilon}_{T+1,T+h}$ that $\epsilon^{mp}_{T+1} = 1$, $\epsilon^{mp}_{T+s} = 0$ for $s > 1$ and $\epsilon^{\ell}_{T+s} = 0$ for $s > 0$, $\ell \neq mp$. Our posterior belief about the actual effects of a US monetary policy shock after h periods is given by

$$f(\boldsymbol{y}_{T+h}|\boldsymbol{y}_{1,T}, \mathcal{I}_a, \boldsymbol{\epsilon}_{T+1,T+h}) \propto p(\boldsymbol{\psi}) \times \ell(\boldsymbol{y}_{1,T}|\boldsymbol{\psi}, \mathcal{I}_a) \times \nu, \tag{2.13}$$

where $p(\boldsymbol{\psi})$ is the prior about the structural VAR parameters, $\mathcal{I}_a$ our identifying assumptions, and ν the volume element of the mapping from the posterior distribution of the structural VAR parameters to the posterior distribution of the impulse response $\boldsymbol{y}_{T+h}$. MRE determines the posterior beliefs about the effects of a US monetary policy shock $\widetilde{\boldsymbol{y}}_{T+h}$ in a counterfactual VAR model with structural parameters $\widetilde{\boldsymbol{\psi}}$ by

$$\begin{gathered} Min_{\widetilde{\boldsymbol{\psi}}} \ \mathcal{D}(f^*||f) \quad s.t. \\ \int f^*(\widetilde{\boldsymbol{y}})\widetilde{y}^{ip^*} d\widetilde{\boldsymbol{y}} = E(\widetilde{y}^{ip^*}) = 0, \ \int f^*(\widetilde{\boldsymbol{y}}) d\widetilde{\boldsymbol{y}} = 1, \ \ f^*(\widetilde{\boldsymbol{y}}) \geq 0, \end{gathered} \tag{2.14}$$

where $\mathcal{D}(\cdot)$ is the Kullback-Leibler divergence—the 'relative entropy'—between the counterfactual and baseline posterior beliefs (the subscripts in Equation (2.14) are dropped for

[35]See Cogley et al. (2005) and Giacomini and Ragusa (2014) for similar applications.

simplicity). In general, there are infinitely many counterfactual beliefs f^* that satisfy the constraint $E(\widetilde{y}^{ip^*}_{T+h}) = 0$. The MRE approach disciplines the choice of the counterfactual posterior beliefs f^* by requiring that they are *minimally* different from the baseline posterior beliefs f in an information-theoretic sense.[36] Intuitively, MRE determines the counterfactual VAR model in which real activity spillovers from US monetary policy are nil but whose dynamic properties in terms of impulse responses are otherwise minimally different from those of the actual VAR model.[37,38]

2.3.4 MRE counterfactuals: Results

The right-hand side column in Figure 2.2 presents the baseline impulse response of US industrial production to the contractionary US monetary policy shock (black solid line) together with the MRE counterfactual (orange line with circles). The results are similar to those from SSA.[39,40]

Appendix B.6 reports placebo tests for our SSA/MRE counterfactuals. In particular, we estimate spillbacks from US monetary policy through some small open economy (SOE) rather than through the entire RoW. As one would plausibly expect, the counterfactual in which only spillovers to the SOE but not the entire RoW are precluded turns out to be hardly different from the baseline. Similarly, when spillovers to the RoW are precluded but spillovers to the SOE are allowed for, the results hardly differ from the counterfactuals in Figure 2.2.

2.3.5 Spillback transmission channels

To shed light on the channels through which spillbacks from US monetary policy materialise we first examine the responses of US GDP components.

[36]The counterfactual structural VAR parameters $\widetilde{\boldsymbol{\psi}}$ could in principle be obtained from the counterfactual impulse responses $\widetilde{\boldsymbol{y}}$ based on the mapping between impulse responses and structural VAR parameters (see Arias et al. 2018, Appendix B.2).

[37]A brute force alternative for counterfactual analysis in VAR models is to set to zero autoregressive parameters after or before estimation (see, for example, Vicondoa 2019; Degasperi et al. 2020). However, setting to zero VAR coefficients before estimation implies model mis-specification and results in biased estimates; and in general the bias is not informative about the strength of the channel that is being shut down (Georgiadis 2017). In turn, setting to zero VAR coefficients after estimation may be understood similarly as the MRE approach in the sense that it reflects some counterfactual VAR model. However, while the MRE approach determines a counterfactual VAR model that is—roughly speaking—minimally different from the actual VAR model, setting to zero some VAR coefficient does not impose any intuitively appealing discipline on the choice of the counterfactual VAR model (for a discussion see Benati 2010).

[38]See Appendix B.5 for technical details on the implementation of MRE.

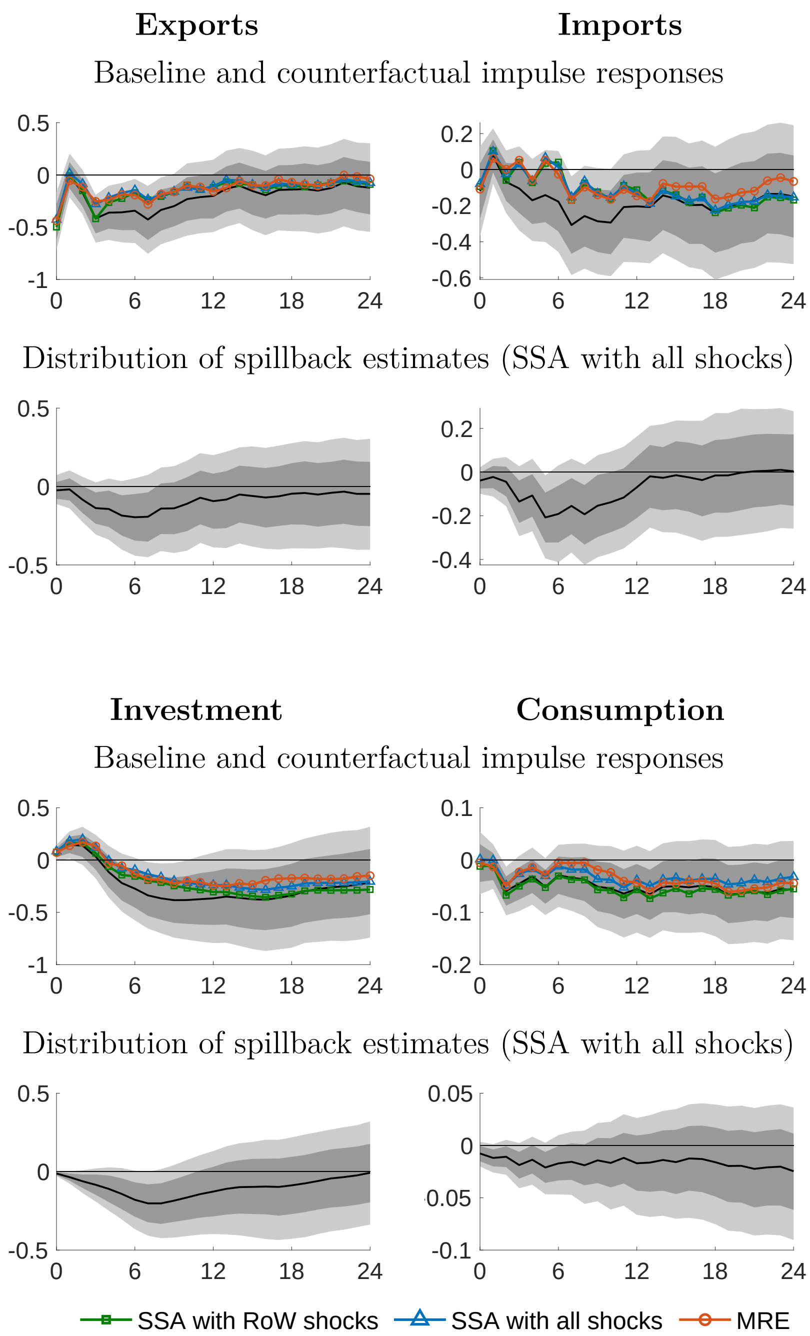
[39]While SSA produces a *distribution of differences*, MRE produces a *different distribution*. Therefore, we cannot report an MRE analogue to Figure 2.3.

[40]When we test H_0 : *Spillbacks are zero* based on the Bayes factor as suggested by Kass and Raftery (1995) for the marginal likelihoods of the baseline and the MRE counterfactual obtained from the harmonic mean estimator of Geweke (1999), the evidence against H_0 : is—in the words of Kass and Raftery (1995)—"very strong" and "decisive". We are grateful to a referee for suggesting this exercise.

GDP components

Figure 2.4 displays the responses of US real exports, imports, investment and consumption to a monetary policy shock in the baseline and the counterfactual in which real activity spillovers from US monetary policy are precluded. All GDP components decline in response to a contractionary US monetary policy shock in the baseline. In the counterfactual the decline is weaker for all GDP components. The results in the first two rows suggest that net trade cannot account for spillbacks to the US: Exports and imports decline by less in the counterfactual to roughly the same degree. The panels in the last two rows suggest that spillbacks instead arise through consumption and in particular investment. The underlying mechanisms are explored in more detail next.

Figure 2.4: Responses of monthly GDP components to US monetary policy shock for the baseline and the counterfactual



Notes: In the first and third row the figure shows the baseline (black solid lines) and counterfactual impulse responses based on SSA with RoW shocks (green lines with squares), SSA with all shocks (blue lines with triangles) and MRE (orange lines with circles) for real exports, real imports, real private gross fixed capital investment and real private consumption expenditures to a US monetary policy shock. The grey shaded areas represent 68% and 90% centered point-wise probability bands for the baseline impulse responses. In the second and fourth row the figure shows the point-wise posterior mean of the differences between the baseline and the 'All shocks SSA' counterfactual together with 68% and 90% centered point-wise probability bands. We augment the VAR model by one additional endogenous variable at a time.

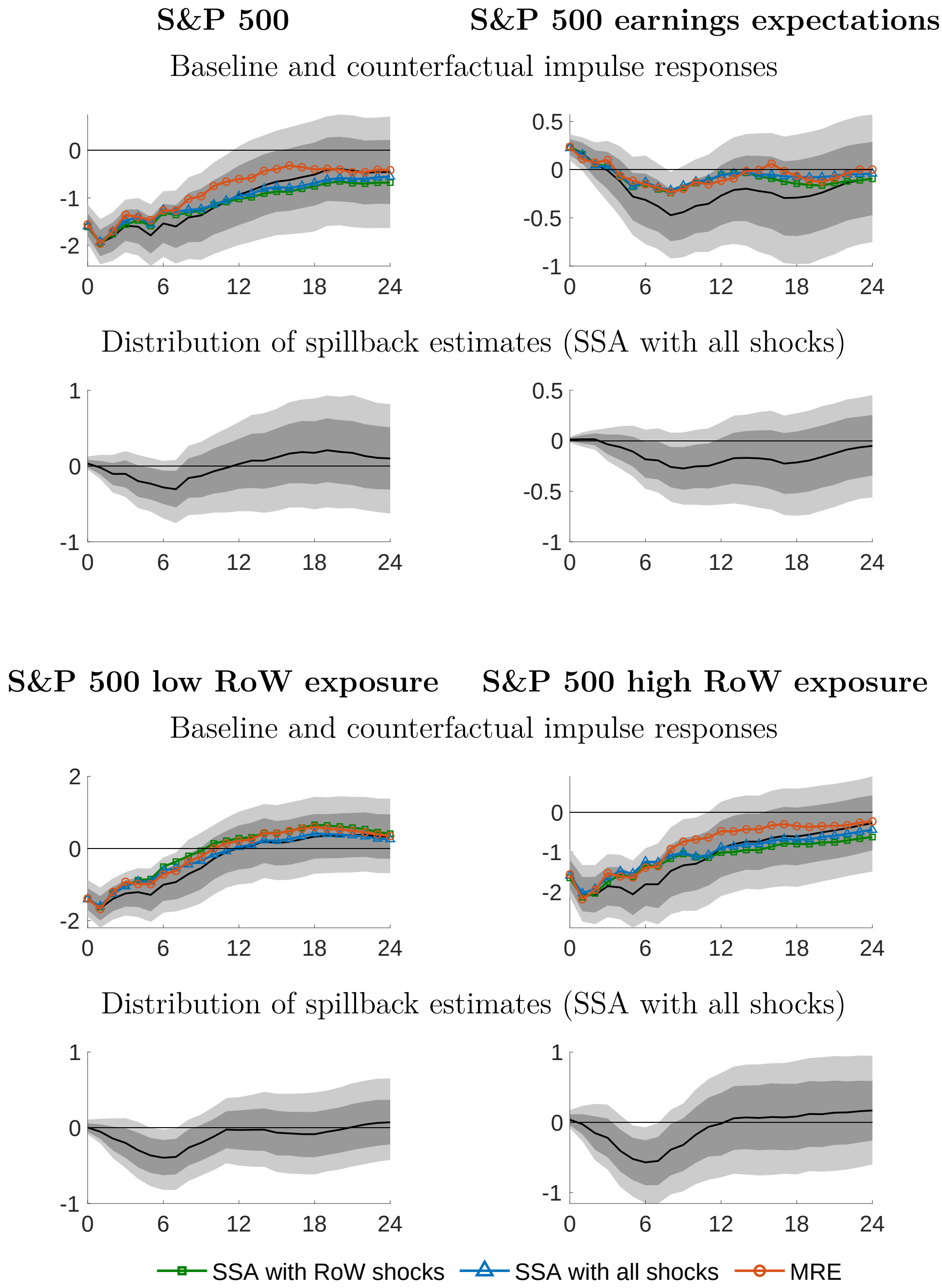
Investment

In theory a key determinant of investment is Tobin's q, the ratio of the market price of capital—a firm's stock market valuation—and its replacement price. A slowdown in RoW real activity in response to a contractionary US monetary policy shock might reduce US firms' foreign sales, their profits, and hence their valuations. Moreover, much work has documented that firm investment is additionally determined by cash flow (Cao et al. 2019; Lian and Ma 2021; Drechsel 2023).

Indeed, Figure 2.5 documents that US equity prices fall by less in the counterfactual when spillovers from US monetary policy are precluded. Similarly, and somewhat more clearly, US firms' cash flows measured by earnings expectations fall by less in the counterfactual. That valuations and cash flows fall by less in the counterfactual in which spillovers are precluded is consistent with US firms' substantial exposure to the RoW: More than 40% (30%) of total sales (revenues) of S&P 500 firms are accounted for by the RoW (Brzenk 2018; Silverblatt 2019). In fact, Figure 2.5 documents that the responses of valuations exhibit greater differences across the baseline and the counterfactual for sectors which are more exposed to the RoW.

Overall, our results are consistent with spillbacks from US monetary policy arising through cutbacks in investment by US firms whose valuations and cash flows fall as they experience a decline in foreign demand. Figure B.9 in the Appendix documents that other possible channels—through probabilities of default, risk premia and uncertainty—do not seem to contribute to spillbacks to investment.

Figure 2.5: Channels of transmission for spillbacks from US monetary policy to investment

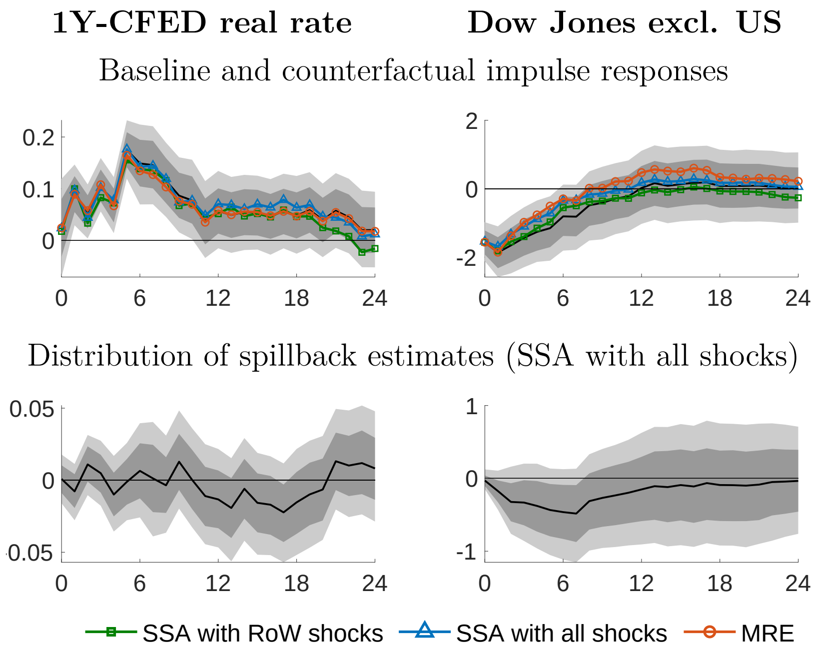


Notes: In the first and third row the figure shows the baseline (black solid lines) and counterfactual impulse responses based on SSA with RoW shocks (green lines with squares), SSA with all shocks (blue lines with triangles) and MRE (orange lines with circles) for the S&P 500 Composite, the 12-months forward S&P 500's earnings per share, and the S&P 500 index for low and high RoW exposures. The latter two are constructed as market-capitalisation-weighted averages of sectoral S&P indices. The low RoW exposure sectors are utilities, telecommunication services, health care and financials, and the high RoW exposure sectors are energy, materials, industrials and information technology; see Brzenk (2018) for data and a discussion of RoW exposures in the S&P 500. The grey shaded areas represent 68% and 90% centered point-wise probability bands for the baseline impulse responses. In the second and fourth row the figure shows the point-wise posterior mean of the differences between the baseline and the 'All shocks SSA' counterfactual together with 68% and 90% centered point-wise probability bands. We augment the VAR model by one additional endogenous variable at a time.

Consumption

The main channel through which monetary policy affects consumption in the standard representative-agent New Keynesian model centers on interest rates and inter-temporal substitution. However, Figure 2.6 documents that the one-year ahead *ex ante* real interest rate responds very similarly in the baseline and the counterfactual. The evidence thus suggests inter-temporal substitution does not play a role in the transmission of spillbacks from US monetary policy.

Figure 2.6: Channels of transmission for spillbacks from US monetary policy to consumption



Notes: In the first row the figure shows the baseline (black solid lines) and counterfactual impulse responses based on SSA with RoW shocks (green lines with squares), SSA with all shocks (blue lines with triangles) and MRE (orange lines with circles) for the Cleveland Fed interest rate-term structure-based one-year real rate and the Dow Jones World excl. US index. The grey shaded areas represent 68% and 90% centered point-wise probability bands for the baseline impulse responses. In the second row the figure shows the point-wise posterior mean of the differences between the baseline and the 'All shocks SSA' counterfactual together with 68% and 90% centered point-wise probability bands. We augment the VAR model by one additional endogenous variable at a time.

Recent research highlights that indirect channels may be quantitatively more important for monetary transmission than direct channels centred on inter-temporal substitution. Kaplan et al. (2018) propose a heterogeneous-agent NK framework in which the effect of monetary policy on consumption that materialises through indirect channels involving labour demand, wages and wealth is large relative to direct channels. Alves et al. (2020) generalise the model of Kaplan et al. (2018) by accounting for aggregate capital adjustment costs and inertia in the monetary policy rule and show that this entails substantial marginal propensities to consume out of illiquid equity wealth. Alves et al. (2020) show that these wealth effects play a quantitatively important role in the transmission of monetary policy.

Recall that Figure 2.5 already documents that US equity prices fall by less in the counterfactual, and that this is plausibly related to spillovers and hence spillbacks due to a weaker fall in US firms' drop in foreign sales. Analogously, Figure 2.6 documents that RoW equity prices also fall less in the counterfactual. Overall, the responses of global equity prices are qualitatively consistent with stock market wealth effects accounting for the spillbacks to US consumption. However, as domestic holdings are more important in US household portfolios, spillbacks through stock market wealth effects on RoW equity arguably only play a lesser role.[41]

Figure B.9 in the Appendix again documents that other channels—through precautionary savings related to variation in consumer confidence and wealth effects through house price variation—do not appear to account for spillbacks to consumption.

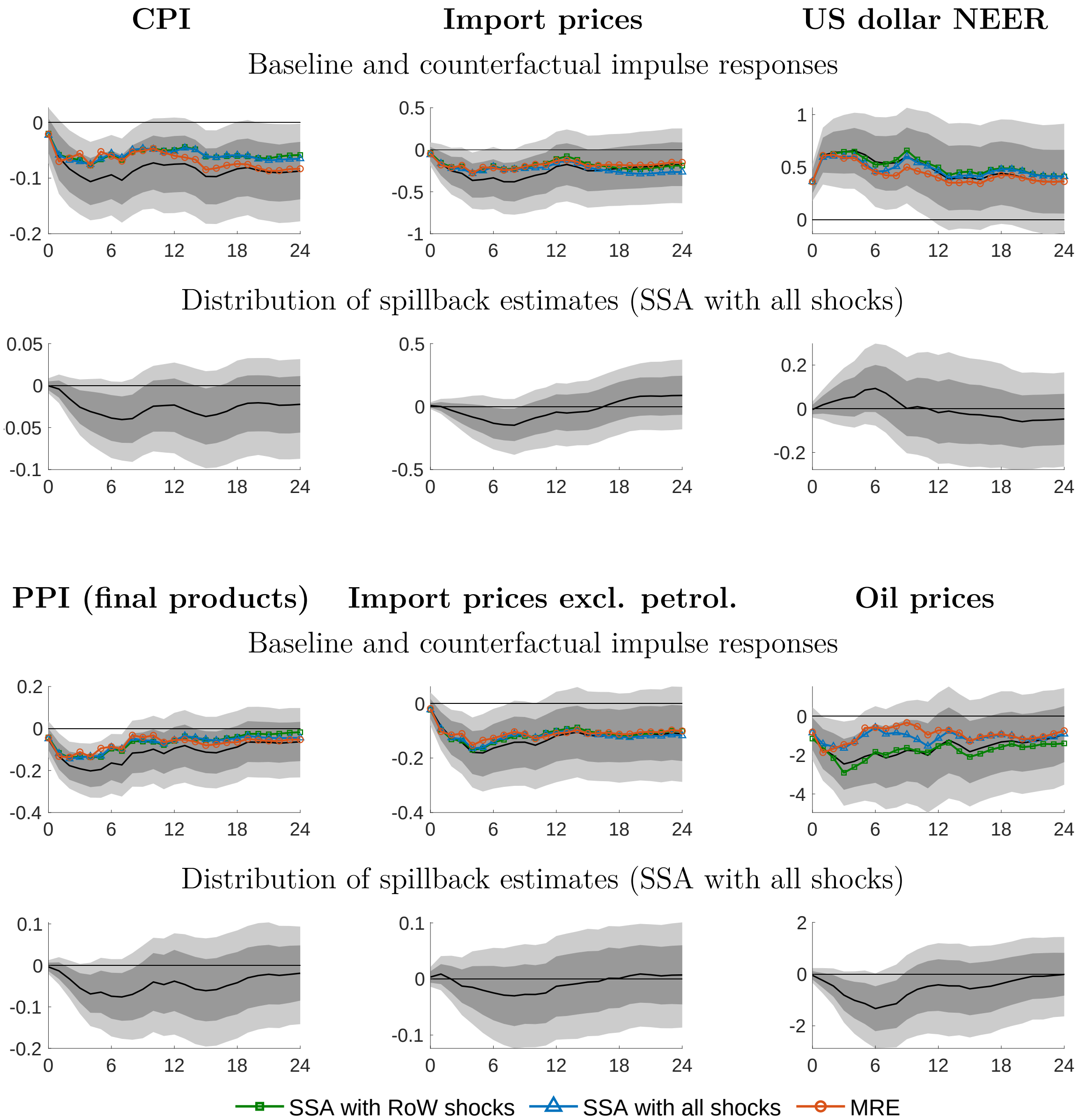
Finally, note two remarks on our findings. First, our analysis of the transmission channels of spillbacks is reduced form and thereby suggestive. In particular, it does not allow us to decompose the overall spillbacks into contributions due to Tobin's q/cash flow and stock market wealth effects. In practice, these channels interact and amplify each other in general equilibrium. Second, one should not interpret the findings in Figure 2.2 as suggesting that US monetary policy would be ineffective domestically if there were no spillbacks. In particular, note that in the counterfactual world in which spillbacks through US firms' foreign sales are nil it would necessarily also be that US firms are not exposed to the RoW. But then in this counterfactual US firms' exposure to the US economy would be commensurately higher, and the corresponding cash flow and stock market wealth effects would play out through domestic rather than foreign sales.

2.3.6 Spillbacks to US consumer prices

The first panel in Figure 2.7 indicates that spillbacks account for almost 50% of the total effect on US consumer prices. The remaining panels in Figure 2.7 suggest that spillbacks to consumer prices materialise through both domestic and import prices. Given that US imports are largely invoiced in dollar (Gopinath et al. 2010), the weaker appreciation of the US NEER in the counterfactual likely only plays a limited role. Instead, the weaker drop in US import prices is primarily due to a weaker fall in oil prices in the counterfactual. Indeed, the reduction in the drop in import prices excluding petroleum in the counterfactual is much smaller than for overall import prices.

[41] Figure B.8 in the Appendix shows that domestic equity accounts for roughly two thirds of total US equity holdings. When we constrain RoW equity prices in the counterfactual to fall by as much as in the baseline, the results hardly differ from the counterfactuals in Figure 2.2.

Figure 2.7: Spillbacks for US consumer prices



Notes: In the first and third row the figure shows the baseline (black solid lines) and counterfactual impulse responses based on SSA with RoW shocks (green lines with squares), SSA with all shocks (blue lines with triangles) and MRE (orange lines with circles) for US CPI, import prices with and without petroleum, PPI, the US dollar NEER and oil prices to a US monetary policy shock. The grey shaded areas represent 68% and 90% centered point-wise probability bands for the baseline impulse responses. In the second and fourth row the figure shows the point-wise posterior mean of the differences between the baseline and the'All shocks SSA' counterfactual together with 68% and 90% centered point-wise probability bands. We augment the VAR model by one additional endogenous variable at a time.

2.3.7 Spillbacks through AEs vs. EMEs

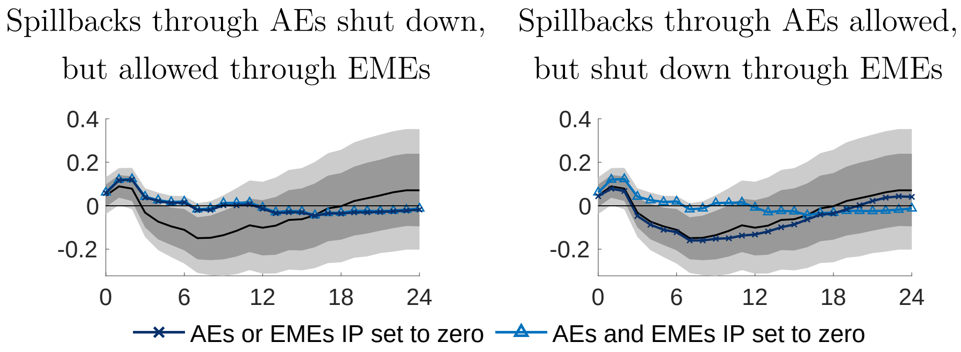
Finally, we explore if spillbacks materialise through AEs or EMEs or both. To do so, we re-estimate the VAR model replacing RoW industrial production by AE and EME analogues.

We then repeat the counterfactual analysis, imposing that real activity spillovers to AEs and EMEs are nil. In order to assess the role of spillovers to AEs and EMEs for the overall spillbacks, we consider two variations of the counterfactual: First we only preclude spillovers from US monetary policy to AEs while constraining spillovers to EMEs to coincide with those in the baseline; hence, in this variation we shut down spillbacks through AEs but allow spillbacks through EMEs. Second, we do the reverse.

The left-hand side panel in Figure 2.8 presents the results for the counterfactual in which we shut down spillovers to AEs but not to EMEs. The light blue line with triangles depicts the domestic real activity response to a US monetary policy shock when spillovers to the entire RoW—i.e. both AEs and EMEs—are precluded, and the dark blue line with crosses when only spillovers to AEs are precluded. The domestic effect of a US monetary policy shock is almost identical when spillovers to the entire RoW or only to AEs are precluded. This implies that spillbacks from US monetary policy arise much more through AEs rather than EMEs. Indeed, the right-hand side panel shows that the domestic effect of a US monetary policy shock is almost identical when spillovers to the entire RoW are unconstrained (black solid line) and when only spillovers to EMEs are precluded (dark blue line with crosses).[42] An important caveat to this finding is that it is based on the entire sample period from 1990 to 2019. As the importance of EMEs has been growing over time, their role for spillbacks may be larger nowadays.

[42]Figure B.11 in the Appendix documents a more important role for AEs than EMEs in US firms' foreign sales (proxied by the country composition of exports) and the US foreign equity portfolio.

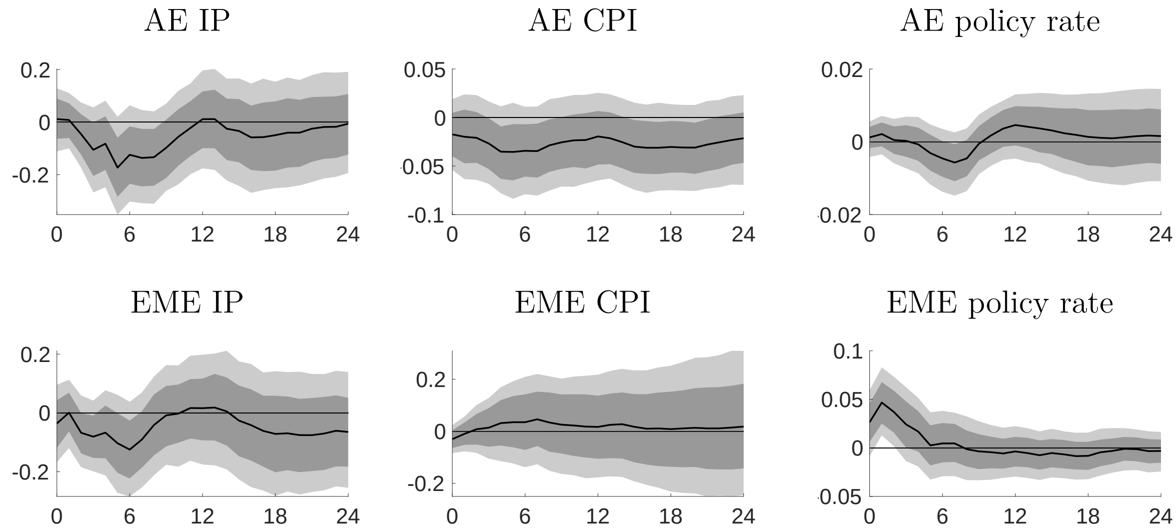
Figure 2.8: Spillbacks from US monetary policy through AEs and EMEs



Notes: The black solid lines depict the response of US industrial production to a US monetary policy shock from VAR models in which RoW industrial production is replaced with separate measures for AEs and EMEs industrial production and their impulse responses to a US monetary policy shock are not constrained; the grey shaded areas represent the associated 68% and 90% centered point-wise probability bands. The light blue lines with triangles depict counterfactual impulse responses in which spillovers to the RoW are precluded, meaning that both AEs and EMEs industrial production are constrained to not respond to a US monetary policy shock. In the left-hand side panel the dark blue line with crosses depicts the counterfactual impulse response of US industrial production when AEs industrial production is constrained to not respond while EMEs industrial production is set to respond as in the unconstrained case. In the right-hand side panel the dark blue line with crosses depicts the counterfactual impulse response when AEs industrial production is constrained to respond as in the unconstrained case while EMEs industrial production is constrained to not respond to a US monetary policy shock. All counterfactual impulse responses shown are based on SSA with all shocks. Figure B.10 in the Appendix provides results for MRE.

This finding has implications for the notion that spillbacks from US monetary policy weaken the case for more extensive forms of international monetary policy coordination (Fischer 2014; Yellen 2019). Figure 2.9 presents estimates of the spillovers from US monetary policy to consumer prices and policy rates in AEs and EMEs. While real activity spillovers are negative both in AEs and EMEs, consumer prices fall in AEs but rise in EMEs. Thus, while US monetary policy spillovers do not induce welfare-reducing trade-offs between output and inflation in AEs, they do so in EMEs. This implies that global welfare could benefit if US monetary policy internalised spillovers to EMEs independently of spillbacks.

Figure 2.9: US monetary policy spillovers to AEs and EMEs



Notes: The figure shows point-wise posterior mean impulse responses (black solid lines) together with 68% and 90% centered point-wise probability bands (grey areas). The impulse responses are estimated from a VAR model in which all additional variables are added simultaneously to the vector of endogenous variables.

2.4 Conclusion

Counterfactual analyses suggest that spillbacks account for a substantial fraction of the overall domestic effects of US monetary policy. They materialise through stock market wealth and Tobin's q/cash flow effects; net trade does not contribute to spillbacks because US monetary policy affects exports and imports similarly. Spillbacks materialise through AEs rather than through EMEs, consistent with the composition of the US foreign equity portfolio and exports. This suggests that US monetary policy only internalises part of the spillovers it emits to the RoW through spillbacks. Moreover, our evidence suggests that while US monetary policy spillovers do not give rise to trade-offs between output and inflation in AEs, they do so in EMEs. This implies that global welfare could benefit if US monetary policy internalised spillovers to EMEs independently of spillbacks.

Chapter 3

Global risk and the dollar

with Georgios Georgiadis and Gernot J. Müller

Abstract

The dollar is a safe-haven currency and appreciates when global risk goes up. We investigate the dollar's role for the transmission of global risk to the world economy within a Bayesian proxy structural vectorautoregressive model. We identify global risk shocks using high-frequency asset-price surprises around narratively selected events. Global risk shocks appreciate the dollar, induce tighter global financial conditions and a synchronized contraction of world economic activity. We benchmark these effects against counterfactuals in which the dollar does not appreciate. In the absence of dollar appreciation, the contractionary impact of a global risk shock is much weaker, both in the rest of the world and the US. For the rest of the world, contractionary financial channels thus dominate expansionary expenditure switching when global risk rises and the dollar appreciates

Keywords: Dollar exchange rate, global risk shocks, international transmission, Bayesian proxy structural VAR

JEL-Classification: F31, F42, F44

3.1 Introduction

According to the received wisdom the dollar appreciates when global risk goes up. Figure 3.1 presents the Global Financial Crisis (GFC) and the COVID-19 pandemic as striking examples. This co-movement is a general pattern of the data and testifies to a fundamental asymmetry in a global financial system centered around the dollar.[43] While the dollar's position can be rationalized on the ground that some assets are particularly safe or liquid (Farhi and Gabaix 2016; He et al. 2019; Gopinath and Stein 2021; Chahrour and Valchev 2022; Eren and Malamud 2022), the role of its appreciation in the transmission of global risk is unclear: Does it help the world economy in coping with global risk shocks or does it amplify their adverse impact?

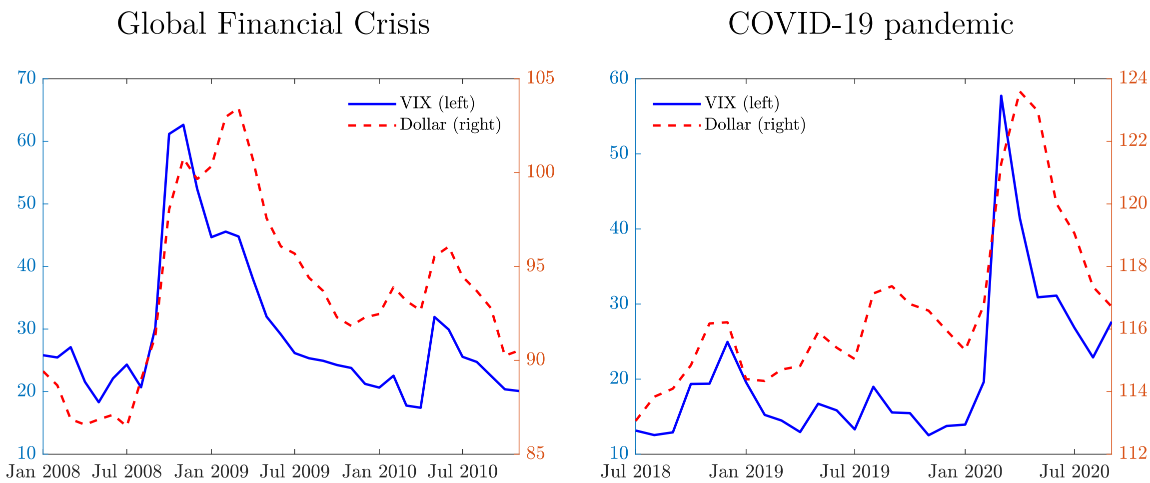
We shed light on this question by exploring the net effect of dollar appreciation in the transmission of global risk. We first upgrade the received wisdom to rigorous causal evidence using a state-of-the-art structural vector-autoregressive (VAR) model identified using narrative external instruments. We show that exogenous global risk shocks induce an appreciation of the dollar. They furthermore contract economic activity in the US and the rest of the world (RoW). Reflecting a trade channel, US net exports fall, suggesting that dollar appreciation induces expenditure switching in the RoW (Gopinath et al. 2020b). Reflecting a financial channel, global equity prices drop, spreads increase and cross-border bank credit contracts (Bruno and Shin 2015; Jiang et al. 2023; Kekre and Lenel 2021).

Second, we construct three conceptually distinct counterfactuals that simulate the effects of a global risk shock in the *absence* of dollar appreciation. The first counterfactual is based on the estimated VAR model and explores the most likely path of the endogenous variables conditional on a global risk shock in a scenario in which the dollar happens to not appreciate because additional, offsetting shocks materialize as well (Antolin-Diaz et al. 2021). The second counterfactual is a VAR-based policy-rule experiment in which the Federal Reserve (Fed) stabilizes the dollar exchange rate conditional on a global risk shock (McKay and Wolf 2023). The third counterfactual is based on a structural model for the US and the RoW in which the deep parameters can be modified so that the dollar does not have a dominant status in cross-border credit and safe assets, which is responsible for the appreciation upon a global risk shock in the first place.

We find that while in all no-appreciation counterfactuals a global risk shock still causes a slowdown in US and RoW activity, the contraction is substantially reduced relative to the baseline by about—depending on the horizon and the methodology—30-50%. Without dollar appreciation the response of US net exports hardly changes. Expenditure switching thus contributes little to the transmission of a global risk shock to the rest of the world through dollar appreciation. While financial conditions still move in the absence of dollar

[43] In a regression of changes in the VIX on changes in the dollar exchange rate over the period 01/1990-12/2020 the t-value is 5.8, and 2.2 when excluding the period 7/2008-12/2009 and after 03/2020. Consistent with the findings in Lilley et al. (2022), the t-value is essentially zero for the time period prior to the GFC, it is 4.3 for the post-GFC period 1/2010-12/2020, and 3.6 for the inter-crises period 1/2010-3/2020.

Figure 3.1: The US dollar and the VIX



Note: VIX is an index of expected stock market volatility compiled by Chicago Board of Options Exchange; dollar is the price of dollar expressed in foreign currency (in effective terms) such that an increase represents an appreciation.

appreciation, they tighten by much less. The financial channel thus plays a key role in the transmission of a global risk shock through dollar appreciation. Put differently, the contractionary effects that materialize through tighter financial conditions thus dominate expansionary effects through expenditure switching.

In more detail, we estimate a Bayesian proxy structural VAR (BPSVAR) model using the approach of Arias et al. (2021). Specifically, we extend the closed-economy VAR model of Gertler and Karadi (2015) which features US industrial production, the 1-Treasury bill rate, the excess bond premium, and consumer prices and include the dollar nominal effective exchange rate, the 5-Treasury bill rate, the VXO, RoW industrial production and policy rates.

In order to identify a global risk shock we rely on an external instrument (Mertens and Ravn 2013). As in Piffer and Podstawski (2018) we use gold-price changes in narrow intra-daily windows around the time stamps of global risk events selected narratively originally by Bloom (2009). We estimate the model on monthly data for the period 1990–2019. In order to speak to the theoretical literature on the dominant role of the dollar, we consider extended specifications with US exports and imports, cross-border bank credit to non-US borrowers, the Emerging Markets Bond Index (EMBI) spread, and RoW equity prices.

We find that a global risk shock appreciates the dollar and contracts US and RoW industrial production. US and RoW monetary policy loosen. Consistent with a trade channel, US net exports fall. Consistent with a financial channel, cross-border credit contracts, RoW equity prices fall and the EMBI spread rises.

We then construct no-appreciation counterfactuals to assess the dollar's contribution to the transmission of a global risk shock to the RoW. The first counterfactual is implemented in the BPSVAR model based on the idea that the dollar does not appreciate because a series of additional, offsetting shocks materialize (Antolin-Diaz et al. 2021). Specifically, we cast impulse responses into a forecast conditioned on a global risk shock occurring in period t and subject to the constraint that the dollar does not appreciate along the forecast horizon $t, t+1, \ldots, t+H$. The offsetting shocks that materialize along the forecast horizon and enforce the constraint are chosen so as to be as small as possible and least correlated, hence deviating minimally from the baseline of a standard global risk shock impulse response. Intuitively, this counterfactual can be thought of as the most likely scenario in which the dollar does not appreciate following a global risk shock and which could be observed in practice.

The second counterfactual assumes the Fed deviates from its actual policy rule and stabilizes the dollar exchange rate. McKay and Wolf (2023) show that even without knowing the true structural model a policy-rule counterfactual can be recovered in a VAR model using a set of period-t policy shocks. To implement this counterfactual, we additionally identify conventional Fed funds rate and forward guidance shocks. Following McKay and Wolf (2023) we then choose the size of these shocks so that when they materialize together with a global risk shock in period t the dollar stays at its baseline value over horizons $t, t+1, \ldots, t+H$.

The third counterfactual is based on a structural model for the US and the RoW in which the dollar appreciates upon a global risk shock because of the interplay between dollar dominance in safe assets and cross-border finance (Georgiadis et al. 2023). In the model, when global risk aversion goes up and the world economy contracts, holding US Treasuries increasingly loosens balance-sheet constraints of RoW banks indebted in foreign currency, which causes the Treasury convenience yield to rise and the dollar to appreciate. To implement a counterfactual in which the dollar does not appreciate upon a global risk shock, we shut down dollar dominance in cross-border finance and safe assets. Intuitively, this can be thought of as showing how a global risk shock would play out in a counterfactual world in which the dollar does not appreciate for structural reasons other than variation in the policy rule.

In all counterfactuals the contractionary effect of a global risk shock still causes a slowdown in US and RoW activity, but the contraction is substantially reduced relative to the baseline by about 30-50%. This implies that in the baseline the contractionary effects that operate via the financial channel dominate the expansionary effects that operate via the trade channel.

Related literature. Our empirical analysis speaks to theoretical work on the special role of the dollar and US assets in the international monetary system (Gopinath et al. 2020b; Jiang et al. 2023; Kekre and Lenel 2021; Bianchi et al. 2021; Devereux et al. 2022). Our analysis assesses the empirical relevance of the mechanisms spelled out in this work. More generally, our analysis informs the theoretical literature on the role of exchange rates for the cross-border transmission of shocks through financial channels (Banerjee et al. 2016;

Aoki et al. 2018; Akinci and Queralto 2019; Croce et al. 2022).

Our paper is also related to empirical work that studies the role of the dollar as a global risk factor (Lustig et al. 2014; Verdelhan 2018), the predictive power of convenience yields (Engel and Wu 2018; Valchev 2020; Jiang et al. 2021) and global risk (Lilley et al. 2022; Hassan et al. 2021) for the dollar, as well as the relationship between global risk, deviations from covered interest parity, the dollar and cross-border credit (Avdjiev et al. 2019; Erik et al. 2020). We complement this work by moving from forecasting and reduced-form regressions to isolating the effects of exogenous variation in global risk.

Our paper furthermore contributes to empirical work on the role of financial channels in the global transmission of risk shocks (Liu et al. 2017; Cesa-Bianchi et al. 2018; Epstein et al. 2019; Shousha 2019; Bhattarai et al. 2020). Relative to existing work, we zoom in on and quantify the role of the dollar within the broader class of financial channels. Our findings on the role of the dollar for financial spillovers complement existing evidence from micro data (Shim et al. 2021; Bruno and Shin 2023; Niepmann and Schmidt-Eisenlohr 2017). Relative to this work, our analysis allows us to contrast trade and financial channels and hence assess the net effect of dollar appreciation.

Finally, our paper is related to the literature on shock identification using external instruments in VAR models (Mertens and Ravn 2013; Gertler and Karadi 2015; Caldara and Herbst 2019). In contrast to much of the existing work we employ the Bayesian estimation approach of Arias et al. (2021) to jointly identify several structural shocks by means of multiple external instruments and use exact finite sample inference in order to bypass questions about the appropriate asymptotic inference in the presence of multiple and potentially weak instruments (Jentsch and Lunsford 2019; Montiel Olea et al. 2021). Moreover, we postulate only relatively weak additional exogeneity assumptions in order to avoid set-identification and difficulties in posterior inference (Baumeister and Hamilton 2015; Giacomini and Kitagawa 2021).

The rest of the paper is structured as follows. Section 3.2 lays out the BPSVAR framework and describes our empirical specification. Section 3.3 presents results for the effects of global risk shocks in the data. Section 3.4 explores no-appreciation counterfactuals. Section 3.5 concludes.

3.2 Empirical strategy

We first outline the BPSVAR framework of Arias et al. (2021) and discuss our specification and identification assumptions. We keep the discussion short and refer to the working paper version of this paper for details (Georgiadis et al. 2021).

3.2.1 The BPSVAR framework

Consider the structural VAR model

$$\boldsymbol{y}_t'\boldsymbol{A}_0 = \boldsymbol{y}_{t-1}'\boldsymbol{A}_1 + \boldsymbol{\epsilon}_t', \tag{3.1}$$

where $\boldsymbol{y}_t$ is an $n \times 1$ vector of endogenous variables and $\boldsymbol{\epsilon}_t$ an $n \times 1$ vector of structural shocks. Assume there is a $k \times 1$ vector of observed proxy variables—or, in alternative jargon, external instruments—$\boldsymbol{p}_t$ that are correlated with the k unobserved structural shocks of interest $\boldsymbol{\epsilon}_t^*$ (relevance condition) and orthogonal to the remaining unobserved structural shocks $\boldsymbol{\epsilon}_t^o$ (exogeneity condition):

$$E[\boldsymbol{p}_t\boldsymbol{\epsilon}_t^{*\prime}] = \boldsymbol{V}, \qquad E[\boldsymbol{p}_t\boldsymbol{\epsilon}_t^{o\prime}] = \boldsymbol{0}. \tag{3.2}$$

Arias et al. (2021) develop a Bayesian algorithm that imposes these assumptions in the estimation of the VAR model in Equation (3.1) augmented with equations for the proxy variables. The estimation thereby identifies the structural shocks.

3.2.2 BPSVAR model specification

Our point of departure is the closed-economy US VAR model of Gertler and Karadi (2015), which includes in $\boldsymbol{y}_t$ the logarithms of US industrial production and consumer prices, the excess bond premium of Gilchrist and Zakrajsek (2012), and the 1-year Treasury bill rate as monetary policy indicator. We augment $\boldsymbol{y}_t$ with the VXO, the logarithm of an index of non-US, RoW industrial production, a weighted average of advanced economies' (AEs) policy rates, the 5-year Treasury bill rate, and the logarithm of the US dollar nominal effective exchange rate (NEER).[44] We use monthly data for the time period from February 1990 to December 2019 and flat priors for the VAR parameters. Data descriptions are provided in Table C.1 in the Appendix.

3.2.3 Identification

For ease of exposition, we first only discuss the identification of the global risk shock given as our key shock of interest. We explain in Section 3.4.2 how we additionally identify the US monetary policy shocks we use in one of the counterfactuals.

We think of a global risk shock as an incident that is associated with an exogenous drop in investors' risk appetite, which can be understood as a shock to the price—as opposed to the quantity—of risk (Miranda-Agrippino and Rey 2020b; Bauer et al. 2023).

[44]We use AE instead of RoW policy rates as the latter exhibit spikes reflecting periods of hyperinflation in some EMEs. In the Appendix we consider an extension in which we include AE and EME industrial production, prices and policy rates separately (Figure C.1). Furthermore, we document in the Appendix that the results are robust to including a measure of RoW prices (Figure C.2).

The proxy variable $p_t^{\epsilon,r}$ for the global risk shock is based on intra-daily data in the spirit of work on the high-frequency identification of monetary policy shocks (e.g. Gertler and Karadi 2015). In particular, we use intra-daily changes in the price of gold around the time stamps of narratively selected events originally selected by Bloom (2009) and later updated by Piffer and Podstawski (2018) and Bobasu et al. (2021). We consider the events labeled as 'global' and 'US' by Piffer and Podstawski (2018). We assume global risk shocks drive gold-price surprises on the narratively selected events, that is in the relevance condition in Equation (3.2) we have $E[p_t^{\epsilon,r}\epsilon_t^r] \neq 0$. The intuition is that an increase in global risk raises the price of the archetypical safe asset of gold (Baur and McDermott 2010; Ludvigson et al. 2021).

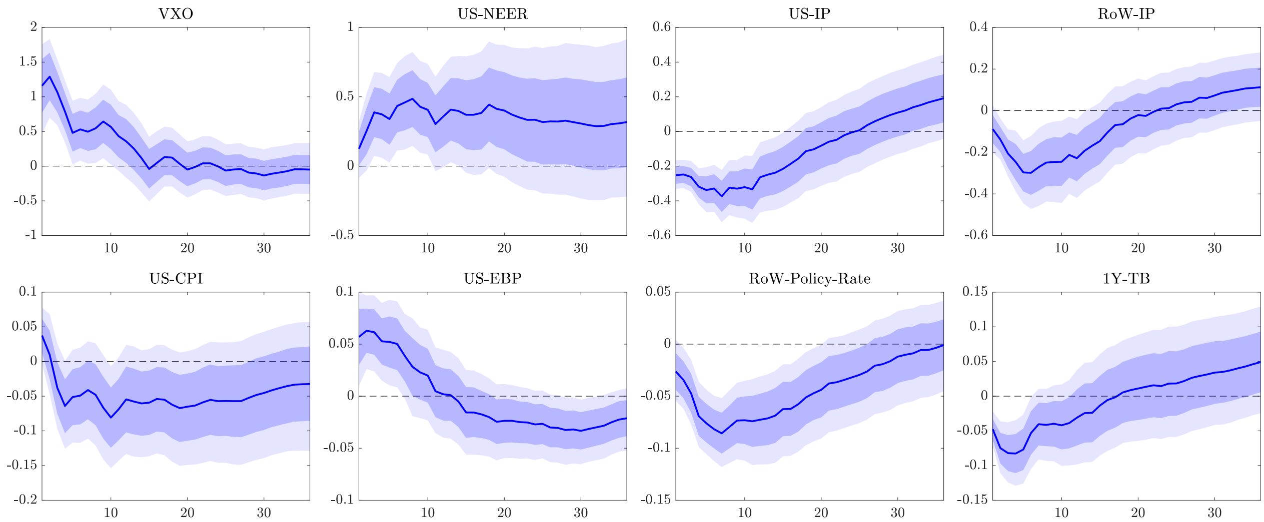
Regarding the exogeneity condition $E[p_t^{\epsilon,r}\boldsymbol{\epsilon}_t^o] = 0$ in Equation (3.2), Piffer and Podstawski (2018) document that the intra-daily gold-price surprises on the narratively selected events are not systematically correlated with a range of measures of non-risk shocks. In other words, we assume the only shock that occurred systematically in the intra-daily windows across the narratively selected events is the global risk shock. Note that what is critical for the exogeneity condition to be satisfied is that across the full list of narratively selected events the gold-price surprises around the intra-daily windows were driven *systematically* only by global risk shocks. For this, the selection of events and the width of the intra-daily windows around the corresponding time stamps rather than the specific asset price are crucial. We explore robustness checks for both aspects below.

Finally, for consistency we follow Caldara and Herbst (2019) as well as Arias et al. (2021) and impose a 'relevance threshold' to express a prior belief that the proxy variables are relevant instruments: We require that at least 10% of the variance of the proxy variables is accounted for by the identified shocks; this is less than the 20% required by Arias et al. (2021), and—although not straightforward to compare conceptually—lies below the 'high-relevance' prior of Caldara and Herbst (2019). Specifying the relevance threshold at 10% implies there is a lot of room for the proxy variable measurement error in the BPSVAR model to account for events on which global risk shocks occurred but the recorded gold-price surprise is zero as they are not selected by Bloom (2009), Piffer and Podstawski (2018) and Bobasu et al. (2021). We explore robustness checks without relevance threshold below.

3.3 The effect of global risk shocks on the world economy

Figure 3.2 shows our first result: A one-standard deviation global risk shock increases the VXO and appreciates the dollar. This implies the positive co-movement between global risk and the dollar shown in Figure 3.1 is at least to some extent accounted for by global risk shocks. US and RoW industrial production both contract, but the effect in the US is more immediate and somewhat larger. US consumer prices fall after a short delay and the excess bond premium rises. US and RoW monetary policy are loosened.

Figure 3.2: Impulse responses to a global risk shock



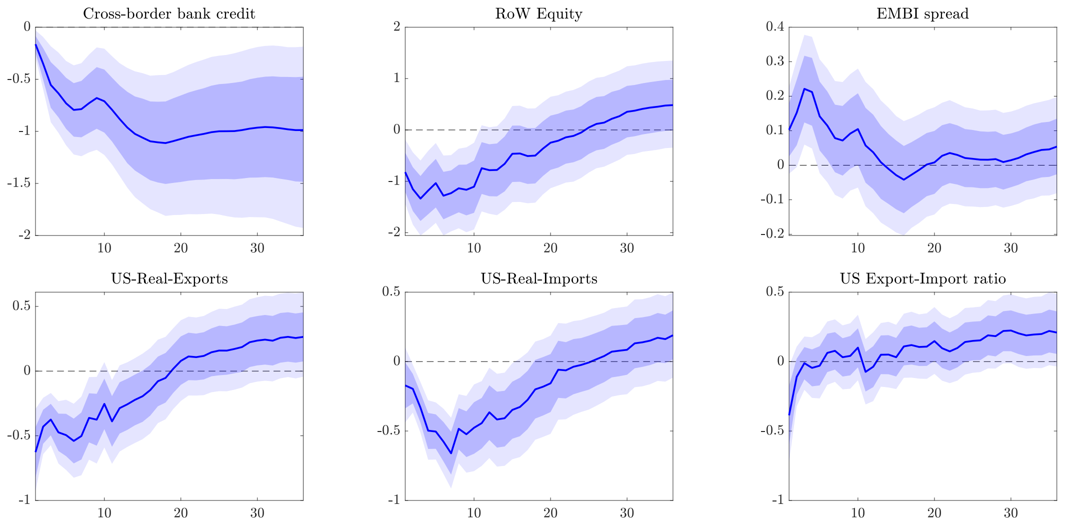
Note: Horizontal axis measures time in months, vertical axis deviation from pre-shock level; size of shock is one standard deviation; blue solid line represents point-wise posterior mean and shaded areas 68%/90% equal-tailed, point-wise credible sets. VXO measured in levels, the dollar NEER, US and RoW industrial production, US consumer prices in logs, and the excess bond premium, the RoW policy as well as the US 1-year Treasury Bill rates in percent.

Figure 3.3 presents the responses of global financial conditions and US trade. Consistent with a financial channel, cross-border bank credit to non-US borrowers declines, RoW equity prices contract and spreads increase. Consistent with a trade channel through expenditure switching US net exports contract.[45]

In the Appendix we present results for several extensions. We document that in response to a global risk shock: also other safe-haven currencies such as the Japanese yen and the Swiss franc appreciate, while non safe-haven currencies such as the euro and the British pound depreciate (Figure C.3); the price of safety in terms of the Treasury premium of Du et al. (2018) increases; consistent with the model of Bianchi et al. (2021) banks raise the ratio of safe and liquid dollar assets to liabilities (Figure C.3); that there is evidence for 'fear-of-floating' as EME monetary policy tightens at the same time as output contracts (Figure C.1); when we additionally impose forecast error variance decomposition restrictions in the spirit of Francis et al. (2014) to disentangle shocks to the price—risk appetite—and the quantity of risk—uncertainty—both shocks appreciate the dollar and exhibit qualitatively similar patterns, but the impulse responses to the global risk appetite shock correspond more closely to those from our baseline (Figure C.4).

[45]That the contraction is more immediate in US exports than imports is consistent with dominant-currency paradigm (DCP) in trade invoicing (Gopinath et al. 2020b). As under DCP US export prices are sticky in dollar, a dollar appreciation induces immediate expenditure switching in the RoW. In contrast, as also RoW export prices are sticky in dollar, there is no expenditure switching in the US; the response of US

Figure 3.3: Impulse responses of trade and financial variables to a global risk shock



Note: See notes to Figure 3.2.

In the Appendix we also document that the estimated effects of global risk shocks hardly change if: as in Ludvigson et al. (2021) we relax the exogeneity condition and only impose $|E[p_t^{\epsilon,r}\epsilon_t^r]| > |E[p_t^{\epsilon,r}\epsilon_t^\ell]|$ for $\ell \neq r$ (Figure C.5); we address concerns that the gold-price surprises calculated over windows of several hours are contaminated by other shocks occurring close to the narratively selected event time stamps by considering long-term Treasury yield and US dollar/euro exchange rate surprises over narrower windows (Figure C.6 and C.7); we abandon the narratively selected events and instead consider monthly changes in the Geopolitical Risk Index of Caldara and Iacoviello (2022) as proxy variable (Figure C.8); we estimate a larger BPSVAR model with many more US and RoW variables (Figure C.9); we do not impose a relevance threshold (Figure C.10).

3.4 The role of the dollar

Our results suggest the dollar appreciation caused by a global risk shock impacts the RoW through both a trade and a financial channel. And given that the appreciation impacts RoW real activity with different signs depending on the channel, its net effect is ambiguous. In this section we determine the net effect by benchmarking the baseline impulse responses against a counterfactual in which the dollar does not appreciate. To robustify our analysis, we consider three conceptually distinct no-appreciation counterfactuals.

imports to a global risk shock is then driven only by the hump-shaped contraction in US demand.

3.4.1 A possible empirical scenario

The first approach is based on structural scenario analysis (SSA; Antolin-Diaz et al. 2021, ADPRR). ADPRR develop SSA as a flexible framework for conditional forecasts. We apply SSA to construct a no-appreciation counterfactual. In particular, we first represent the impulse responses as conditional forecasts for a system that is in its steady state in period $t-1$ and then hit by a single shock in period t. Then we determine the smallest and least correlated shocks that would have to materialize in periods $t, t+1, \ldots, t+h$ in order to offset the effect of the period-t global risk shock on the dollar. Intuitively, this counterfactual can be thought of as the most likely scenario that could be observed in practice in which the dollar does not appreciate upon a global risk shock.

Formally, assume for simplicity but without loss of generality that the VAR model in Equation (3.1) is stationary and that it does not include deterministic terms. After iterating forward from period t to $t+h$ we have

$$\boldsymbol{y}_{t,t+h} = \boldsymbol{b}_{t,t+h} + \boldsymbol{M}'\boldsymbol{\epsilon}_{t,t+h}, \tag{3.3}$$

where the $n(h+1) \times 1$-vectors $\boldsymbol{y}_{t,t+h} \equiv (\boldsymbol{y}_t', \boldsymbol{y}_{t+1}', \ldots, \boldsymbol{y}_{t+h}')'$ and $\boldsymbol{\epsilon}_{t,t+h} \equiv (\boldsymbol{\epsilon}_t', \boldsymbol{\epsilon}_{t+1}', \ldots, \boldsymbol{\epsilon}_{t+h}')'$ stack the endogenous variables and structural shocks for periods $t, t+1, \ldots, t+h$, respectively, the $n(h+1) \times n(h+1)$ matrix $\boldsymbol{M} = \boldsymbol{M}(\boldsymbol{A}_0, \boldsymbol{A}_1)$ represents the effects of these structural shocks in terms of impulse responses, and $\boldsymbol{b}_{t,t+h}$ period-(t–1) initial conditions. Assume further the VAR model is in steady state in period $t-1$ so that $\boldsymbol{b}_{t,t+h} = \boldsymbol{0}$. The impulse responses to a period-t global risk shock are then given by the forecast $\boldsymbol{y}_{t,t+h}$ conditional on $\boldsymbol{\epsilon}_{t,t+h}$, with $\epsilon_t^r = 1$, $\epsilon_{t+s}^r = 0$ for $s > 0$ and $\epsilon_{t+s}^\ell = 0$ for $s \geq 0$, $\ell \neq r$.

In order to obtain the counterfactual conditional forecast $\widetilde{\boldsymbol{y}}_{t,t+h}$ SSA determines a series of additional shocks $\widetilde{\boldsymbol{\epsilon}}_{t,t+h}$ that materialize over periods $t, t+1, \ldots, t+h$ and whose effects offset the dollar appreciation caused by the period-t global risk shock. This no-appreciation constraint can be written as $\widetilde{\boldsymbol{C}}\widetilde{\boldsymbol{y}}_{t,t+h} = \boldsymbol{0}$, where the $(h+1) \times n(h+1)$ matrix $\widetilde{\boldsymbol{C}}$ selects the conditional forecast of the dollar over periods $t, t+1, \ldots, t+h$.[46] Constraints on the structural shocks are written as $\boldsymbol{\Xi}\widetilde{\boldsymbol{\epsilon}}_{t,t+h} = \boldsymbol{g}_{t,t+h}$, where the $k_s \times n(h+1)$ matrix $\boldsymbol{\Xi}$ first selects the period-t global risk shock and then any $k_s - 1$ structural shocks over periods $t, t+1, \ldots, t+h$ that shall take on specific values to enforce the no-appreciation constraint.

ADPRR show how to obtain the SSA solution $\widetilde{\boldsymbol{\epsilon}}_{t,t+h}$ which satisfies the counterfactual no-appreciation constraint $\widetilde{\boldsymbol{C}}\widetilde{\boldsymbol{y}}_{t,t+h} = \boldsymbol{0}$ and the constraint on the set of structural shocks $\boldsymbol{\Xi}\widetilde{\boldsymbol{\epsilon}}_{t,t+h} = \boldsymbol{g}_{t,t+h}$. The solution implies the counterfactual impulse response $\widetilde{\boldsymbol{y}}_{t,t+h} = \boldsymbol{M}'\widetilde{\boldsymbol{\epsilon}}_{t,t+h}$.

In order to stay agnostic and let the data select the most likely offsetting shocks—see below for the intuition—we perform SSA without constraint on the set of structural shocks.[47]

[46] Ordering the dollar last in $\boldsymbol{y}_t$, we have $\widetilde{\boldsymbol{C}} = \boldsymbol{I}_{h+1} \otimes \boldsymbol{e}_n'$, where $\boldsymbol{e}_i$ is $n \times 1$-vector of zeros with unity at the i-th position.

[47] Ordering the global risk shock last in $\boldsymbol{\epsilon}_t$, we have $\boldsymbol{g}_{t,t+h} = 1$ and $\boldsymbol{\Xi} = [\boldsymbol{e}_n', \boldsymbol{0}_{1\times hn}]$, where $\boldsymbol{e}_i$ is an $n \times 1$ vector of zeros with unity at the i-th position.

Incidentally, this also means we do not have to identify any additional structural shocks. Technically, this is because any orthogonal decomposition of the reduced-form shocks (i.e. any set of additionally identified structural shocks) that satisfies the exogeneity condition in Equation (3.2) would produce the same result (Section 2.1 of ADPRR).

Because in every period $t, t+1, \ldots, t+h$ we have up to $n > 1$ shocks to impose the no-appreciation constraint, there is a multiplicity of SSA solutions. ADPRR show that in this case the SSA solution minimizes the Frobenius norm of the deviation of $\widetilde{\boldsymbol{\epsilon}}_{t,t+h}$ from their baseline value of zero and their baseline variance matrix. This means the solution selects the smallest and least correlated shocks that enforce the no-appreciation constraint. We therefore interpret the SSA counterfactual as reflecting the most likely scenario in which the dollar does not appreciate following a global risk shock which could be observed in practice.

The first column in Figure 3.4 shows the SSA counterfactual together with the baseline impulse responses. In response to a global risk shock the dollar does not appreciate by assumption, and both US and RoW real activity drop less than in the baseline; the reduction in the recessionary impact of the global risk shock amounts to up to 30%.[48]

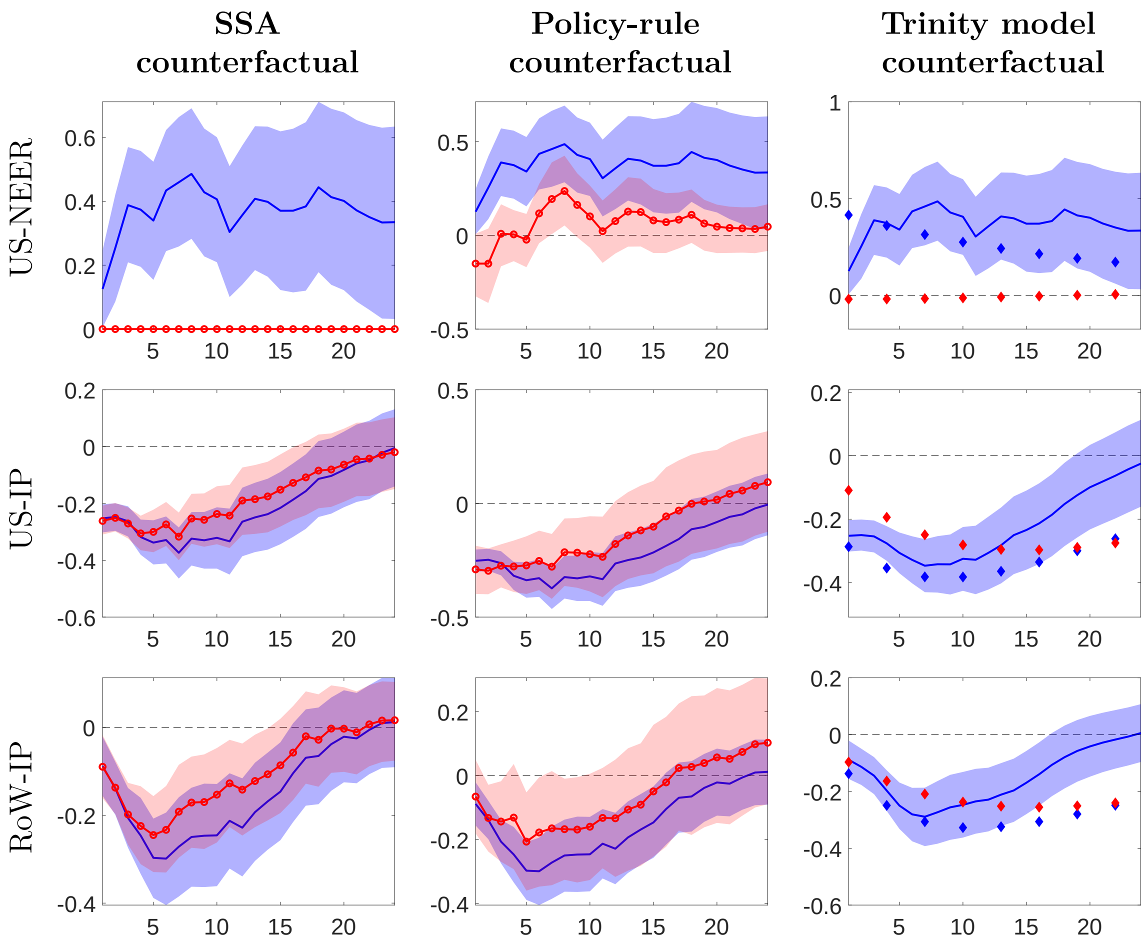
The first column in Figure 3.5 shows the SSA counterfactual together with the baseline impulse responses for variables reflecting the trade and financial channels. Two results stand out. First, consistent with the absence of expenditure switching, when the dollar does not appreciate US net exports drop by a little less. This suggests dollar appreciation is expansionary for the RoW in the baseline through the trade channel, although the latter is not very powerful. Second, although financial conditions still move in the counterfactual, they tighten by much less. This suggests dollar appreciation is contractionary through the financial channel in the baseline, and that the latter is rather powerful. Together with our findings for RoW activity, these results suggest the net effect of dollar appreciation upon a global risk shock is contractionary for the RoW and that the financial channel dominates the trade channel.[49]

The SSA counterfactual is appealing at a conceptual level because it uses those offsetting shocks which are most likely to materialize in practice and is otherwise agnostic about the nature of these shocks. Yet for this very reason it is not possible to tell why the dollar does not appreciate in the counterfactual. In what follows, we therefore complement the SSA counterfactual with two alternatives which allow for a structural interpretation. The first alternative counterfactual we consider has a concrete economic interpretation as a monetary-policy-rule counterfactual. In particular, we next explore how a global risk shock would affect the RoW if the Fed were to stabilize the dollar.

[48]In Figure C.11 in the Appendix we show that about 90% of the posterior probability mass of the difference between the baseline and the counterfactual is larger than zero.

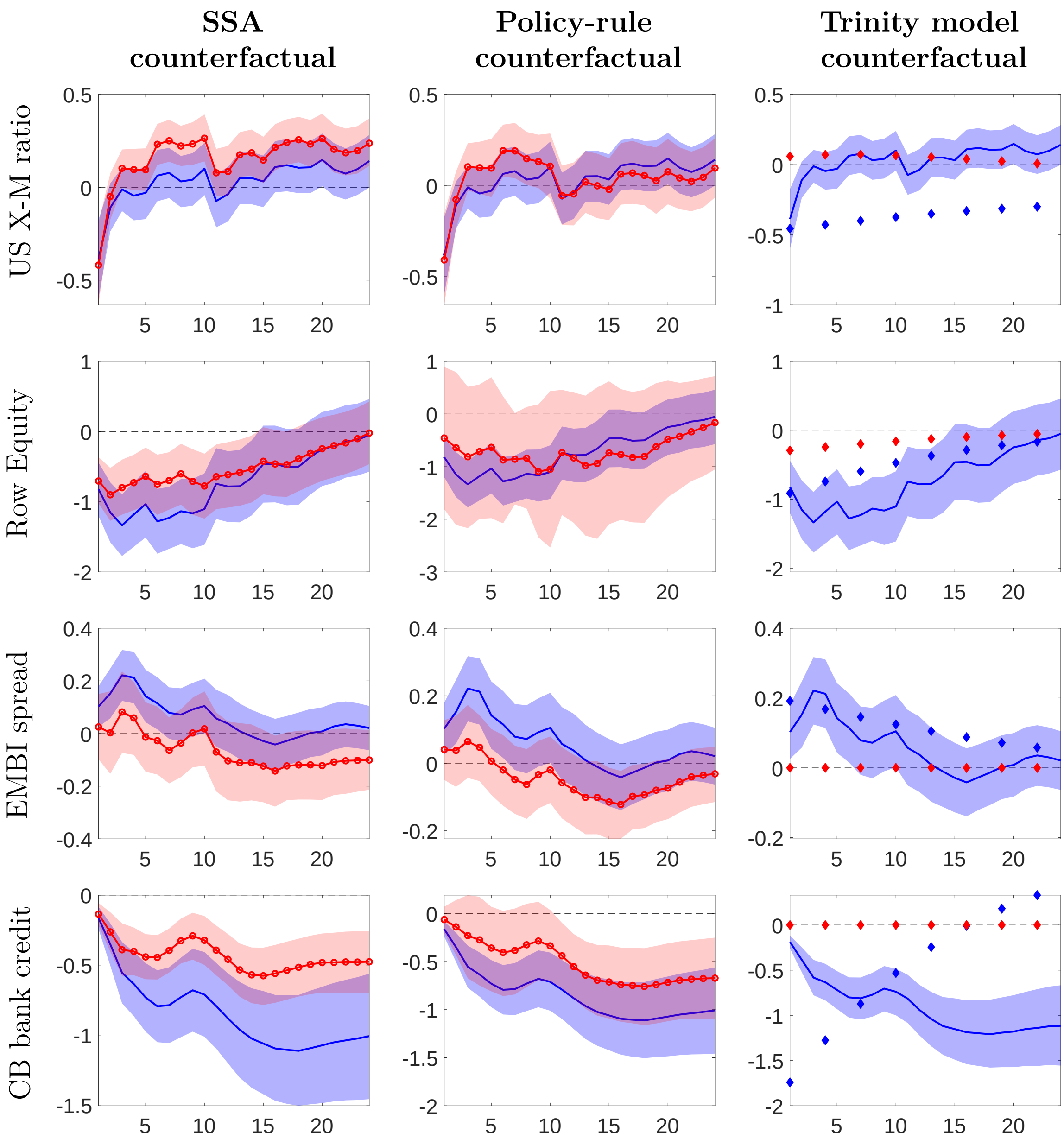
[49]We report the counterfactual impulse responses for the remaining variables in the BPSVAR model in the Appendix (Figure C.13).

Figure 3.4: Baseline and counterfactual responses to a global risk shock



Note: The figure shows the baseline BPSVAR model (blue solid) and counterfactual (red circled) impulse responses to a global risk shock. SSA counterfactuals are shown in the first column, policy-rule counterfactuals in the second column, and the trinity-model counterfactuals in the third column. The red (blue) shaded areas represent 68% credible sets obtained from computing the counterfactual (impulse responses) for each draw from the posterior distribution. In the third column, the blue (red) diamonds depict the baseline (counterfactual) impulse responses to a global risk aversion shock in the trinity model. We do not connect the dots depicting the counterfactual because the trinity model is calibrated to quarterly frequency while the BPSVAR model is estimated at the monthly frequency. The global risk aversion shock in the trinity model is scaled such that the average of the response of the dollar over the first year is the same as the response from the BPSVAR model. The real GDP (output) response in the trinity model is multiplied by 2.5 to make it comparable to the industrial production response from the BPSVAR model given that in the data the latter is 2.5 times more volatile than the former. In the Appendix we document that the BPSVAR model impulse response of S&P Global's US monthly GDP is indeed about 2.5 times smaller than for US industrial production (Figure C.12), while their time profiles are rather similar.

Figure 3.5: Baseline and counterfactual responses of trade and financial variables to a global risk shock



Note: See notes to Figure 3.4. As the trinity model does not include an exact match for equity prices (the EMBI spread) we plot the response of the price of capital (RoW cross-border credit spread) instead. In the counterfactual structural model dollar dominance is absent so that standard UIP holds. Therefore any exchange-rate-adjusted cross-border border return differential is zero.

3.4.2 What if the Fed stabilized the dollar?

Standard uncovered interest rate parity (UIP) logic suggests the Fed could prevent dollar appreciation upon a global risk shock by loosening more than in the baseline. One would then expect the additional Fed loosening to be expansionary for the RoW and hence mitigate the contractionary effects of the global risk shock. We next consider a policy-rule counterfactual in the BPSVAR model to explore this rigorously.

VAR-based policy counterfactuals using structural shocks have a long history in the literature (e.g. Sims and Zha 2006). Typically, these counterfactuals are constructed in an SSA-like fashion with unexpected policy shocks materializing every period along the entire impulse-response horizon $t, t+1, \ldots, t+h$. These counterfactuals are often conceived as a change in the policy rule (for example Kilian and Lewis 2011). However, this approach may be subject to the Lucas critique and in general does not recover the true policy-rule counterfactual McKay and Wolf (2023, henceforth MW). Intuitively, this is because it is assumed that although agents are being repeatedly surprised they do not adjust their expectations about future policy behaviour. Put differently, this approach ignores a possible expectations channel through which a policy-rule change may impact the economy.

MW develop an approach for constructing policy-rule counterfactuals in VAR models that is robust to the Lucas critique and recovers the true policy-rule counterfactual for a broad range of underlying structural frameworks, including standard representative and heterogeneous-agent New Keynesian models. In particular, they show that using appropriate impact-period-t *news* shocks about current and future policy recovers the impulse responses that would prevail under a counterfactual policy rule. Put differently, MW show that using appropriate period-t policy news shocks is an equivalent approach to explicitly simulating a shock under a counterfactual policy rule.

Formally, motivated by the representation of structural models in sequence space introduced by Auclert et al. (2021), MW consider a linear, perfect-foresight, infinite-horizon economy in terms of deviations from the deterministic steady state for periods $t = 0, 1, 2, \ldots$ summarized by

$$\boldsymbol{\mathcal{H}}_x \boldsymbol{x} + \boldsymbol{\mathcal{H}}_z \boldsymbol{z} + \boldsymbol{\mathcal{H}}_\epsilon \boldsymbol{\epsilon} = \mathbf{0}, \tag{3.4}$$

$$\boldsymbol{\mathcal{A}}_x \boldsymbol{x} + \boldsymbol{\mathcal{A}}_z \boldsymbol{z} + \boldsymbol{\nu} = \mathbf{0}, \tag{3.5}$$

where $\boldsymbol{x} \equiv (\boldsymbol{x}_1', \boldsymbol{x}_2', \ldots, \boldsymbol{x}_{n_x}')'$ stacks the time paths of the n_x endogenous variables, analogously $\boldsymbol{z}$ the n_z policy instruments, $\boldsymbol{\epsilon}$ the n_ϵ non-policy structural shocks and $\boldsymbol{\nu}$ the n_ν policy news shocks; the latter are deviations from the policy rule announced at date t but implemented only in some future period $t+s$, $s > 0$. The key assumption reflected in Equations (3.4) and (3.5) is that $\{\boldsymbol{\mathcal{H}}_x, \boldsymbol{\mathcal{H}}_z, \boldsymbol{\mathcal{H}}_\epsilon\}$ do not depend on the coefficients of the policy rule $\{\boldsymbol{\mathcal{A}}_x, \boldsymbol{\mathcal{A}}_z\}$, so that policy affects the private sector's decisions only through the path of the instrument $\boldsymbol{z}$, rather than through the policy rule *per se*. Under some mild assumptions the solution to Equations (3.4) and (3.5) can be written using impulse response coefficients $\boldsymbol{\Theta}_{\mathcal{A}}$ as

$$\begin{pmatrix} \boldsymbol{x} \\ \boldsymbol{z} \end{pmatrix} = \boldsymbol{\Theta}_{\mathcal{A}} \times \begin{pmatrix} \boldsymbol{\epsilon} \\ \boldsymbol{\nu} \end{pmatrix}, \quad \boldsymbol{\Theta}_{\mathcal{A}} \equiv (\boldsymbol{\Theta}_{\epsilon,\mathcal{A}}, \boldsymbol{\Theta}_{\nu,\mathcal{A}}) \equiv \begin{pmatrix} \boldsymbol{\Theta}_{x,\epsilon,\mathcal{A}} & \boldsymbol{\Theta}_{x,\nu,\mathcal{A}} \\ \boldsymbol{\Theta}_{z,\epsilon,\mathcal{A}} & \boldsymbol{\Theta}_{z,\nu,\mathcal{A}} \end{pmatrix}. \tag{3.6}$$

MW show that knowledge of the impulse responses $\boldsymbol{\Theta}_{\mathcal{A}}$ under the baseline policy rule is sufficient to determine the impulse responses to the structural shock ϵ under any counterfactual policy rule $\tilde{\boldsymbol{\mathcal{A}}}_x \boldsymbol{x} + \tilde{\boldsymbol{\mathcal{A}}}_z \boldsymbol{z} = \boldsymbol{0}$ as

$$\boldsymbol{x}_{\tilde{\boldsymbol{\mathcal{A}}}}(\epsilon) = \boldsymbol{x}_{\boldsymbol{\mathcal{A}}}(\epsilon) + \boldsymbol{\Theta}_{x,\nu,\mathcal{A}} \times \tilde{\boldsymbol{\nu}}, \qquad \boldsymbol{z}_{\tilde{\boldsymbol{\mathcal{A}}}}(\epsilon) = \boldsymbol{z}_{\boldsymbol{\mathcal{A}}}(\epsilon) + \boldsymbol{\Theta}_{z,\nu,\mathcal{A}} \times \tilde{\boldsymbol{\nu}}. \tag{3.7}$$

In particular, the impulse response to the structural shock ϵ under the counterfactual policy rule is given by the sum of the corresponding impulse responses to the same structural shock under the baseline policy rule $\boldsymbol{x}_{\boldsymbol{\mathcal{A}}}(\epsilon)$ and the impulse responses to some policy news shocks $\tilde{\boldsymbol{\nu}}$. The latter are chosen so that the counterfactual policy rule

$$\tilde{\boldsymbol{\mathcal{A}}}_x \left[\boldsymbol{x}_{\boldsymbol{\mathcal{A}}}(\epsilon) + \boldsymbol{\Theta}_{x,\nu,\mathcal{A}} \times \tilde{\boldsymbol{\nu}}\right] + \tilde{\boldsymbol{\mathcal{A}}}_z \left[\boldsymbol{z}_{\boldsymbol{\mathcal{A}}}(\epsilon) + \boldsymbol{\Theta}_{z,\nu,\mathcal{A}} \times \tilde{\boldsymbol{\nu}}\right] = \boldsymbol{0} \tag{3.8}$$

holds. The intuition is that as long as the private sector's decisions depend on the path of the policy instrument rather than the rule *per se* it does not matter whether the path comes about due to the systematic conduct of policy or due to policy news shocks.

A practical challenge of this approach is that news shocks $\tilde{\boldsymbol{\nu}}$ which communicate changes in future policy over all possible horizons $t, t+1, t+2, \ldots$ are rarely available to the econometrician. However, MW show that in practice one can use a set of standard monetary policy shocks $\boldsymbol{s}$ and their impulse responses $\boldsymbol{\Omega}_{\boldsymbol{s},\mathcal{A}}$ from the empirical literature as long as each entails a different future path of the policy instrument. And MW show that rather than requiring impulse responses to as many shocks as horizons over which the counterfactual policy-rule is assumed, using even only a small number of shocks $\boldsymbol{s}$ that solve

$$\min_{\boldsymbol{s}} \; || \tilde{\boldsymbol{\mathcal{A}}}_x \left[\boldsymbol{x}_{\boldsymbol{\mathcal{A}}}(\epsilon) + \boldsymbol{\Omega}_{x,s,\mathcal{A}} \times \boldsymbol{s}\right] + \tilde{\boldsymbol{\mathcal{A}}}_z \left[\boldsymbol{z}_{\boldsymbol{\mathcal{A}}}(\epsilon) + \boldsymbol{\Omega}_{z,s,\mathcal{A}} \times \boldsymbol{s}\right] ||, \tag{3.9}$$

produces a reliable "best Lucas-critique-robust approximation".

Against this background, we explore how a global risk shock would affect the RoW if the Fed were to stabilize the dollar. As in Wolf (2023), we specify the counterfactual policy rule implicitly as $\boldsymbol{e}_{usd}\boldsymbol{x} = \boldsymbol{0}$, where $\boldsymbol{e}_{usd}$ is a $1 \times n_x$-vector of zeros with unity at the position of the dollar in $\boldsymbol{x}_t$. Confining the counterfactual to periods $t = 0, 1, 2, \ldots, h$, Equation (3.9) becomes

$$\min_{\boldsymbol{s}} \; || \boldsymbol{e}_{usd}\boldsymbol{x}_{\boldsymbol{\mathcal{A}},t,t+h}(\epsilon) + \boldsymbol{\Omega}_{x,s,\mathcal{A}} \times \boldsymbol{s} ||, \tag{3.10}$$

which boils down to solving a least-squares minimization problem for n_s unknown period-t Fed policy shocks $\boldsymbol{s}$ in $h+1$ equations.

We implement this policy-rule counterfactual using $n_s = 2$ distinct US monetary policy shocks in $\boldsymbol{s}$, just like MW do in their illustration. In particular, in addition to the global risk shock we jointly identify a conventional monetary policy and a forward guidance shock using similar proxy variables as Miranda-Agrippino and Rey (2020a) and Miranda-Agrippino and Nenova (2022), namely intra-daily surprises in the 3-month Federal funds futures and the 5-year Treasury bill rate in a narrow window around FOMC announcements as proxy variables.[50]

We estimate that a contractionary conventional monetary policy shock raises interest rates at relatively shorter horizons, as given by the 1-year Treasury bill rate (Figure C.14). In turn, we estimate that a contractionary forward guidance shock raises interest rates at relatively longer horizons, as given 5-year Treasury bill rate (Figure C.15). Both shocks slow down real activity in the US and the RoW, tighten global financing conditions, and appreciate the dollar.[51, 52] To implement the policy-rule counterfactual we let both a conventional and a forward guidance Fed policy shock occur in period t together with the global risk shock, and choose their size so that they offset as much as possible—in a least squares sense—the response of the dollar.

The middle columns in Figures 3.4 and 3.5 present the results for this policy-rule counterfactual. Despite the conceptual difference, the results of this policy-rule counterfactual are quite similar to those of the SSA counterfactual.[53] In the Appendix we show that a non-trivial additional Fed easing in terms of the 1 and 5-year Treasury bill rates is required

[50]This means that in Equation (3.1) the structural shocks of interest are given by $\boldsymbol{\epsilon}_t^* \equiv (\epsilon_t^r, \epsilon_t^{cmp}, \epsilon_t^{fg})'$, where ϵ_t^{cmp} and ϵ_t^{fg} denote the conventional monetary policy and forward guidance shocks, respectively, and the corresponding proxy variables are given by $\boldsymbol{p}_t \equiv (p_t^{\epsilon,r}, p_t^{\epsilon,3m}, p_t^{\epsilon,5y})'$. In Equation (3.2) we impose the additional identifying assumptions that the 3-month and 5-year-rate surprises are not driven by the global risk shock, $E[p_t^{\epsilon,3m}\epsilon_t^r] = E[p_t^{\epsilon,5y}\epsilon_t^r] = 0$ (Gertler and Karadi 2015; Miranda-Agrippino and Rey 2020b). Note that these assumptions imply two zeros in the first column of $\boldsymbol{V}$ in Equation (3.2), which are sufficient to point-identify of the global risk shock. It would be intuitive to go further and impose that $\boldsymbol{V}$ is diagonal to disentangle the conventional monetary policy and forward guidance shocks, but this would imply over-identifying restrictions and cannot be implemented in the estimation algorithm of Arias et al. (2021). To nonetheless disentangle the two monetary policy shocks we impose magnitude restrictions. In particular, we assume that the 3-month-rate (5-year-rate) surprise is affected more strongly by the conventional monetary policy (forward guidance) shock than by the forward guidance (conventional monetary policy) shock, that is $E[p_t^{\epsilon,3m}\epsilon_t^{cmp}] > E[p_t^{\epsilon,3m}\epsilon_t^{fg}]$ and $E[p_t^{\epsilon,5y}\epsilon_t^{fg}] > E[p_t^{\epsilon,5y}\epsilon_t^{cmp}]$.

[51]Because we only have data on the 5-year-rate surprises from 1996, as in Känzig (2021) we replace the missing values in the 3-month Federal funds futures and the 5-year Treasury bill rate surprises by zero (see Noh 2017, for a formal justification of this approach). However, Figures C.16 and C.17 in the Appendix document that our results are robust to starting the estimation in 1996. We follow Miranda-Agrippino and Nenova (2022) and apply the poor-man's approach of Jarocinski and Karadi (2020) and purge these surprises from central bank information effects on the basis of the sign of the corresponding equity-price surprise.

[52]We document that results are similar if instead of the 3-month and 5-year-rate surprises we use as proxy variables the conventional monetary policy and forward guidance surprises of Jarocinski (2021) or Lewis (2024), which both also account for central bank information effects (Figures C.18 to C.21).

[53]The dollar is not perfectly stabilized because we use only $n_s = 2$ rather than $h + 1$ policy shocks in Equation (3.10). However, our results are similar if we use a third US monetary policy shock (i.e. $n_s = 3$) identified by 10-year Treasury bill rate surprises as proxy variable so that the dollar is more stable upon a global risk shock (Figure C.22). Moreover, in the Appendix C.6 we document that using the structural model we discuss in Section 3.4.3 below as a laboratory, the true policy-rule counterfactual is approximated fairly well with the approach of MW and only two distinct policy (news) shocks (Figure C.30).

to stabilize the dollar (Figure C.23) and that at the posterior mean the required policy shocks are both expansionary and equal about half of their standard deviation (Figure C.24).

3.4.3 A world economy without structural dollar dominance

The VAR-based counterfactuals take the non-policy structure of the world economy as given and explore what would happen if offsetting shocks materialized or if the Fed were to stabilize the dollar. Although the latter provides a well-defined structural explanation for the absence of appreciation, it explicitly leverages changes in policy and thereby intertwines the effect of the dollar appreciation with the change in the policy rate. Therefore, as an alternative, one may consider changing the non-policy features of the world economy that underpin the dollar's response to a global risk shock in the first place. Hence, in what follows we construct a third counterfactual based on a structural two-country model for the US and the RoW. The model matches the empirical impulses responses in Figure 3.2 and allows us to modify the non-policy structural features so that the dollar does not appreciate upon a global risk shock.

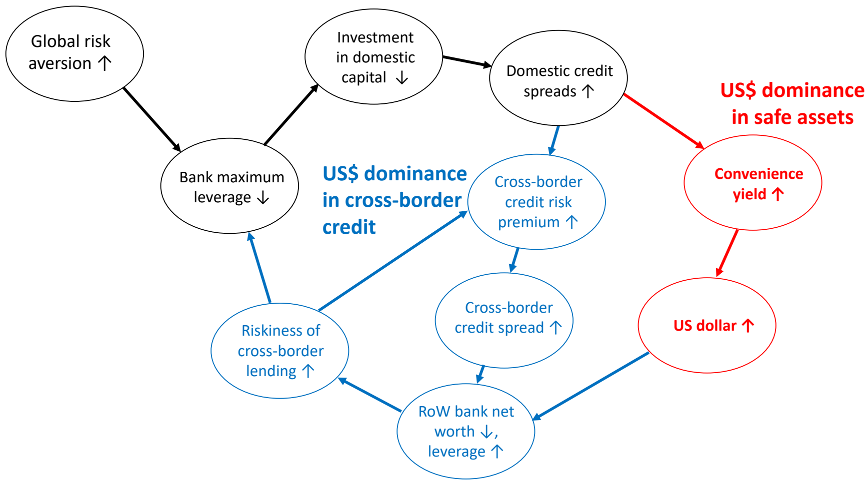
We draw on the two-country model for the US and the RoW with dollar dominance in cross-border credit, safe assets and trade invoicing developed in Georgiadis et al. (2023). Laying out the structure of this 'trinity model' is beyond the scope of this paper, and so we only provide an intuitive description.[54] In the model, US banks intermediate domestic dollar funds to banks in the RoW. Cross-border dollar borrowing is cheap but also risky relative to domestic funding, and therefore tightens RoW banks' balance-sheet constraints. Because they are viewed as the global safe asset, US Treasuries are held as liquidity-buffers by RoW banks to loosen balance-sheet constraints and thereby earn an additional, indirect pecuniary return that can be interpreted as a convenience yield.

In the trinity model dollar dominance in cross-border credit and safe assets interact so that the dollar appreciates in response to a global risk shock. In particular, an increase in global risk aversion—modeled as an exogenous reduction in the willingness of creditors to provide funding to banks for a given level of net worth—raises domestic credit spreads so that leveraging up by loosening the balance-sheet constraint becomes more profitable, which causes the Treasury convenience yield to rise, and eventually the dollar to appreciate. This dollar appreciation triggers a global financial accelerator. In particular, as RoW banks' cross-border dollar borrowing is not perfectly hedged by holdings of Treasuries, dollar appreciation reduces their net worth. As a result, the balance-sheet constraint of US banks as the lenders of RoW banks tightens and forces them to deleverage, which raises US and RoW domestic credit spreads further.

Figure 3.6 summarizes the mechanics of this global financial accelerator and highlights how dollar dominance in safe assets and cross-border credit interact to give rise to dollar

[54]We provide a detailed discussion of the model, all equations and the calibration in Appendix C.4.

Figure 3.6: The global financial accelerator in the trinity model of Georgiadis et al. (2023)



Note: The figure presents a schematic overview of the global financial accelerator in the dollar trinity model of Georgiadis et al. (2023).

appreciation when risk aversion rises: Dollar dominance in safe assets underpins a dollar appreciation when global risk aversion rises, and dollar dominance in cross-border credit a global financial accelerator when the dollar appreciates.[55]

The right-hand side columns in Figures 3.4 and 3.5 show that the impulse responses to a global risk aversion shock for the baseline calibration of the trinity model (blue dots) match the BPSVAR model impulse responses (blue solid lines) fairly well.[56]

For the counterfactual we assume the dollar does not hold any dominant position in the world economy: There is no cross-border dollar credit and RoW banks do not demand Treasury securities as safe asset. The counterfactual impulse responses (red dots) show that without dollar dominance the dollar does not appreciate when global investors' risk aversion increases (Figure 3.4), that global financial conditions in terms of equity valuations, spreads and cross-border credit tighten by less (Figure 3.5), and that output drops by less both in the US and the RoW (Figure 3.4). The reason is that without dollar dominance in safe assets, holding Treasuries no longer loosens balance-sheet constraints of RoW banks and hence does not earn a convenience yield. As a result, the dollar does not appreciate when global investors' risk aversion increases. And without dollar appreciation and dollar dominance in cross-border credit there is no global financial accelerator mechanism that amplifies the effect of a global risk aversion shock on the RoW. Finally, US net exports fall

[55]There is an additional amplification channel shown in the middle of Figure 3.6 that arises because also cross-border credit spreads rise as US banks' balance-sheet constraints tighten, which reduces RoW banks' net worth independently from the dollar appreciation.

[56]In order to make percentage deviations of flow variables—such as output—from the quarterly business-cycle model comparable to those from the monthly BPSVAR model we report the three-month trailing moving average of the latter's impulse responses as suggested by Born and Pfeifer (2014).

by less—in fact rise—in the absence of dollar dominance.[57]

Taken together, the results from the trinity-model counterfactuals are consistent with those for the SSA and the policy-rule counterfactuals. We uniformly find that the contractionary financial channel dominates the expansionary trade channel. The net effect of dollar appreciation upon a global risk shock is contractionary for the RoW.

3.5 Conclusion

In this paper we provide evidence that global risk shocks cause a dollar appreciation and a slowdown in world real activity. In order to shed light on the dollar's role in the international transmission of global risk, we construct three conceptually distinct no-appreciation counterfactuals. The results suggest robustly that without dollar appreciation the slowdown in global economic activity would be much weaker. This raises important normative questions about the design of the international financial architecture that underpin the key role of the dollar in the global economy. These are, however, beyond the scope of the present paper.

[57]In the Appendix we document that results are similar in a trinity-model policy-rule counterfactual in which the Fed stabilizes the dollar exchange rate rather than US output and inflation (Figures C.27 to C.29).

Chapter 4

Dominant-currency pricing and the global output spillovers from US dollar appreciation

with Georgios Georgiadis

Abstract

We test for the empirical relevance of partial and asymmetric dominant-currency pricing (DCP), the hypothesis that a large share of global export and import prices are sticky in US dollar. We first set up a structural three-country New Keynesian dynamic stochastic general equilibrium model which nests DCP, producer-currency pricing (PCP) and local-currency pricing (LCP). Under DCP, the output spillovers from shocks that appreciate the US dollar decline with an economy's export-import US dollar pricing share differential, i.e. the difference between the share of an economy's export and import prices that are sticky in US dollar. Underlying this prediction is variation in an economy's net exports in response to US dollar appreciation that arises because the shares of export and import prices that are sticky in US dollar are different. We then document that this prediction from partial and asymmetric DCP is consistent with the data in a sample of up to 45 advanced and emerging market economies for the time period from 1995 to 2018. We moreover document that our findings are robust to considering US demand, US monetary policy and exogenous exchange rate shocks as a trigger of US dollar appreciation, zooming in on the responses of economies' exports and imports, as well as to accounting for the role of commodity trade in US dollar invoicing.

Keywords: Dominant-currency pricing, US dollar appreciation, output spillovers.
JEL-Classification: F42, E52, C50.

4.1 Introduction

In the classic Mundellian international macroeconomic model it is assumed that export prices are sticky in the producer's currency (producer-currency pricing; PCP; Mundell 1963; Fleming 1962; Obstfeld and Rogoff 1995). An alternative suggested more recently is to assume that export prices are sticky in the importer's currency (local-currency pricing; LCP; Betts and Devereux 1996, 2000; Engel 2000; Devereux and Engel 2003). And even more recently, it has been suggested as yet another alternative that export prices are sticky in a dominant currency regardless of their destination, for example in US dollar (dominant-currency pricing; DCP; Gopinath et al. 2020b). Importantly, the dynamic relationships between key macroeconomic variables in response to domestic and US shocks are rather different under DCP, PCP and LCP. For example, in contrast to widespread intuitions shaped by the implicit assumption of PCP, under DCP pass-through of changes in the exchange rate against non-dominant currencies is small, while that of a change in an economy's exchange rate against the dominant currency is high. As a consequence, under DCP the US dollar exchange rate is driving global trade rather than only transactions in which the US is involved. Because different export pricing paradigms have different implications for a wide variety of issues including business cycle co-movement, optimal monetary policy and currency areas, and international monetary policy co-ordination, it is critical for academics and policymakers to know which pricing paradigm fits the data best.

We contribute to the literature by providing new evidence for the empirical relevance of DCP and by addressing some important shortcomings of existing work. In particular, we study an exogenous trigger of exchange rate appreciation instead of reduced-form regressions, partial rather than full DCP, and the role of DCP for output rather than trade price and volume spillovers. Under partial DCP we understand a setting in which the share of exports and imports priced in the dominant currency is in general less than 100%, and a setting in which the shares in general differ across exports and imports; hence, strictly speaking, we consider partial and asymmetric DCP instead of full and symmetric DCP. We first show that in a standard structural open-economy model with DCP the output spillovers from shocks that appreciate the US dollar multilaterally decline with an economy's non-US export-import DCP share differential, i.e. the difference between the share of an economy's non-US exports and imports that are subject to DCP. The intuition is that as an economy's share of exports that is subject to DCP rises relative to the corresponding share of imports, a multilateral appreciation of the US dollar elicits stronger expenditure switching in the economy's trading partners than domestically; hence the multilateral appreciation of the US dollar produces a larger decline in the economy's exports than in its imports, which implies that its net exports and hence output fall.

We then exploit this prediction from the model in order to test for the relevance of partial DCP in the data. In particular, because of the lack of data on economies' export-import US dollar DCP share differentials, we test the *joint* hypothesis that the data are characterised by partial DCP and that the export *invoicing* currency coincides with the *pricing* currency, that is the currency in which export prices are sticky. More concretely, we test empirically whether economies' non-US export-import US dollar *invoicing* share differentials are negatively related to cross-country differences in the estimated output spillovers from shocks that appreciate the US dollar multilaterally. We find that the data are consistent with the hypothesis that a large share of global trade is subject to partial DCP and that the currency of invoicing coincides with the currency in which export prices are sticky: The estimated output spillovers from shocks that appreciate the US dollar multilaterally are statistically significantly and negatively related to economies' non-US export-import US dollar invoicing share differentials.

In more detail, we first set up a three-country New Keynesian dynamic stochastic general equilibrium model for the US and the rest of the world. We split the rest of the world in two regions of equal size in order to study the effects of US dollar appreciation on non-US trade under partial DCP. Importantly, the model we consider nests the cases of PCP, LCP and DCP as well as combinations thereof. We then explore the output spillovers from US demand, monetary policy and uncovered interest rate parity (UIP) shocks that appreciate the US dollar multilaterally, and in particular how these vary across different assumptions regarding the currency in which export prices are sticky. In particular, in case of PCP the output spillovers from multilateral US dollar appreciation to the rest of the world are positive: Multilateral US dollar appreciation increases the rest of the world's net exports vis-à-vis the US, as it reduces (raises) the rest of the world's imports from (exports to) the US through expenditure switching. In contrast, in case of LCP the output spillovers from US dollar appreciation to the rest of the world are muted. Specifically, as both US and rest of the world export prices are sticky in the importer's currency, multilateral US dollar appreciation does not trigger significant expenditure switching under LCP. Most importantly, however, both in case of PCP and LCP there is no expenditure switching in trade between the economies in the rest of the world as their exchange rates depreciate symmetrically against the US dollar.

In case of *full* DCP as considered in Gopinath et al. (2020b) all export—and hence all import—prices are sticky in US dollar regardless of the destination of exports and the origin of imports, which implies that multilateral US dollar appreciation elicits expenditure switching in the bilateral trade between the economies in the rest of the world. Specifically, assume for simplicity that the rest of the world does not trade with the US. Because prices of intra-regional imports in the rest of the world are sticky in US dollar under full DCP, multilateral US dollar appreciation triggers expenditure switching away from imports from other economies in the rest of the world towards domestically-produced goods. As a mirror image, an economy's exports to other economies in the rest of the world accordingly decline due to these expenditure switching effects. Most importantly, however, because multilateral US dollar appreciation reduces intra-regional exports and imports through expenditure switching to the exact same degree in all economies in the rest of the world under full DCP, intra-regional net exports and hence output are unchanged.

In case of *partial* DCP that we consider in this paper—that is the shares of an economy's exports and imports that are subject to DCP being generally below 100% and different from each other—the output spillovers from multilateral US dollar appreciation that stem from intra-regional trade in the rest of the world are in general different from zero. For example, in case of a non-US economy featuring a positive export-import DCP share differential—i.e. a larger share of exports than imports being subject to DCP—US dollar appreciation reduces its imports from other economies in the rest of the world by less through expenditure switching than it reduces its exports to other economies in the rest of the world. As a result, the output spillovers from US dollar appreciation that arise through intra-regional trade with other economies in the rest of the world are smaller—more negative or less positive—for an economy with a positive non-US export-import DCP share differential. The model thus implies a testable prediction regarding the magnitude of output spillovers from US dollar appreciation under partial DCP: The output spillovers from US dollar appreciation decline with economies' non-US export-import DCP share differentials. We document that this prediction from partial DCP is qualitatively robust to changes in the model parametrisation, considering a variety of shocks that induce US dollar appreciation such as positive US demand, contractionary US monetary policy and exogenous UIP shocks, as well as to the inclusion of additional model features such as trade in intermediate inputs for production, sticky wages, capital and financial frictions.

Testing the prediction of partial DCP in the data is not straightforward. For one thing, the shares of economies' export and import prices that are *sticky* in different currencies and which are needed in order to calculate the export-import DCP share differentials are not observed. In fact, if these shares were observed, it would be trivially obvious whether the data are characterised by partial DCP or not. Therefore, we exploit data on the shares of economies' export and import prices that are *invoiced* in different currencies in order to test for the empirical relevance of partial DCP. In particular, we consider the *joint* hypothesis that (i) the data are characterised by partial DCP and that (ii) the currency in which export prices are sticky coincides with the currency in which exports are invoiced. Against the background of the theoretical analysis, the prediction from this joint hypothesis is that output spillovers from US dollar appreciation are negatively related to economies' non-US export-import US dollar *invoicing*—rather than DCP—share differentials. It is important to notice that the currency in which export prices are sticky does not need to coincide with the currency in which exports are invoiced; a case in point are commodity exports, which are typically invoiced in US dollar but whose prices are typically believed to not be sticky in US dollar (or in any other currency). To the best of our knowledge, while there is some evidence for selected advanced economies, evidence on the relationship between the currency in which exports are invoiced and in which their prices are sticky does not exist for trade between emerging market economies.

We then confront the prediction of the joint hypothesis that the data are characterised by partial DCP and that the currency in which export prices are sticky coincides with the currency in which exports are invoiced with the data. Specifically, we first estimate the spillovers from US dollar appreciation using two-country VAR models for up to 45 advanced and emerging market economies for the time period from 1995 to 2018. The two-country VAR models consist of the US and one non-US economy at the time. We identify US demand, US monetary policy and exogenous UIP shocks using model-consistent sign restrictions or, alternatively, using external instruments in proxy structural VAR models. In a second step, we run cross-sectional regressions of the cross-country differences in the estimated output spillovers from US dollar appreciation on economies' non-US export-import US dollar invoicing share differentials. In the regressions we control for cross-country differences in the susceptibility to other spillover transmission channels that are relevant in the structural model: Economies' susceptibility to bilateral US and multilateral demand channels of spillovers reflected by their overall openness to trade as well as their bilateral trade integration with the US. We also control for differences in the strength of expenditure switching conditional on trade pricing reflected by economies' exchange rate flexibility vis-à-vis the US dollar.

Our empirical findings are consistent with the predictions of partial DCP from the structural model and the hypothesised coincidence between the currency in which export prices are sticky and the currency of invoicing: Economies that exhibit a larger non-US export-import US dollar invoicing share differential exhibit smaller output spillovers from US dollar appreciation. We document that these findings are robust to a range of alternative specifications of the empirical framework both regarding the two-country VAR models and the cross-sectional regressions, in particular zooming in on the responses of exports and imports, controlling for the role of commodity trade in US dollar invoicing. Finally, we also consider the output spillovers from multilateral euro instead of US dollar appreciation as a placebo test. In particular, we document that the data provide less support for DCP in case of the euro, which is consistent with what one would expect given the much less prominent role of the euro in global trade invoicing.

Our paper is related to existing literature. Devereux et al. (2007) discuss the "US dollar standard" as a combination of PCP and LCP in a two-country model for the US and the rest of the world. Cook and Devereux (2006) and Goldberg and Tille (2009) are early contributions discussing the role of DCP for third-country trade. In particular, Cook and Devereux (2006) explore the effects of DCP in a partial equilibrium three-country New Keynesian model and show that it is critical in order to account for the muted dynamics of South East Asian exports against the background of the sizable currency depreciations during the Asian financial crisis. Goldberg and Tille (2009) explore welfare and international monetary policy co-ordination under full DCP in a static three-country general equilibrium model. More recently, Gopinath et al. (2020b) consider a three-country New Keynesian DSGE model that nests PCP, LCP as well as DCP, and they derive testable predictions in order to assess the empirical relevance of the latter. In particular, Gopinath et al. (2020b) show that under DCP trade prices and quantities respond to the US dollar rather than to bilateral exchange rates. Furthermore, Gopinath et al. (2020b) provide extensive empirical evidence based on global trade data for 2500 country pairs as well as customs data for Colombia that is consistent with these predictions from DCP.[58] Finally, Zhang (2018) sets up a two-period three-country general equilibrium model and shows that under DCP exchange rates of economies with larger shares of the consumption basket priced in US dollar exhibit smaller depreciations and larger interest rate rises in response to contractionary US monetary policy shocks. Zhang (2018) then provides empirical evidence based on the response of asset prices to US monetary policy shocks at the daily frequency for a set of advanced economies with developed financial markets that are consistent with the predictions of DCP.

We contribute to this literature by providing new evidence for the empirical relevance of DCP in several dimensions and by addressing some gaps and important shortcomings of existing work. Relative to Gopinath et al. (2020b), rather than considering trade price or volume spillovers from changes in the US dollar exchange rate we contribute to the literature by focusing on spillovers to *output*. This is an important contribution as output is arguably a key variable in economic and policy analysis over and above exports and imports. Moreover, we consider *exogenous* shocks that drive changes in the US dollar exchange rate rather than reduced-form regressions. Considering exogenous shocks is more consistent with the experiments in the theoretical models from which the testable predictions of DCP are derived, and it reduces the risk of omitted variable bias in the empirical analysis. Also, we consider *partial* rather than full DCP and—in particular—*asymmetric* rather than symmetric DCP, which seems to be more relevant empirically given the invoicing landscape documented in Gopinath (2015).[59] Relative to Zhang (2018), we contribute to the literature by assessing the empirical relevance of DCP testing for its predictions for output spillovers from US shocks at the *business cycle* frequency rather than for asset prices at the daily frequency. This is an important addition to the literature because high-frequency analysis does not indicate how persistent the estimated effects are at horizons that are relevant for macroeconomic models and policymakers. Moreover, relative to Zhang (2018) we also include in our analysis a broad range of *emerging market economies* instead of focusing on advanced economies with developed financial markets. This is particularly informative for assessing the empirical relevance of DCP as US dollar invoicing of trade is much more widespread in emerging market than in advanced economies. We also provide evidence for DCP in imports and *exports* using *not only* monetary policy shocks as drivers of US dollar appreciation. Finally, relative to both Gopinath et al. (2020b) and Zhang (2018) we contribute to the literature by highlighting expenditure switching and net exports in non-US trade relationships as a transmission channel for the output spillovers from US dollar appreciation under partial and in particular asymmetric DCP. The mechanism we

[58] Similar evidence is found by Chen et al. (2018) using highly disaggregated customs data for the UK.

[59] While the model considered in Gopinath et al. (2020b) allows for partial DCP in principle, their theoretical analysis focuses on full DCP. Their analysis also does not consider net exports and, in this context, asymmetric DCP.

study is thus distinct from that under which output in a non-US economy is affected by US dollar appreciation because monetary policy tightens as import and hence consumer prices rise, which is the mechanism highlighted in Gopinath et al. (2020b) and Zhang (2018).

The rest of the paper is organised as follows. Section 4.2 presents the structural three-country model and derives the testable prediction. Section 4.3 carries out the empirical analysis. Finally, Section 4.4 concludes.

4.2 The model

In this section we set up a three-country New Keynesian DSGE model to derive predictions regarding the differences in spillovers from shocks that appreciate the US dollar multilaterally across different configurations of partial DCP in economies' imports and exports. The model we set up is similar to the one considered in Gopinath et al. (2020b), but we leave out some features which are not critical for our purposes. Specifically, the model we consider includes the minimum set of features necessary to produce predictions that set partial DCP apart from LCP and PCP: Three economies which trade with each other as well as sticky prices.

4.2.1 Model set up

We consider three large economies, namely E, F and D. D is the issuer of the dominant currency, and E and F are two economies that together form the rest of the world.[60] For expositional convenience, throughout this section we refer to these economies as the US instead of D, EME instead of E, and the rest of the world (RoW) instead of F. Economies generally feature a symmetric structure. Figure 4.1 provides a schematic overview of each domestic economy. Each economy is inhabited by five types of agents: Households, intermediate good producers, domestic intermediate and import final good bundlers, consumption good bundlers, and a monetary authority.

Given the symmetry of the three countries, for expositional convenience we only report the relevant equations for EME (E). The equations are analogous for US (D) and RoW (F).[61]

Households

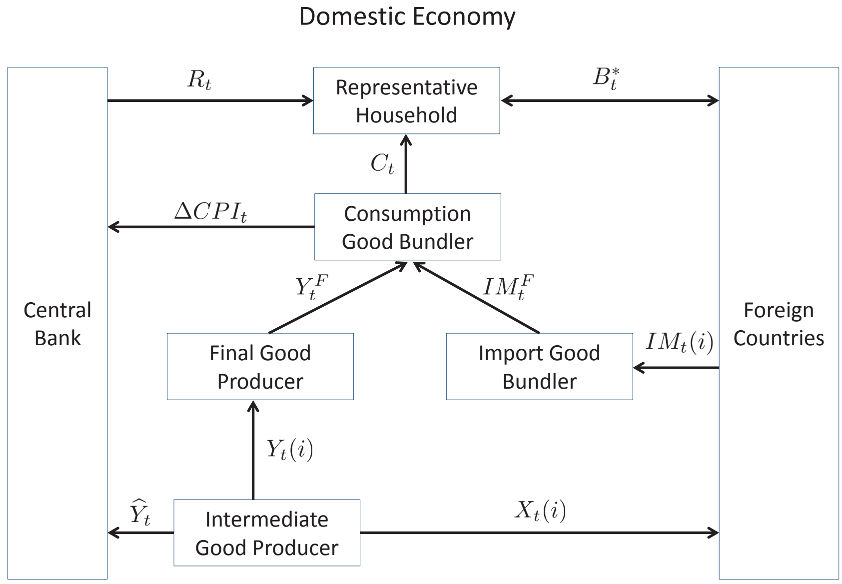
We assume each economy is inhabited by a continuum of symmetric households. In each period a household h consumes a non-traded consumption good bundle $C_{E,t}(h)$ and provides labor service $N_{E,t}(h)$; in the following we omit the reference to household h for simplicity. The period-by-period utility function is given by

$$U(C_{E,t}, N_{E,t}) = \frac{1}{1-\sigma^c} C_{E,t}^{1-\sigma^c} - \frac{\kappa}{1+\varphi} N_{E,t}^{1+\varphi}. \tag{4.1}$$

[60] We discuss this modelling choice in detail below.

[61] A detailed model appendix is available from the authors upon request.

Figure 4.1: Schematic overview of the domestic economy



Note: The figure provides a schematic overview of the structure of each economy in the model laid out in Section 4.2. Import good bundlers exist for imports from each trading partner, in the figure these are collapsed into one bundler only for expositional simplicity.

The household's expected lifetime utility is given by

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \varepsilon_{E,t}^{\beta} U(C_{E,t}, N_{E,t}), \tag{4.2}$$

with β being the household's discount factor and $\varepsilon_{E,t}^{\beta}$ a preference shock that evolves according to a first-order autoregressive process.

We assume financial markets are incomplete internationally and households can trade two risk-free bonds: A domestic bond, $B_{E,t}^E$, denominated in the domestic currency and a foreign bond $B_{E,t}^D$ denominated in US dollar.[62] Finally, households own domestic firms and receive their profits $\Pi_{E,t}$. The household's budget constraint is given by

$$\frac{B_{E,t}}{P_{E,t}^c R_{E,t}} + \frac{\mathcal{E}_{E,t}^D B_{E,t}^D}{P_{E,t}^c R_{E,t}^D} + C_{E,t} = \frac{B_{E,t-1}}{P_{E,t}^c} + \frac{\mathcal{E}_{E,t}^D B_{E,t-1}^D}{P_{E,t}^c} + \frac{W_{E,t} N_{E,t}}{P_{E,t}^c} + \frac{\Pi_{E,t}}{P_{E,t}^c}, \tag{4.3}$$

where we choose the price of the consumption bundle $P_{E,t}^c$ as the numéraire. $R_{E,t}$ and $R_{E,t}^D$ are the domestic and the US dollar interest rate, respectively, and $\mathcal{E}_{E,t}^j$ is the nominal bilateral exchange rate between currencies of EME and j, quoted as units of EME currency required to purchases one unit of currency j.

As in Schmitt-Grohe and Uribe (2003) and Gopinath et al. (2020b), in order to render

[62] Note that the assumption that the international bond is denominated in US dollar does not change the results up to a first-order approximation. We could also allow for trade in two or three internationally traded bonds denominated in different currencies.

the linearised solution of the model stationary and for the foreign bonds to have a well-defined steady state, we attach a risk premium to the US dollar interest rate faced by EME households; the risk premium depends on the aggregate level of US dollar debt relative to its steady-state value. Also, we postulate a shock $\varepsilon_{E,t}^{\mathcal{E}_E^D}$ to the risk premium in order to allow for exogenous changes in the exchange rate that are unrelated to fundamentals (Itskhoki and Mukhin 2017). Specifically, the risk premium is given by

$$R_{E,t}^{D} = R_{D,t} \cdot e^{\left(-\chi \frac{\hat{B}_{E,t}^{D} - \bar{B}_{E}^{D}}{Y_{E,t}} + \varepsilon_{E,t}^{\mathcal{E}_E^D}\right)}. \tag{4.4}$$

Consumption good bundlers

Consumption good bundlers operate under perfect competition and produce a non-traded consumption good that is consumed by domestic households. To do so, they combine the domestically produced final good bundle $Y_{E,t}^{E}$, the non-traded import final good bundle $Y_{F,t}^{E}$ composed of imports from RoW, and the non-traded import final good bundle $Y_{D,t}^{E}$ composed of imports from US according to

$$Y_{E,t}^{c} = \left(n_E^{\frac{1}{\psi_f}} Y_{E,t}^{E\,\frac{\psi_f - 1}{\psi_f}} + \upsilon_E^{F\,\frac{1}{\psi_f}} Y_{F,t}^{E\,\frac{\psi_f - 1}{\psi_f}} + \upsilon_E^{D\,\frac{1}{\psi_f}} Y_{D,t}^{E\,\frac{\psi_f - 1}{\psi_f}} \right)^{\frac{\psi_f}{\psi_f - 1}}, \tag{4.5}$$

where n_E reflects the degree of home bias in the non-traded consumption good bundle, and υ_E^F and υ_E^D represent the steady-state shares of the consumption good bundle accounted for by imports from RoW and US, respectively. Obviously, $n_E + \upsilon_E^F + \upsilon_E^D = 1$, and as bilateral trade is balanced in the steady state $\upsilon_E^F = \upsilon_F^E$ and $\upsilon_E^D = \upsilon_D^E$. The consumer-price index associated with Equation (4.5) is given by

$$P_{E,t}^{c} = \left(n_E P_{E,t}^{E\,1-\psi_f} + \upsilon_E^F P_{F,t}^{E^I\,1-\psi_f} + \upsilon_E^D P_{D,t}^{E^I\,1-\psi_f} \right)^{\frac{1}{1-\psi_f}}, \tag{4.6}$$

where $P_{E,t}^{E}$ defined in Equation (4.8) below denotes the price of the domestic final good bundle, $P_{F,t}^{E^I}$ defined in Equation (4.11) below the EME currency price of the import final good bundle from RoW, and $P_{D,t}^{E^I}$ the EME currency price of the import final good bundle from US.

Cost minimisation implies demand schedules faced by the domestic and import final good bundlers from the consumption good bundler given by

$$Y_{E,t}^{E} = n_E \left(\frac{P_{E,t}^{E}}{P_{E,t}^{c}} \right)^{-\psi_f} Y_{E,t}^{c}, \quad Y_{F,t}^{E} = \upsilon_E^F \left(\frac{P_{F,t}^{E^I}}{P_{E,t}^{c}} \right)^{-\psi_f} Y_{E,t}^{c}, \quad Y_{D,t}^{E} = \upsilon_E^D \left(\frac{P_{D,t}^{E^I}}{P_{E,t}^{c}} \right)^{-\psi_f} Y_{E,t}^{c}. \tag{4.7}$$

Domestic and import final good bundlers

Final good bundlers operate under perfect competition and are specialised based on the origin of the goods they combine. Specifically, the final good composed of domestically produced differentiated intermediate goods and its price index are given by

$$Y_{E,t}^{E}=\left(\int_{0}^{1}Y_{E,t}^{E}(i)^{\frac{\psi_i-1}{\psi_i}}di\right)^{\frac{\psi_i}{\psi_i-1}},\qquad P_{E,t}^{E}=\left(\int_{0}^{1}P_{E,t}^{E}(i)^{1-\psi_i}di\right)^{\frac{1}{1-\psi_i}}. \tag{4.8}$$

Cost minimisation implies the demand schedule faced by domestic differentiated intermediate goods producers from the domestic final good bundler

$$Y_{E,t}^{E}(i)=\left(\frac{P_{E,t}(i)}{P_{E,t}}\right)^{-\psi_i}Y_{E,t}^{E}. \tag{4.9}$$

In turn, the import final good bundle composed of PCP ($\widetilde{Y}_{F,t}^{E}$) and DCP ($\widehat{Y}_{F,t}^{E}$) import good bundles from RoW is given by

$$Y_{F,t}^{E}=\left(\gamma_F^{E\,\frac{1}{\psi_d}}\widetilde{Y}_{F,t}^{E\,\frac{\psi_d-1}{\psi_d}}+(1-\gamma_F^{E})^{\frac{1}{\psi_d}}\widehat{Y}_{F,t}^{E\,\frac{\psi_d-1}{\psi_d}}\right)^{\frac{\psi_d}{\psi_d-1}}, \tag{4.10}$$

where γ_F^E is the share of PCP exports in total RoW exports to EME. The parameter γ_F^E is going to be the focus further below, as we are going to vary $(1-\gamma_F^E)$ in order to explore the variations in the output spillovers from US dollar appreciation that are due to variations in the DCP share. The associated aggregate EME import-price index in EME currency is given by

$$P_{F,t}^{EI}=\left(\gamma_F^{E}\mathcal{E}_{E,t}^{F}\widetilde{P}_{F,t}^{E1-\psi_d}+(1-\gamma_F^{E})\mathcal{E}_{E,t}^{D}\widehat{P}_{F,t}^{E1-\psi_d}\right)^{\frac{1}{1-\psi_d}}, \tag{4.11}$$

where $\widetilde{P}_{F,t}^{E}$ and $\widehat{P}_{F,t}^{E}$ denote the PCP and DCP export-price indices of RoW exports to EME in RoW currency and US dollar, respectively, defined below in Equations (4.19) and (4.20). Cost minimisation implies the demand schedule of the EME import final good bundler for the RoW PCP and DCP import good bundles

$$\widetilde{Y}_{E,t}^{F}=\gamma_F^{E}\left(\frac{\mathcal{E}_{E,t}^{F}\widetilde{P}_{F,t}^{E}}{P_{F,t}^{EI}}\right)^{-\psi_d}Y_{F,t}^{E},\quad \widehat{Y}_{E,t}^{F}=(1-\gamma_F^{E})\left(\frac{\mathcal{E}_{E,t}^{D}\widehat{P}_{F,t}^{E}}{P_{F,t}^{EI}}\right)^{-\psi_d}Y_{F,t}^{E}. \tag{4.12}$$

Intermediate good producers

There is a continuum of firms producing different varieties of intermediate goods under monopolistic competition using labour as the only input

$$Y_{E,t}(i)=N_{E,t}(i). \tag{4.13}$$

We assume all intermediate good producers produce for the domestic and export markets. We also assume intermediate good producers are subject to price-setting frictions à la Calvo. Specifically, each producer i faces a constant probability of $(1-\theta_p)$ of being able to adjust its price. Moreover, we assume that a fraction γ_E^F of EME intermediate good producers is subject to PCP regarding their exports, while the remainder fraction $1-\gamma_E^F$ is subject to DCP.[63,64] Hence, denoting by $Y_{E,t}^E(i)$ domestic sales, by $\widetilde{Y}_{E,t}(i)$ PCP exports and by $\widehat{Y}_{E,t}(i)$ DCP exports, we have that firm production is split into domestic sales and PCP exports $Y_{E,t}(i) = Y_{E,t}^E(i) + \widetilde{Y}_{E,t}(i)$ or into domestic sales and DCP exports $Y_{E,t}(i) = Y_{E,t}^E(i) + \widehat{Y}_{E,t}(i)$, respectively. Finally, we allow for pricing-to-market, i.e. export prices are specific to the destination market.

An EME firm sets its price $P_{E,t}^E(i)$ for domestic sales so as to maximise future expected discounted profits, i.e. according to

$$\underset{P_{E,t}^E(i)}{max}\, \mathbb{E}_t \sum\nolimits_0^\infty \theta_p^s \Theta_{E,t,t+s} Y_{E,t}^E(i) \left[P_{E,t}^E(i) - MC_{E,t}\right], \tag{4.14}$$

where $\Theta_{E,t,t+s} \equiv \beta^s \frac{\Lambda_{E,t+s}}{\Lambda_{E,t}} \frac{P_{E,t}^c}{P_{E,t+s}^c}$ denotes the household's stochastic discount factor, $Y_{E,t}^E(i)$ demand for domestic sales from the domestic final good bundler in Equation (4.9), and $MC_{E,t}$ real marginal costs.

An EME firm subject to PCP sets the EME currency price of its exports to RoW $\widetilde{P}_{E,t}^F(i)$ according to

$$\underset{\widetilde{P}_{E,t}^F(i)}{max}\, \mathbb{E}_t \sum\nolimits_0^\infty \theta_p^s \Theta_{E,t,t+s} \widetilde{Y}_{E,t}^F(i) \left[\widetilde{P}_{E,t}^F(i) - MC_{E,t}\right], \tag{4.15}$$

where $\widetilde{Y}_{E,t}^F(i)$ is demand for differentiated PCP export goods from the domestic PCP export good bundler defined in Equation (4.21) below. Analogously, an EME firm subject to DCP sets the US dollar price of its exports to RoW $\widehat{P}_{E,t}^F(i)$ according to

$$\underset{\widehat{P}_{E,t}^F(i)}{max}\, \mathbb{E}_t \sum\nolimits_0^\infty \theta_p^s \Theta_{E,t,t+s} \widehat{Y}_{E,t}^F(i) \left[\mathcal{E}_{E,t}^D \widehat{P}_{E,t}^F(i) - MC_{E,t}\right], \tag{4.16}$$

[63] As we discuss below, all model predictions are qualitatively robust to assuming that the non-DCP component of bilateral trade between EME and RoW is subject to LCP instead of PCP or combinations thereof. To save space, we discuss only the equations for the case of PCP. The analogous equations for the case of LCP are provided in the detailed model appendix available from the authors upon request.

[64] A different strand of the literature is concerned with firms' choice of pricing currency, see Corsetti and Pesenti (2004), Devereux et al. (2004), Bacchetta and van Wincoop (2005), Goldberg and Tille (2008), Gopinath et al. (2010), Devereux and Shi (2013), Mukhin (2018), as well as Gopinath and Stein (2018). As in Gopinath et al. (2020b), in this paper we take pricing currency as given. Gopinath et al. (2010) as well as Mukhin (2018) show that firms' currency pricing decision is effectively a zero-one and once-and-for-all decision for a given economic environment that determines the currency in which firms' optimal price is most stable, such as the monetary policy regime, the structure of marginal costs due to domestic wages vs. imported intermediates, whether the export market features competition from local producers or exporters from third economies (see Figure 1 in Mukhin 2018). Importantly, Gopinath et al. (2010) as well as Mukhin (2018) show that firms' choice of pricing currency does in general not change during the adjustment to *temporary* shocks. Hence, as argued also in Gopinath et al. (2020b), we expect the predictions from DCP to be insensitive to endogenising currency choice. Finally, as documented in Gopinath (2015), the invoicing currency shares in the data are rather stable over the sample period we consider.

where $\widehat{Y}^F_{E,t}(i)$ is demand for differentiated DCP export goods from the domestic DCP export good bundler defined in Equation (4.22) below. Equation (4.16) shows that DCP exporters optimise taking into account that their prices are sticky in US dollar.

Before being exported to RoW, EME firms' differentiated PCP and DCP intermediate export goods are combined into a PCP ($\widetilde{Y}^F_{E,t}$) and DCP ($\widehat{Y}^F_{E,t}$) export good bundle, respectively. The PCP and DCP export good bundlers in EME operate under perfect competition using the CES technologies

$$\widetilde{Y}^F_{E,t} = \left[\left(\frac{1}{\gamma^F_E}\right)^{\frac{1}{\psi_i}}\left(\int_0^{\gamma^F_E}\widetilde{Y}^F_{E,t}(i)^{\frac{\psi_i-1}{\psi_i}}di\right)\right]^{\frac{\psi_i}{\psi_i-1}}, \tag{4.17}$$

$$\widehat{Y}^F_{E,t} = \left[\left(\frac{1}{1-\gamma^F_E}\right)^{\frac{1}{\psi_i}}\left(\int_{\gamma^F_E}^{1}\widehat{Y}^F_{E,t}(i)^{\frac{\psi_i-1}{\psi_i}}di\right)\right]^{\frac{\psi_i}{\psi_i-1}}, \tag{4.18}$$

with price indices

$$\widetilde{P}^F_{E,t} = \frac{1}{\gamma^F_E}\left(\int_0^{\gamma^F_E}(\widetilde{P}^F_{E,t}(i))^{1-\psi_d}di\right)^{\frac{1}{1-\psi_i}}, \tag{4.19}$$

$$\widehat{P}^F_{E,t} = \frac{1}{(1-\gamma^F_E)}\left(\int_{\gamma^F_E}^{1}(\widehat{P}^F_{E,t}(i))^{1-\psi_d}di\right)^{\frac{1}{1-\psi_i}}. \tag{4.20}$$

The demand schedule EME firms subject to PCP face for their differentiated exports in Equation (4.15) is given by

$$\widetilde{Y}^F_{E,t}(i) = \gamma^E_F\upsilon^E_F\left(\frac{\widetilde{P}^F_{E,t}(i)}{\widetilde{P}^F_{E,t}}\right)^{-\psi_d}\left(\frac{\mathcal{E}^F_{E,t}\widetilde{P}^F_{E,t}}{P^{F^I}_{E,t}}\right)^{-\psi_i}\left(\frac{P^{F^I}_{E,t}}{P^c_{F,t}}\right)^{-\psi_f}Y^c_{F,t}, \tag{4.21}$$

where we make use of the RoW analogues of Equations (4.7) and (4.12). Equation (4.21) shows that demand for an EME firm's PCP exports depends on the relative price of its differentiated export good and those of other EME PCP exporters reflected in $\widetilde{P}^F_{E,t}$, on the relative price of the aggregate EME PCP export good bundle $\widetilde{P}^F_{E,t}$ and the aggregate import price of total—i.e. PCP and DCP—RoW imports from EME reflected in $P^{F^I}_{E,t}$, as well as on the relative price between the aggregate import price of total RoW imports from EME reflected in $P^{F^I}_{E,t}$ and the RoW aggregate consumer price index $P^c_{F,t}$. Hence, EME PCP exporters compete with all other producers of goods that enter the RoW consumption bundle, including EME DCP exporters, US exporters as well as RoW firms.

Analogously, the demand schedule EME firms subject to DCP face for their differentiated exports in Equation (4.16) from the domestic DCP export good bundler is given by

$$\widehat{Y}^F_{E,t}(i) = (1-\gamma^E_F)\upsilon^E_F\left(\frac{\widehat{P}^F_{E,t}(i)}{\widehat{P}^F_{E,t}}\right)^{-\psi_i}\left(\frac{\mathcal{E}^D_{F,t}\widehat{P}^F_{E,t}}{P^{F^I}_{E,t}}\right)^{-\psi_d}\left(\frac{P^{F^I}_{E,t}}{P^c_{F,t}}\right)^{-\psi_f}Y^c_{F,t}, \tag{4.22}$$

where we again make use of the RoW analogues of Equations (4.7) and (4.12). As in Equation (4.21), Equation (4.22) shows that EME DCP exporters compete with all other producers of goods that enter the RoW consumption bundle, including EME PCP exporters, US exporters as well as RoW firms.

Monetary policy and market clearing

The monetary authority sets the domestic interest rate and follows a Taylor rule

$$\widehat{r}_{E,t} = \rho_r \widehat{r}_{E,t-1} + (1-\rho_r)\left(\phi_\pi \widehat{\pi}^c_{E,t} + \phi_y \widehat{y}_{E,t}\right) + \epsilon^r_{E,t}, \tag{4.23}$$

where $\epsilon^r_{E,t}$ represents a monetary policy shock and $\widehat{y}_{E,t}$ and $\widehat{\pi}^c_{E,t}$ percentage deviations of output from the undistorted (flexible-price) steady state and the steady-state consumer-price inflation, respectively.

Finally, we impose market clearing in all goods, labor and bond markets. We assume bonds are in zero net supply and US bonds are the only bonds traded internationally.

4.2.2 Parametrisation and model solution

Regarding the shares of foreign inputs in the consumption bundle we replicate a world with three country blocks, namely the US, an EME region and the rest of the world. Specifically, in the country sample we study below the average share of economies' exports accounted for by the US is about 12%. Moreover, imports account for about 12% of the US consumption bundle. Hence, we choose the bundle shares reported in Table 4.1. For the other parameter values we follow Gopinath et al. (2020b) and choose the values reported in Table 4.2. We solve the model using a first-order approximation around the non-stochastic steady state.

Table 4.1: Parameter values for bundle shares

EME (E)		
Share of EME consumption bundle accounted for by EME final product	n_E	0.6
Share of EME consumption bundle accounted for by RoW final product	υ_E^F	0.34
Share of EME consumption bundle accounted for by US final product	υ_E^D	0.06
RoW (F)		
Share of RoW consumption bundle accounted for by RoW final product	n_F	0.6
Share of RoW consumption bundle accounted for by EME final product	υ_F^E	0.34
Share of RoW consumption bundle accounted for by US final product	υ_F^D	0.06
US (D)		
Share of US consumption bundle accounted for by US final product	n_D	0.88
Share of US consumption bundle accounted for by EME final product	υ_D^E	0.06
Share of US consumption bundle accounted for by RoW final product	υ_D^F	0.06

Table 4.2: Parameter values

Consumption good bundlers		
Demand elasticity for domestic and foreign final goods	ψ_f	2
Import goods bundler		
Demand elasticity for DCP and PCP import goods	ψ_d	2
Intermediate goods producers		
Calvo probability for prices	θ_p	0.75
Demand elasticity for differentiated intermediate	ψ_i	2
Households		
Risk aversion	σ_c	2
Discount factor	β	0.99
Demand Shock persistence	ρ^β	0.5
Inverse Frisch elasticity of labour	φ	2
Sensitivity of foreign interest rates to NFA	μ	0.001
Central bank		
Central bank smoothing parameter	ρ_R	0.5
Central bank inflation sensitivity	ϕ_π	1.5
Central bank output sensitivity	ϕ_y	0.5

4.2.3 Global spillovers from US demand shocks

To derive model predictions in order to test for partial DCP against combinations of PCP and LCP we consider a multilateral US dollar appreciation induced by a positive US demand shock as the baseline laboratory experiment. One reason for the choice of a US demand shock for the baseline is that these are typically found to account for a substantial fraction of the variation in macroeconomic data (see, for example, Smets and Wouters 2007), so that we expect to exploit a favourable signal-to-noise ratio in our empirical analysis. In contrast, it is typically found that monetary policy shocks only account for a rather small share of the variance of many variables of interest implying a low signal-to-noise ratio in the context of our empirical analysis. For exchange rates the evidence suggests that UIP shocks are particularly important drivers (Engel and West 2010; Itskhoki and Mukhin 2017). However, UIP shocks are also much harder to identify empirically. Moreover, UIP shocks are also not straightforward to motivate from a structural perspective, notwithstanding some recent progress (Itskhoki and Mukhin 2017; Engel and Wu 2018; Jiang et al. 2018); recall for example that we introduce an exogenous exchange rate shock in an ad hoc manner in the model in Equation (4.4). In extensions and robustness checks below we nevertheless do examine the spillovers from US dollar appreciation induced by contractionary US monetary policy and UIP shocks.[65]

In the following we first discuss the transmission of a multilateral US dollar appreciation

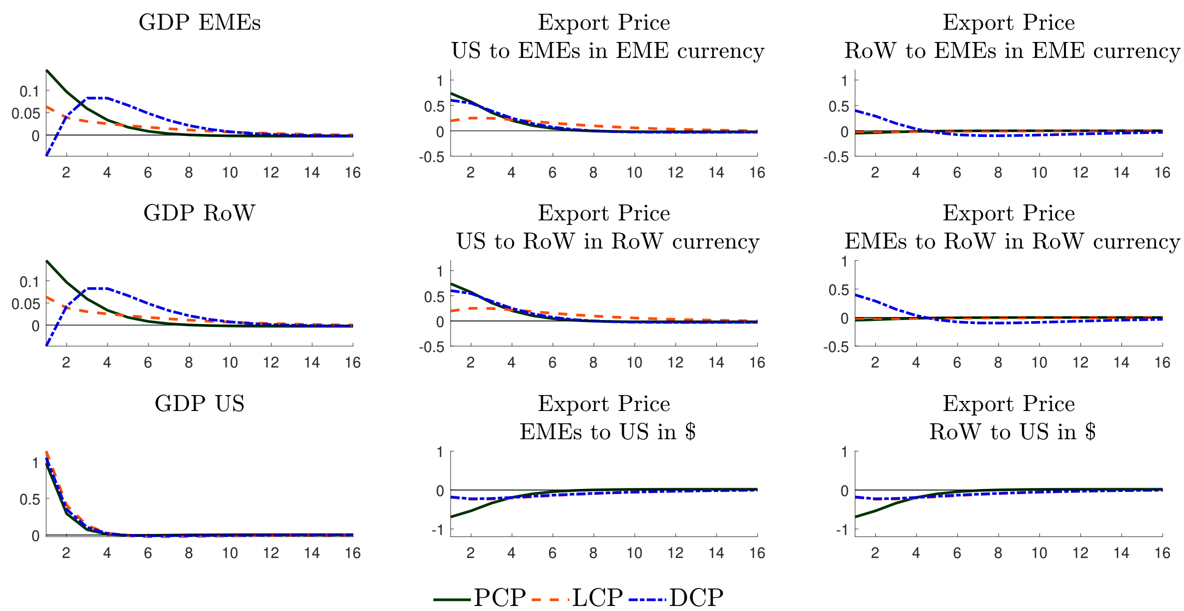
[65] We do not consider cost-push, productivity and fiscal policy shocks as the literature has not reached a consensus on their effects on the exchange rate.

to EME and RoW for the polar cases in which *all* exporting firms in US, EME and RoW are either PCP, LCP or DCP exporters. After having illustrated the transmission channels in these polar cases, we discuss the case of partial DCP, in which only a fraction of exporters in EME and RoW exhibits DCP. and in which the fraction of exports that is subject to DCP may differ from the fraction of imports that is subject to DCP.

Full PCP, LCP or DCP

To establish a reference point we first briefly consider the simple cases of *full* PCP, LCP and DCP. Figure 4.2 presents the dynamic effects of a multilateral US dollar appreciation induced by a positive US demand shock on output and bilateral export prices. Regardless of the pricing paradigm assumed for exporters in US, EME and RoW, the positive US demand shock increases output in the US. Together with the increase in consumer-price inflation this causes US monetary policy to tighten. The resulting interest rate differentials with respect to EME and RoW elicit a multilateral appreciation of the US dollar. Also regardless of the pricing paradigm assumed for exporters in the US, EME and RoW, the rise in US output raises US demand for imports from EME and RoW for a given level of exchange rates and export prices, and hence implies a positive output spillover through a bilateral demand channel. Next we discuss the transmission of the US dollar appreciation to EME and RoW through expenditure switching.

Figure 4.2: Dynamic effects of a positive US demand shock under PCP, LCP and DCP



Note: The figure presents the dynamic effects of a positive US demand shock on real GDP and bilateral export prices denoted in the currency of the importer. The figure provides the impulse responses for three polar parameterisations of the model: All exporters exhibit PCP (black solid lines), LCP (red dashed lines), and DCP (blue dash-dotted lines).

Under full PCP (black solid line), the US dollar appreciation raises local-currency import prices for EME and RoW imports from US, while it lowers local-currency prices of US imports from EME and RoW. Therefore, expenditure switching raises bilateral net exports vis-à-vis US and thus output in EME and RoW. In contrast, under full LCP (red dashed line) local-currency import prices are rather stable. As a result, little expenditure switching occurs and output spillovers from US dollar appreciation are smaller under LCP than under PCP. Finally, notice that both under PCP and LCP the bilateral exchange rate between EME and RoW is stable, as their currencies depreciate symmetrically against the US dollar. As a result, there is no expenditure switching in bilateral trade between EME and RoW in response to multilateral US dollar appreciation.

For the case of full DCP (blue dash-dotted line) it is useful to split the discussion between trade that does and does not involve the US. First, because US import prices are sticky in US dollar as in the case of LCP, local-currency import prices in US remain rather stable. In contrast, EME and RoW local-currency prices of imports from US rise, given that US export prices are sticky in US dollar. As a result, EME and RoW bilateral net exports with US rise. Notice, however, that under DCP the latter rise by less than in case of PCP, because expenditure switching occurs only in EME and RoW but not in US. Second, unlike in the cases of PCP and LCP, under DCP bilateral trade between EME and RoW does respond to US dollar appreciation. Specifically, because prices of EME exports to RoW and vice versa are sticky in US dollar, the latter's appreciation reduces gross bilateral trade between EME and RoW due to expenditure switching in both economies. Note, however, that bilateral net exports between EME and RoW are stable, as their bilateral exports and imports decline to the exact same degree.

Overall, thus, in the case of *full* DCP the output spillovers from the US dollar appreciation to EME and RoW are determined by the response of their bilateral net exports with US. And because bilateral net exports with US rise due to expenditure switching in EME and RoW but not in US, the *peak* output spillovers are in between those obtained under PCP and LCP. The output spillovers in EME and RoW are smaller under DCP than under LCP in the very short term because of different endogenous monetary policy reactions. Specifically, recall that under full DCP all import prices are sticky in US dollar. Therefore, US dollar appreciation has a stronger impact on consumer-price inflation in EME and RoW than in the cases of PCP and LCP, especially because of the large shares of the consumption bundles being accounted for by imports from non-US sources (see Table 4.1), and thereby induces monetary policy to tighten; these dynamics in the very short term are consistent with the mechanisms highlighted by Mukhin (2018) and Zhang (2018).

Partial DCP

We move beyond the polar cases discussed above and in the literature so far and consider *partial* DCP. Specifically, for illustration in the baseline partial DCP parametrisation of the model we assume 50% of bilateral trade between EME and RoW is subject to DCP and PCP, respectively: 50% of EME exports to RoW are priced in US dollar and the remaining

50% in EME currency; analogously, 50% of RoW exports to EME are priced in US dollar and the remaining 50% in RoW currency. Notice that this partial DCP in exports translates into partial DCP in imports: 50% of EME imports from RoW are priced in US dollar, and the remaining 50% in RoW currency, and vice versa. At the same time, and consistent with the findings in Gopinath and Rigobon (2008), we assume trade with US is subject to full DCP, that is all exports to and all imports from US are priced in US dollar.

Figure 4.3 presents the dynamic effects of US dollar appreciation for this baseline partial DCP parametrisation of the model (black solid line). For this baseline case of partial DCP the output spillovers from US dollar appreciation are essentially a combination of those from the cases of full DCP and PCP above. Specifically, given that trade with US is entirely subject to DCP, the bilateral spillovers from US to EME—that is bilateral demand effects from US and expenditure switching in EME away from imports from US towards domestically produced goods—are identical under full and partial DCP. As under full DCP, there is also some expenditure switching between EME and RoW, as 50% of their bilateral exports are subject to DCP. Most importantly, however, as the DCP shares in bilateral exports between EME and RoW are symmetric in this baseline parametrisation of partial DCP, their bilateral exports fall in tandem with their bilateral imports, and hence their bilateral net exports are again unchanged. The non-DCP component of bilateral trade between EME and RoW is unaffected by expenditure switching, because it is subject to PCP and because the bilateral exchange rate between EME and RoW is stable. The overall output spillovers from US dollar appreciation thus again arise exclusively through bilateral spillovers from US, as under the case of full DCP. Notice for later reference that in this symmetric baseline parametrisation of partial DCP, the *non-US export-import DCP share differentials* of EME and RoW are zero, as their shares of exports and imports that are subject to DCP are both 50%.

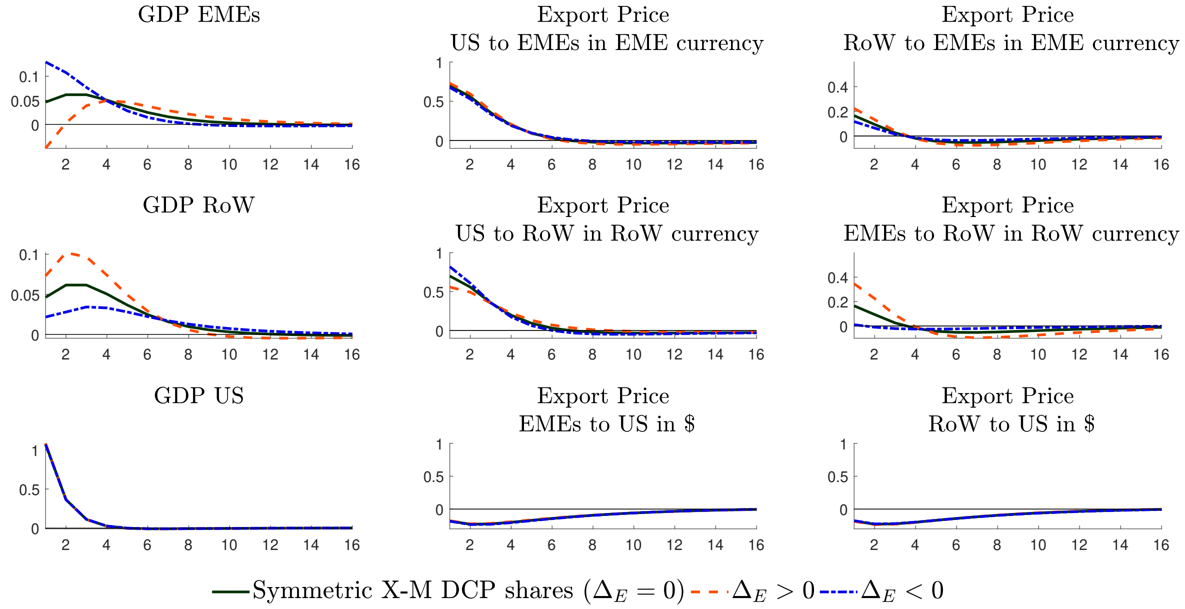
Figure 4.3 also depicts the dynamic effects of US dollar appreciation for the case in which the baseline scenario is altered by raising the share of EME exports to RoW which is subject to DCP to 100% (red dashed line).[66] This parametrisation implies a *positive* EME (and a negative RoW) non-US export-import DCP share differential. In this case, EME bilateral export prices in RoW currency rise by more than RoW bilateral export prices in EME currency. As a result, expenditure switching in bilateral trade is stronger in RoW than in EME, so that EME bilateral net exports with RoW fall and RoW bilateral net exports with EME rise. The fall in EME bilateral net exports with RoW reduces the overall output spillover from US dollar appreciation to EME, in fact even turning it negative for the first few quarters.[67]

Finally, Figure 4.3 also depicts the dynamic effects of US dollar appreciation for the

[66]The findings are the same if we induce variation in the non-US export-import DCP share differentials by varying the import rather than the export DCP share.

[67]The EME currency depreciates somewhat against the RoW currency in this partial DCP parametrisation. Hence, the PCP component of bilateral trade between EME and RoW mitigates somewhat the reduction in the output spillovers to EME.

Figure 4.3: Dynamic effects of a positive US demand shock under partial DCP



Note: The figure presents the dynamic effects of a positive US demand shock on real GDP and bilateral export prices denoted in the currency of the importer for the case of partial DCP. The figure provides the impulse responses for three alternative parameterisations of export pricing in EMEs and the RoW: 50% of EME exports to the RoW are priced in US dollar and vice versa (black solid line, "symmetric X-M DCP shares ($\Delta_E = 0$)"); 100% of EME exports to the RoW and 50% of RoW exports to EMEs are priced in US dollar (red dashed line, $\Delta_E > 0$); 0% of EME exports to the RoW and 50% of RoW exports to EMEs are priced in US dollar (blue dashed-dotted line, $\Delta_E < 0$). In all cases tarde with the US is subject to full DCP, i.e. all exports to and of the US are priced in US dollar.

case in which the baseline scenario is altered by lowering the share of EME exports to RoW which is subject to DCP to 0% (blue dash-dotted line). This partial DCP parametrisation implies a *negative* EME (and a positive RoW) non-US export-import DCP share differential. In this case, EME export prices in RoW currency rise by less than RoW export prices in EME currency. As a result, expenditure switching in bilateral trade is weaker in RoW than in EME, so that EME bilateral net exports with RoW rise and RoW bilateral net exports with EME fall. The rise in EME bilateral net exports with RoW increases the overall output spillovers from US dollar appreciation to EME.

Notice that while—as discussed in Section 4.2.3—the *level* of the output spillovers in the very short term is largely determined by the monetary policy response to the rise in consumer prices also under partial DCP (see Mukhin 2018; Zhang 2018), the *variation* in the output spillovers to EME across different values of the non-US export-import DCP share differentials is due to differences in the response of bilateral net exports with RoW. This can be seen from noting that as the EME non-US export-import DCP share differential falls—for example due to a larger share of imports from RoW priced in US dollar—the output spillovers to EME *rise*, consistent with a rise in EME bilateral net exports driven by a stronger decline in imports from RoW than in exports to RoW. Moreover, larger output spillovers to EME in case of a smaller non-US export-import DCP share differential due to a larger share of imports from RoW priced in US dollar arise *although* EME monetary policy tightens more strongly in the face of a more pronounced rise in import and hence consumer prices. Thus, the variation in the output spillovers to EME across different non-US export-import DCP share differentials is unambiguously due to variation in bilateral net exports and not due to variation in the spillovers to consumer prices and associated monetary policy responses as discussed in Mukhin (2018) and Zhang (2018).[68] Another way to see this is to note that in the scenarios we vary the share of EME exports to RoW that are subject to DCP, while we keep *constant* the share of EME imports from RoW that are subject to DCP.

The key prediction from partial DCP is thus that the output spillovers from multilateral US dollar appreciation are negatively related to economies' non-US export-import DCP share differential. In Appendices D.1 and D.2 we document the robustness of the model prediction from partial DCP in several dimensions. In particular, first, we document by means of a Monte Carlo experiment that the prediction from partial DCP is not specific to the particular parametrisation of the model. We also show that the prediction from partial DCP obtains in case of a contractionary US monetary policy or a UIP shock that induce the multilateral appreciation of the US dollar. Moreover, we document that the prediction from partial DCP also obtains in more elaborate versions of the model. Specifically, we consider a model with trade in intermediate goods used as inputs to production rather than only for consumption. Second, we consider a version of the model with capital, frictions in domestic financial markets, and sticky wages. And finally, we discuss that additional model features such as hedging, non-constant demand elasticity and local distribution services do not change qualitatively the predictions from partial DCP.

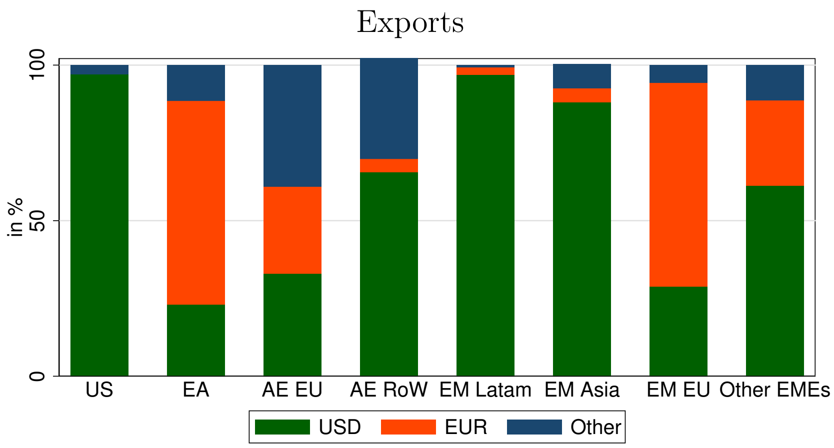
One could argue a more plausible version of the model would feature the US, a small open economy and a large rest of the world block; this would also link more closely to the empirical analysis with two-country VAR models involving the US and a non-US economy below. Notice, however, that in the context of our paper such a specification would not account for the role of DCP in trade *between* the economies that make up the rest of the world in the data, as the small open economy would be of infinitesimal size relative to the rest of the world block in the model. Eventually, that would blur the exposition of the mechanisms we are interested in. Specifically, when the US dollar appreciates multilaterally import prices in the small open economy would soar given its large degree of openness, inducing a substantial tightening in monetary policy. In contrast, prices of imports from the small open economy in the rest of the world block would be stable, given its size and hence limited degree of openness. As a result, the small open economy's currency would appreciate

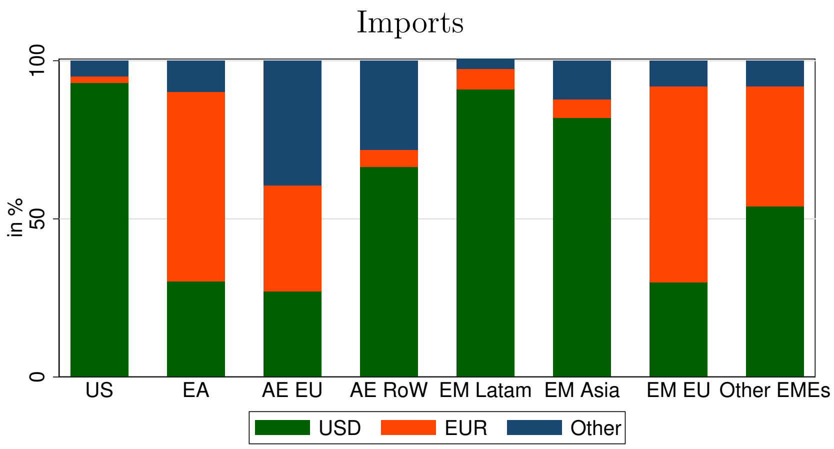
[68] Interestingly, we do find empirical evidence that is consistent with the variation in spillovers to *consumer prices* through the mechanism discussed in Mukhin (2018) and Zhang (2018): The spillovers from US dollar appreciation induced by a US monetary policy shock to consumer prices and short-term interest rates in the rest of the world is larger for economies which have a larger share of imports invoiced in US dollar. That these larger consumer-price and short-term interest rate spillovers do not translate into larger—i.e. more *negative*—output spillovers suggests that the mechanism arising through monetary policy responding with a tightening to the rise in inflation is weaker than the expenditure switching/net exports channel. Our findings for the estimates of the spillovers from US monetary policy to consumer prices are available on request.

against that of the rest of the world, inducing expenditure switching also in their bilateral non-DCP trade. Moreover, the small open economy's currency would depreciate less against the US dollar, mitigating expenditure switching in DCP trade with the rest of the world block. Both effects would add transmission channels that contribute to the overall effects of US dollar appreciation in the small open economy, and thereby blur the exposition of the role of DCP. Most importantly, however, such a scenario would be empirically implausible, because it would not account for the fact that the rest of the world is an aggregate of small open economies which trade with each other subject to DCP. And this means that import prices in each individual small open economy in the rest of the world would soar as much as in the small open economy considered specifically, contrary to the implications of a model with a large rest of the world block. A possible solution would be to specify that a share of domestic sales in the rest of the world block are priced in US dollar. Our view is that such a model would feature several additional layers of expositional complexity without important conceptual gains relative to the model we consider.

In the next section we confront the model prediction with the data. Before doing so, we however have to amend the hypothesis we intend to test. In particular, dominant-currency *pricing* shares—that is the share of export and import prices of non-US trade that is sticky in US dollar—are not observed in the data; and, in fact, if they were, there would not be any need to test for the empirical relevance of DCP in the first place. What we do observe are export and import *invoicing* rather than pricing shares (Figure 4.4). However, the currency of *invoicing* does not need to coincide with the currency in which trade prices are *sticky*, if prices are sticky at all; a case in point are commodity exports, which are typically invoiced in US dollar but whose prices are believed to not be sticky in US dollar (or any other currency). Some evidence for a few individual countries suggests that the currency in which trade prices are sticky is closely related to the currency of invoicing, see Gopinath and Rigobon (2008) for US imports and exports, Fitzgerald and Haller (2014) for Irish exports to the UK as well as Friberg and Wilander (2008) for Swedish exports. However, to the best of our knowledge, there is no evidence for trade between EMEs, for which DCP may be particularly relevant. Therefore, in the next section we test the *joint* hypothesis that the data are characterised by partial DCP *and* that the currency of invoicing coincides with the currency in which trade prices are sticky. The associated alternative hypothesis is that *either* the data are not characterised by partial DCP *or* that the currency of invoicing does not coincide with the currency in which trade prices are sticky. Thus, if the evidence was such as to reject the null hypothesis, this may be the case even if partial DCP is empirically relevant, namely in case the invoicing currency does not coincide with the currency in which trade prices are sticky. If anything, thus, incorporating the assumption that the currency in which trade prices are sticky coincides with the currency of trade invoicing stacks the deck against finding evidence for partial DCP.

Figure 4.4: Global trade invoicing landscape





Note: The figure presents the shares of economies' exports and imports invoiced in US dollar, Euro and other currencies. Data for euro area countries includes intra-euro area trade. The data are taken from Gopinath (2015).

4.3 Empirical evidence

4.3.1 Estimating the spillovers from US dollar appreciation

We estimate spillovers from US shocks using two-country VAR models involving the US and one spillover-recipient economy at a time. Specifically, consider a two-country VAR model

$$\boldsymbol{x}_t = \boldsymbol{A}(L)\boldsymbol{x}_{t-1} + \boldsymbol{P}\boldsymbol{\epsilon}_t, \qquad \boldsymbol{\epsilon}_t \sim (\boldsymbol{0}, \boldsymbol{I}), \tag{4.24}$$

with

$$\boldsymbol{x}_t \equiv \begin{bmatrix} \boldsymbol{x}_{us,t} \\ s_t^{usd} \\ \boldsymbol{x}_{it} \end{bmatrix}, \qquad \boldsymbol{x}_{jt} \equiv \begin{bmatrix} y_{jt} \\ \pi_{jt} \\ i_{jt} \end{bmatrix}, \; j \in \{us, i\}, \tag{4.25}$$

and where i indexes non-US economies, y_{it} denotes real GDP, π_{it} consumer-price inflation, i_{it} the nominal short-term interest rate, and s_t^{usd} the nominal effective US dollar exchange rate. We estimate the VAR models in Equation (4.24) for up to 45 economies for the time period from—at most, given data availability—1995q1 to 2018q3. The set of economies we consider is reported in Table 4.3 and covers around 57% of non-US global GDP. We only consider economies for which we also have data on US dollar invoicing shares.[69] We combine data from the updated GVAR toolbox (see Smith and Galesi 2011), the IMF's International Financial Statistics (IFS) as well as national sources accessed through Haver. We use data on shadow rates of Wu and Xia (2016) for the US, the euro area economies and the UK, as well as from Krippner (2013) for Japan.[70] We include four lags of $\boldsymbol{x}_t$ in Equation (4.24). We consider Bayesian estimation with flat normal diffuse priors. In robustness checks below we consider alternative specifications of the VAR models as regards *inter alia* the priors used in Bayesian estimation, the choice of endogenous variables, accounting for country-specific, regional and global crises, the sample period, and imposing a small open economy assumption.

In our baseline we impose standard sign restrictions to identify US demand, monetary policy, supply and UIP shocks in $\boldsymbol{\epsilon}_t$ using the algorithm of Arias et al. (2018). In particular, consistent with the model in Section 4.2, we impose the restrictions that in case of a positive US demand shock US real GDP, consumer-price inflation and the nominal short-term interest rate rise, and that the nominal effective exchange rate of the US dollar appreciates.[71] In

[69]Despite the availability of export and import currency invoicing share data in Gopinath (2015), in our baseline we do not consider Pakistan due to lack of data on quarterly real GDP. We also do not consider economies with a population of less than one million in our baseline, i.e. Iceland and Luxembourg. Finally, we do not consider Algeria in our baseline, as 99% of its exports are accounted for by oil. However, we document in the robustness checks below that our results do not change when we include these countries. China is not included because there is no invoicing data. Greece is not included because the VAR model featured explosive dynamics.

[70]For euro area economies we combine national with euro area interest rate data to obtain a time series for the pre and post-euro periods.

[71]Note that a positive *global* demand shock mitigates investors' risk aversion and thereby *depreciates* the US dollar. The early phase of the COVID pandemic is a case in point for US dollar appreciation in

Table 4.3: Economies included

Advanced	AUS, AUT, BEL, CAN, CHE, CYP, DEU, DNK, ESP, FIN, FRA, GBR, *ISL*, IRL, ITA, JPN, *LUX*, NLD, NOR, PRT, SWE
EM Europe	BGR, CZE, EST, HRV, HUN, LTU, LVA, POL, ROU, SVK, SVN, UKR
EM Asia	IDN, IND, KOR, *PAK*, THA
EM Latin America	ARG, COL, BRA, PER
EM Middle East and Africa	*DZA*, ISR, MAR, TUR

case of a contractionary US monetary policy shock we impose the restrictions that US real GDP and consumer-price inflation fall, that the nominal short-term interest rate rises, and that the US dollar nominal effective exchange rate appreciates. In case of a positive supply shock we impose the restrictions that US real GDP falls, that consumer-price inflation rises, and that the nominal short-term interest rate rises. Finally, and consistent with the model in Section 4.2, we also identify a UIP shock imposing the restrictions that it appreciates the US dollar, that it lowers US interest rates, real GDP and consumer-price inflation; as mentioned before, identification of the UIP shock is not consequential for our results.[72] We impose all sign restrictions on impact. While we again focus on the output spillovers from a US demand shock in the baseline, we consider the spillovers from US monetary policy and UIP shocks in extensions below. Moreover, in robustness checks below instead of sign restrictions we consider identification using external instrumental variables for all these shocks (Stock and Watson 2012; Mertens and Ravn 2013).

Figure 4.5 reports the estimates of the output spillovers from a positive US demand shock at the country level. Specifically, Figure 4.5 reports the average of the posterior medians of the response of real GDP over the first eight quarters. Dark-shaded bars indicate EMEs, and light-shaded bars AEs. The estimates suggest that there are pronounced cross-country heterogeneities in the output spillovers. Most economies benefit from a positive US demand shock in terms of output, but the spillovers are negative in several cases. In fact, among EMEs, with the exception of Israel and Turkey, spillovers from a positive US demand shock are positive only for European EMEs, while they are negative to non-European EMEs as well as for Ukraine.

response to a negative global demand shock.

[72] One might argue that identifying additional shocks via sign restrictions sharpens the inference for the shock in question. However, notice that the columns of the rotation matrices are drawn sequentially and independently from each other except for the orthogonality restriction in the algorithm of Arias et al. (2018). See also Rubio-Ramirez et al. (2010), Uhlig (2017) as well as Giacomini et al. (2022).

Figure 4.5: Estimated GDP responses to a positive US demand shock

Note: The figure shows the averages over eight quarters of the point estimates of the real GDP effects of a positive US demand shock. The spillover estimates are obtained from estimation of the two-country VAR models described in Section 4.3.1.

4.3.2 Relationship between spillovers from US dollar appreciation and export-import US dollar invoicing share differentials

Data on US dollar export and import invoicing

In order to obtain economies' non-US export-import US dollar invoicing share differentials we combine data from various sources. In particular, our baseline source for invoicing data is Gopinath (2015), who in turn builds and improves on the data collected by Kamps (2006), Goldberg and Tille (2008), as well as Ito and Chinn (2014). We combine the latter with data collected by the ECB in collaboration with national statistical authorities for several euro area and EU economies, data from Devereux et al. (2017) for Canada, and from Ito et al. (2016) for Japan. Second, to the extent possible we combine invoicing data from various sources and calculate time averages of economies' invoicing shares for the same sample periods for which we estimate the corresponding VAR models. Third, starting from economies' *total* export and import US dollar invoicing shares we construct *non-US* export and import US dollar invoicing shares. To do so, we use, to the extent available, information on the share of an economy's exports to and imports from the US that is invoiced in US dollar from Gopinath (2015); if this information is not available, we assume that all of an economy's exports to and imports from the US are invoiced in US dollar.[73]

Figure 4.6 displays the resulting shares of non-US exports and imports invoiced in US dollar as well as the non-US export-import US dollar invoicing share differentials. The

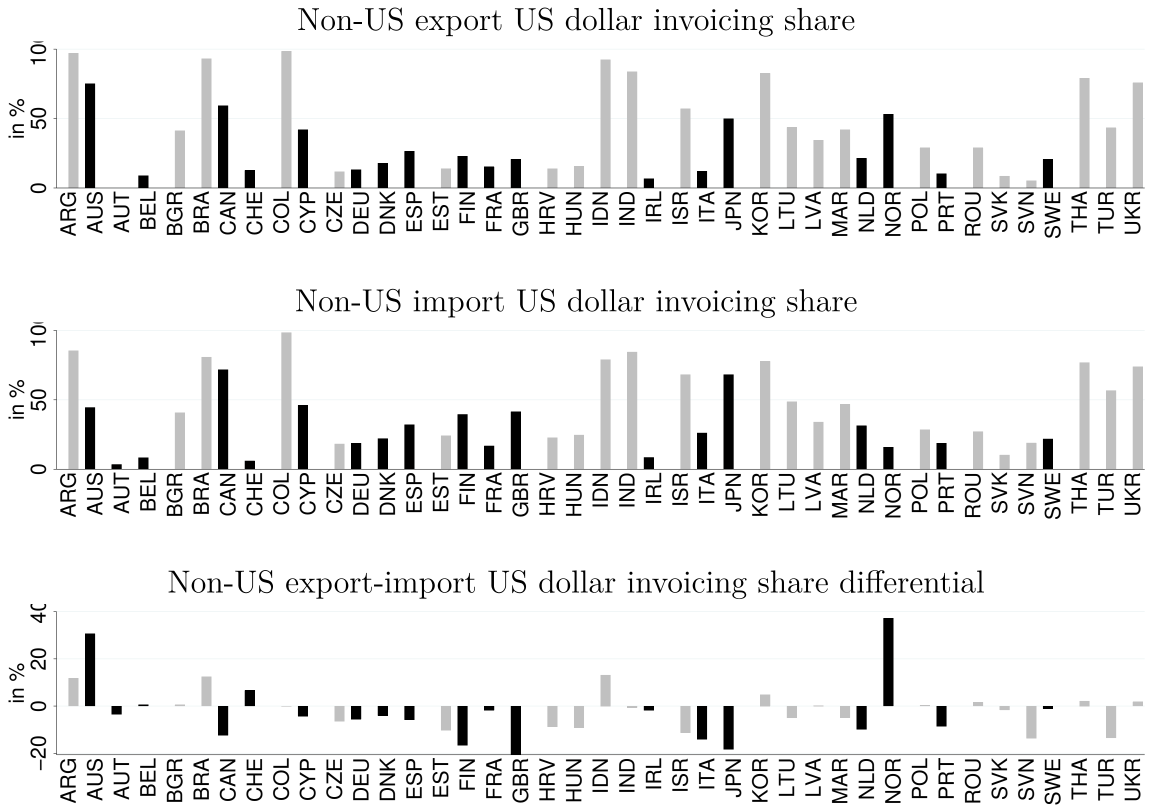
[73]In particular, the non-US export-import US dollar invoicing share differential is given by

$$\begin{aligned}
\Delta_{i,nonus}^{us\$} &\equiv \frac{X_{i,nonus}^{us\$}}{X_{i,nonus}} - \frac{M_{i,nonus}^{us\$}}{M_{i,nonus}} = \frac{X_i^{us\$} - X_{i,us}^{us\$}}{X_{i,nonus}} - \frac{M_i^{us\$} - M_{i,us}^{us\$}}{M_{i,nonus}} \\
&= \frac{X_i^{us\$} - X_{i,us}^{us\$}}{X_i} \cdot \frac{X_i}{X_{i,nonus}} - \frac{M_i^{us\$} - M_{i,us}^{us\$}}{M_i} \cdot \frac{M_i}{M_{i,nonus}} \\
&= \left(\delta_i^{x,us\$} - \frac{X_{i,us}^{us\$}}{X_{i,us}} \cdot \frac{X_{i,us}}{X_i}\right) \cdot \frac{1}{1 - s_i^{x,us}} - \left(\delta_i^{m,us\$} - \frac{M_{i,us}^{us\$}}{M_{i,us}} \cdot \frac{M_{i,us}}{M_i}\right) \cdot \frac{1}{1 - s_i^{m,us}} \\
&= \left(\delta_i^{x,us\$} - \delta_{i,us}^{x,us\$} \cdot s_i^{x,us}\right) \cdot \frac{1}{1 - s_i^{x,us}} - \left(\delta_i^{m,us\$} - \delta_{i,us}^{m,us\$} \cdot s_i^{m,us}\right) \cdot \frac{1}{1 - s_i^{m,us}}, \qquad (4.26)
\end{aligned}$$

where $\delta_i^{\ell,us\$}$ represents economy i's share of total exports/imports invoiced in US dollar, $\delta_{i,us}^{\ell,us\$}$ represents economy i's share of exports/imports in trade with the US invoiced in US dollar, and $s_i^{\ell,us}$ represents the share of exports to/imports from the US in economy i's total exports/imports.

data again exhibit substantial cross-country heterogeneity. Large and/or positive non-US export-import US dollar invoicing share differentials are typically observed for commodity exporters such as Argentina, Australia, Brazil, Indonesia, or Norway, an issue to which we return in more detail below.

Figure 4.6: US dollar invoicing shares



Note: The figure presents the time averages of non-US export and import invoicing shares and the implied differential we use in the empirical analysis in Section 4.3. To the extent possible given data availability, the averages are calculated over the same time period for which the corresponding two-country VAR models are estimated. The data are taken from Gopinath (2015), the ECB, Devereux et al. (2017) as well as Ito et al. (2016), see Section 4.3.2 for more details. Data for euro area economies includes intra-euro area trade.

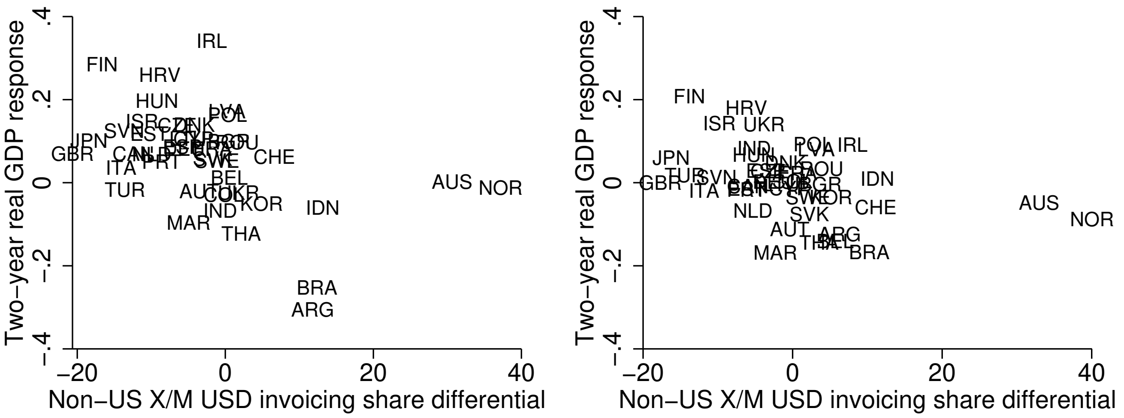
Unconditional relationship between US dollar invoicing share differentials and spillovers from US dollar appreciation

The left-hand side panel in Figure 4.7 displays the unconditional correlation between the output spillovers and the non-US export-import US dollar invoicing share differential of spillover-recipient economies at the country level.[74] Consistent with the prediction of partial DCP in the structural model, the correlation between the output spillovers from US dollar appreciation and economies' non-US export-import US dollar invoicing share differentials is negative: The output spillovers are smaller for economies which feature a larger share of

[74]All of our results are robust to considering the difference between the shares of an economy's share of non-US exports and imports invoiced in US dollar *scaled by GDP*, thereby accounting for differences in the total value of exports and imports.

non-US exports than imports invoiced in US dollar; the negative relationship is also highly statistically significant.

Figure 4.7: Unconditional and conditional scatterplot for the relationship between spillovers from a positive US demand shock and non-US export/import US dollar invoicing share differential



Note: The figure depicts the correlation between the estimated spillovers from a positive US demand shock and economies' non-US export-import US dollar invoicing share differential. The spillover estimates are obtained from the two-country VAR models and are give by the average over the first eight quarters after the US demand shock has occurred. The left-hand side panel depicts the unconditional correlation, and the right-hand side panel the conditional correlation obtained after controlling for the other explanatory variables included in the regression in Equation (4.27), that is spillover-recipient economies' overall non-US trade integration measured by the ratio of the sum of imports and exports to GDP taken from the World Bank's World Development Indicators (WDI); the ratio of spillover-recipient economies' bilateral trade with the US to GDP calculated based on data from the IMF's Direction of Trade statistics; and spillover-recipient economies' exchange rate flexibility using the data from Ilzetzki et al. (2019).

Regression analysis controlling for other transmission channels and cross-country heterogeneity

The transmission channel for the output spillovers from US dollar appreciation that we highlight in the structural model in Section 4.2 focuses on expenditure switching between economies in the rest of the world. However, even in the model the spillovers in general materialise through additional channels. While these additional channels are irrelevant for the variation in the spillovers that we obtain when varying the export-import DCP share differential in the model because they are held constant across scenarios, this is not the case for the sample of economies we consider in the empirical analysis: In the data, economies do not only differ in their non-US export-import US dollar invoicing share differentials, but also in many dimensions that imply differential susceptibility to these additional transmission channels. We thus run regressions in which we control for economies' differential susceptibility to additional transmission channels of spillovers from US dollar appreciation.

In the baseline regression specification we control only for transmission channels that are present in the structural model in Section 4.2. First, recall that a positive US demand shock raises US demand for imports for a given level of the exchange rate and export prices.

Second, the multilateral appreciation of the US dollar gives rise to expenditure switching effects. Because of DCP of exports of the rest of the world to the US, the appreciation of the US dollar is inconsequential in terms of expenditure switching. In contrast, PCP of exports of the US to the rest of the world implies that US dollar appreciation elicits expenditure switching away from imports from the US towards domestically produced goods, and hence positive output spillovers. Third, in order for the latter bilateral and the multilateral DCP expenditure switching effects to unfold, economies' exchange rates must be flexible in the first place. Hence, exchange rate flexibility raises the output spillovers from a positive US demand shock that appreciates the US dollar through expenditure switching.[75] Finally, conditional on the bilateral demand effects emanating from the US as well as the bilateral and multilateral expenditure switching, economies which are more open to trade overall benefit more from the latter through higher-order spillovers in general equilibrium. To summarise, we control for economies' exposure to bilateral trade with the US, multilateral trade with the rest of the world, and their exchange rate flexibility.

More specifically, we run the regression

$$\Delta\widehat{y}_i = \alpha + \beta \cdot \delta_i + \boldsymbol{\gamma}'\boldsymbol{z}_i + \eta_i, \tag{4.27}$$

where $\Delta\widehat{y}_i$ denotes the estimated average real GDP spillover from a positive US demand shock over the first eight quarters depicted in Figure 4.5, and δ_i the non-US export-import US dollar invoicing share differential depicted in the bottom row in Figure 4.6; $\boldsymbol{z}_i$ includes spillover-recipient economies' overall non-US trade openness measured by the ratio of the sum of non-US imports and exports to GDP using data from the World Bank's World Development Indicators (WDI) and the IMF Direction of Trade Statistics (DoTS); the ratio of spillover-recipient economies' bilateral trade with the US to GDP based on data from the IMF DoTS; and spillover-recipient economies' exchange rate flexibility using the data from Ilzetzki et al. (2019). In the robustness checks below we report results from regressions in which we consider alternative definitions of the dependent variable (such as the horizon over which the spillover is defined), additional controls that account for other cross-border transmission channels, alternative country samples, GDP-weighted and robust regressions. Table 4.4 reports summary statistics of the data we use in the estimation of Equation (4.27).[76]

[75]The same applies to the exchange rates of economies' trading partners. We do not control for trading-partner exchange rate flexibility, however, because there is very little cross-country variation. From every individual economy's perspective, trading-partner exchange rate flexibility is an average over the same set of economies except for two, and hence very similar across economies.

[76]Equation (4.27) is not subject to classical measurement error as it is the dependent and not the explanatory variable that is obtained from a previous estimation. In particular, the measurement error in the dependent variable does not imply omitted variable bias due to correlation between the composite regression error term and the observed explanatory variables as in a classical measurement error setting. Instead, measurement error in the dependent variable only inflates the variance of the composite regression error. To the extent that the measurement error in the dependent variable is different across observations, inference needs to account for heteroskedasticity to be efficient. See Lewis and Linzer (2005) for a discussion.

Table 4.4: Summary statistics

	Mean	Min	p5	p95	Max	SD	count
Two-year real GDP response to US demand shock	0.06	-0.31	-0.12	0.26	0.34	0.12	41
Non-US trade rel. to GDP	78.38	19.84	21.93	138.34	144.73	37.38	41
Trade with US rel. to GDP	6.54	1.46	1.78	16.08	47.80	8.40	41
Exchange rate flexibility against USD	3.63	1.58	2.15	4.00	4.00	0.67	41
Non-US exports USD invoicing share	38.52	0.00	6.78	93.23	98.31	29.81	41
Non-US imports USD invoicing share	40.74	3.53	8.43	84.51	98.52	26.45	41
Non-US X/M USD invoicing share differential	-2.22	-20.64	-16.66	13.16	37.20	11.42	41

Table 4.5: Results of regressions of real GDP spillovers from a positive US demand shock

	(1) Baseline	(2) Baseline	(3) Standar-dised	(4) Invoicing rel. to GDP
Non-US exports USD invoicing share	-0.003***			
	(0.00)			
Non-US imports USD invoicing share	0.003**			
	(0.03)			
Non-US X/M USD invoicing share differential		-0.003***	-0.292***	
		(0.00)	(0.00)	
Non-US X/M USD invoicing share differential rel. to GDP				-0.008***
				(0.00)
Non-US trade rel. to GDP	0.001	0.001*	0.231*	0.001
	(0.10)	(0.06)	(0.06)	(0.11)
Trade with US rel. to GDP	0.004**	0.004**	0.240**	0.004**
	(0.04)	(0.03)	(0.03)	(0.04)
Exchange rate flexibility against USD	0.084*	0.096***	0.516***	0.103***
	(0.05)	(0.00)	(0.00)	(0.00)
Adjusted R-squared	0.52	0.53	0.53	0.53
Observations	41	41	41	41

Table 4.6: Regression results for export and import responses to a positive US demand shock

	(1) Real GDP	(2) Exports	(3) Imports
Non-US X/M USD invoicing share differential	-0.003***		
	(0.00)		
Non-US exports USD invoicing share		-0.004***	
		(0.00)	
Non-US imports USD invoicing share			-0.005*
			(0.08)
Non-US trade rel. to GDP	0.001*	-0.000	
	(0.06)	(0.67)	
Trade with US rel. to GDP	0.004**	0.007***	
	(0.03)	(0.00)	
Exchange rate flexibility against USD	0.096***		0.002
	(0.00)		(0.99)
Adjusted R-squared	0.53	0.30	0.15
Observations	41	42	41

Table 4.5 reports the results of the estimation of the baseline specification of Equation (4.27). In column (1) we include the share of economies' non-US exports and imports invoiced in US dollar separately. Both coefficient estimates are highly statistically significant and have the expected signs. Specifically, a larger share of exports invoiced in US dollar is associated with smaller output spillovers from US dollar appreciation; in contrast, a larger share of imports invoiced in US dollar is associated with larger—i.e. more positive or less negative—output spillovers from US dollar appreciation. Column (2) reports the regression results for the specification with the non-US export-import US dollar invoicing share differential. The right-hand side panel in Figure 4.7 presents the conditional correlation—i.e. the slope estimate $\widehat{\beta}$ reported in column (2)—after controlling for the variables in $\boldsymbol{z}_i$. The results document that the coefficient estimate of the non-US export-import US dollar invoicing share differential reported in column (2) in Table 4.5 is not driven by outliers. Column (3) in Table 4.5 reproduces the results from column (2) with the difference that the regression is estimated on standardised data. The results document that varying the non-US export-import US dollar invoicing share differential has an economically large impact on the output spillovers from US dollar appreciation. Overall, our findings are consistent with the model prediction and hence provide evidence for the empirical relevance of partial DCP.

We next consider several extensions to the baseline specification, including zooming in on the responses of economies' exports and imports, a closer examination of the role of commodity trade in US dollar invoicing, considering the effects of US dollar appreciation induced by a contractionary US monetary policy and a UIP rather than a positive US demand shock, and the (lack of) evidence for partial DCP obtained from considering the relationship between spillovers from a euro area demand shock and economies' export-import euro invoicing share differential.

4.3.3 The responses of exports and imports to US dollar appreciation

The mechanism underlying the output spillovers from US dollar appreciation is based on differential responses of exports and imports driven by differences in the importance of US dollar pricing in each. Hence, in principle we could instead of focusing on output estimate the responses of exports and imports, and correlate those with the corresponding export and import US dollar invoicing shares, respectively. In our baseline, we however focus on output for various reasons. First, quarterly data on real GDP are available for a larger number of economies than data for real exports and imports. Second, the output spillover is a parsimonious summary statistic for the mechanism we want to test for in the data. Third, overall real activity is an eventually more important variable, at least from a welfare and policy perspective. Nevertheless, exploring exports and imports separately is a useful exercise that could corroborate our evidence for partial DCP.

Table 4.6 reports results from regressions in which we replace the dependent variable in

Equation (4.27) by the response of economies' real exports and imports to the positive US demand shock that appreciates the US dollar.[77] As controls we only consider variables that we expect to be relevant for exports and imports, respectively.

The results in column (2) suggest that economies with a larger share of non-US exports being invoiced in US dollar experience a stronger drop in their exports in response to US dollar appreciation, consistent with partial DCP.[78] In turn, and again consistent with partial DCP, the results in column (3) suggest that economies with a larger share of non-US imports invoiced in US dollar also experience a stronger drop in their imports. The estimates for the regressions for imports are less precise than for exports, which might be because of the presence of countervailing forces. In particular, as imports decline due the increase in their home-currency price in case of US dollar invoicing, expenditure switching shifts demand towards domestically produced goods, which stimulates production and income and hence again raises demand for imports; admittedly a second-round effect. However, when we additionally control for economies' output spillover in column (4) the estimate on the share of non-US imports invoiced in US dollar becomes much more precise.[79] Overall, the analysis of the responses of exports and imports to US dollar appreciation corroborate our baseline results for the relevance of partial DCP based on output spillovers.

4.3.4 The role of commodity trade

One might claim that the transactions invoiced in US dollar captured by the invoicing share data reported in Gopinath (2015) are primarily related to commodity exports and imports. If this claim was true this would beg two questions. First, given our estimates are consistent, are the DCP effects we find also relevant for non-commodities trade, or are they driven only by commodity trade? Second, one might argue that even though commodities are invoiced in US dollar their prices are not *sticky* in US dollar. If so, we should expect to find stronger evidence consistent with the predictions from partial DCP if we focus on export-import US dollar invoicing share differentials relating only to non-commodity trade? While these two questions are not entirely mutually consistent, it is important to ensure that our results are not unduly specific to or driven by commodity trade. Note that in doing so, we neither take a stand on whether commodities are invoiced in US dollars nor whether their prices are less sticky.

We consider two exercises to account for the possible role of commodity trade in US dollar invoicing. First, we additionally control for the share of economies' commodity exports and

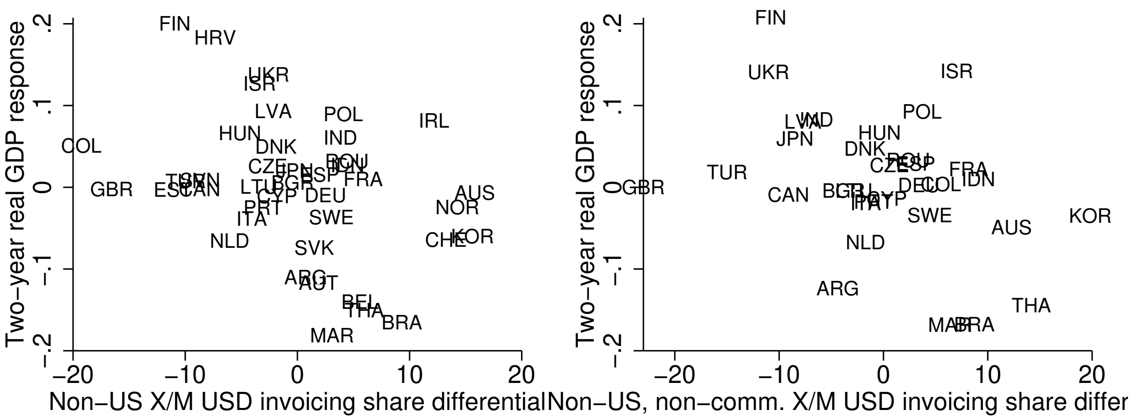
[77] We obtain these responses from extensions of the baseline two-country VAR models in which we append one at a time real exports and imports to the vector of endogenous variables. Because many of the time series for the trade variables are shorter than for the remaining variables in the VAR, we reduce the lag order of the VAR models to one. Quarterly data on real trade variables are not available for Morocco.

[78] Invoicing data is available only for exports in case of Malaysia and South Africa, and only for imports in case of Peru. These economies are not included in the baseline regressions using the invoicing share differentials, but included in the regressions using only the export or import invoicing shares.

[79] Notice that in order to avoid simultaneity problems in column (4) we control for the *one*-year output spillover and consider the *three*-year response of real imports.

imports in total trade in the regression in Equation (4.27). Specifically, we consider the sum of the shares of “agricultural raw materials exports/imports", “fuel exports/imports" and “ores and metals exports/imports" in total exports/imports constructed in the World Bank’s WDI and based on United Nations COMTRADE data.[80] Column (2) in Table 4.7 reports the corresponding regression results, and the left-hand side panel in Figure 4.8 presents the conditional correlation analogous to the right-hand side panel in Figure 4.7. The coefficient estimate on the non-US export-import US dollar invoicing share differential continues to be negative and statistically significant. Another noteworthy finding is that in the conditional correlation plot Australia and Norway—two prominent commodity exporters—do not stand out anymore as outliers.

Figure 4.8: Conditional scatterplots for the relationship between spillovers from a positive US demand shock and country characteristics accounting for commodity exports and imports



Note: The left-hand side panel depicts the conditional correlation between the estimated spillovers from a positive US demand shock and control variables included in the regression in Equation (4.27) as well as economies’ commodity export-import share differential. The right-hand side panel presents the conditional scatterplot from a regression in which the non-US, non-commodity US dollar export-import invoicing share differential is used.

Second, instead of the US dollar invoicing share differential related to economies’ *total* non-US trade, we consider the non-US, *non-commodity* export-import US dollar invoicing share differential. Given that we do not have information on the invoicing currency of commodity trade, we assume that 100% is invoiced in US dollar.[81] Unfortunately, with the latter assumption we obtain that for a few economies the implied non-US, non-commodity

[80] Agricultural raw materials comprise SITC section 2 (crude materials except fuels) excluding divisions 22, 27 (crude fertilizers and minerals excluding coal, petroleum, and precious stones), and 28 (metalliferous ores and scrap); ores and metals comprise the commodities in SITC sections 27 (crude fertilizer, minerals); 28 (metalliferous ores, scrap); and 68 (non-ferrous metals); and fuels comprise the commodities in SITC section 3 (mineral fuels, lubricants and related materials). Our results hardly change when we identify commodities as in Gopinath et al. (2020b) as trade of goods in HS chapters 1-27 and 72-83, or if we apply the classification of traded goods that are sufficiently homogeneous to feature organised markets or reference prices at the three-digit SITC level proposed by Rauch (1999).

[81] The non-US, *non-commodity* export US dollar invoicing share is given by

$$\Delta^{us\$}_{i,nonus,noncom} = \left[\delta_i^{x,us\$} - \delta_i^{x,com} - s_i^{x,us}\delta_{i,us}^{x,us\$}\left(1-\delta_i^{x,us,com}\right)\right]\cdot\frac{1}{\left(1-\delta_i^{x,nonus,com\$}\right)\left(1-s_i^{x,us}\right)}, \tag{4.28}$$

Table 4.7: Results of regressions of real GDP spillovers from a positive US demand shock when controlling for commodity export and import shares

	(1)	(2)	(3)
Non-US X/M USD invoicing share differential	-0.003***	-0.004**	
	(0.00)	(0.02)	
Commodity export/import share differential		0.001	
		(0.47)	
Non-US, non-comm. X/M USD invoicing share differential			-0.004**
			(0.04)
Non-US trade rel. to GDP	0.001*	0.001**	0.001**
	(0.06)	(0.05)	(0.02)
Trade with US rel. to GDP	0.004**	0.003	-0.002
	(0.03)	(0.12)	(0.52)
Exchange rate flexibility against USD	0.096***	0.090***	0.071**
	(0.00)	(0.01)	(0.03)
Adjusted R-squared	0.53	0.52	0.55
Observations	41	41	32

export and import US dollar invoicing shares are negative; this is most plausibly because only less than 100% of the goods we classify as commodities based on the WDI data are invoiced in US dollar. We hence drop these economies for this exercise. Column (3) in Table 4.7 reports the regression results, and the right-hand side panel in Figure 4.8 presents the corresponding conditional correlation. Again, the results are very similar to the baseline, even though the estimates are somewhat less precise which is however not surprising given the substantial reduction in the size of the sample. Finally, it is again noteworthy that Australia and Norway do not stand out anymore as outliers in the conditional correlation plot. Overall, these results suggest that the evidence for partial DCP is not driven by US dollar invoicing of commodity trade (which would also be hard to rationalise given that it is widely believed that commodity prices are—while quoted—not sticky in US dollar).

4.3.5 US monetary policy and UIP shocks

Figures D.3 and D.4 in Appendix D.2 document that the predictions from partial DCP that we discuss in the context of a positive US demand shock in the structural model in Section 4.2 also obtain in case of a contractionary US monetary policy and a UIP shock. In short, the reason for this is that the US dollar appreciates multilaterally in all cases. Specifically, while bilateral demand effects emanating from the US are negative in case of a

where $\delta_i^{x,com}$ represents economy i's share of commodity exports in total exports, $\delta_i^{x,\ell,com}$ the share of commodity exports in total bilateral exports to economy ℓ, where

$$\delta_i^{x,nonus,com} = \left(\delta_i^{x,com} - \delta_i^{x,us,com} s_i^{x,us}\right) \cdot \frac{1}{1 - s_i^{x,us}}. \tag{4.29}$$

We calculate $\delta_i^{x,\ell,com}$ based on COMTRADE data. The non-US, non-commodity import US dollar invoicing share is calculated analogously. See Figure D.2 in Appendix D.2 for a depiction of the non-US, non-commodity export and import US dollar invoicing shares.

contractionary US monetary policy and a UIP shock rather than positive as in case of a positive US demand shock, in both cases output spillovers materialise through expenditure switching in trade between the economies in the rest of the world under partial DCP in the face of an appreciation of the US dollar. Hence, the variation in the output spillovers across different partial DCP scenarios is qualitatively identical across a variety of shocks that appreciate the US dollar multilaterally.

Against this background, column (3) in Table 4.8 reports results from the regression of Equation (4.27) in which the dependent variable is the average response of economies' real GDP over the first eight quarters after a contractionary US monetary policy shock that appreciates the US dollar has materialised. Column (5) in Table 4.8 reports analogous results for the output spillovers from a UIP shock. For both alternative shocks the results are consistent with the baseline, even if they are not estimated less precisely. In robustness checks below we report results for the output spillovers from US demand, monetary policy and UIP shocks based on identification using external instruments instead of sign restrictions.

Table 4.8: Results of regressions for real GDP spillovers from US monetary policy and UIP shocks

	Demand		Monetary policy		UIP	
	(1) SRs	(2) IV	(3) SRs	(4) IV	(5) SRs	(6) IV
Non-US X/M USD invoicing share differential	-0.003***	-0.012**	-0.002**	-0.003***	-0.007**	-0.011***
	(0.00)	(0.03)	(0.04)	(0.01)	(0.02)	(0.00)
Non-US trade rel. to GDP	0.001*	0.003*	0.000	0.001*	-0.001	0.001
	(0.06)	(0.09)	(0.55)	(0.08)	(0.19)	(0.33)
Trade with US rel. to GDP	0.004**	-0.002	-0.004*	-0.005***	-0.004	0.004
	(0.03)	(0.68)	(0.06)	(0.00)	(0.19)	(0.57)
Exchange rate flexibility against USD	0.096***	0.181*	-0.045	-0.055	-0.041	0.330***
	(0.00)	(0.08)	(0.45)	(0.11)	(0.26)	(0.01)
Adjusted R-squared	0.53	0.25	0.02	0.20	0.10	0.39
Observations	41	41	41	41	41	41

4.3.6 Spillovers from euro area demand shocks as a placebo test

DCP effects are in theory not constrained to arise in the context of a single dominant—or in this context better referred to as "vehicle"—currency that is used in trade between third countries. Against this background, it is in principle possible to test for the empirical relevance of partial DCP using export-import invoicing share differentials in any such vehicle currency and the spillovers from shocks in the corresponding currency issuer. Unfortunately, extensive cross-country data is only available for the shares of trade invoiced in the exporter's currency, the US dollar, and the euro. Hence, we can examine at most one additional vehicle currency besides the US dollar.

Looking more closely at the role of the euro in trade invoicing, the data suggest it is not an important vehicle currency, at least from a global perspective. Specifically, the panels in the top row of Figure 4.9 compare the share of economies' exports that is destined to the euro area with the share of economies' exports that is invoiced in euro (left-hand side panel) as well as the share of economies' exports destined to the US with the share of economies' exports invoiced in US dollar (right-hand side panel); the bottom panels present the corresponding comparisons for imports. The data document that in contrast to the US dollar, for most economies invoicing in euro occurs to an extent that matches economies' trade with the euro area. In fact, at best only Eastern European economies invoice a larger share of their exports in euro than the share of their exports that are destined to the euro area. The data for imports are similar.[82] In other words, in contrast to the US dollar the euro does not seem to be—at least globally—used as an invoicing currency in third-country trade, which is what makes a currency a dominant currency.

Against this background, we should not expect to find evidence that supports the predictions from partial DCP when considering the euro as a vehicle currency. In this sense, one may view the analysis of the empirical evidence for partial DCP using euro area demand shocks and euro invoicing shares as a placebo test. Table 4.9 reports results from regressions of Equation (4.27) analogous to those reported in Table 4.5. The sample size is reduced as all euro area economies are dropped while only the US is added. Columns (1) and (2) report results for the full sample, and columns (3) and (4) for a sample that includes only European and neighbouring economies. In the full sample, the coefficient estimate of the export-import euro invoicing share differential has a negative sign but is estimated rather imprecisely. In the extended European sample, the coefficient estimate is negative as well, but even less precisely estimated, which is however not surprising given the very small number of observations. Overall, the DCP hypothesis passes the placebo test based on euro appreciation and euro invoicing shares.

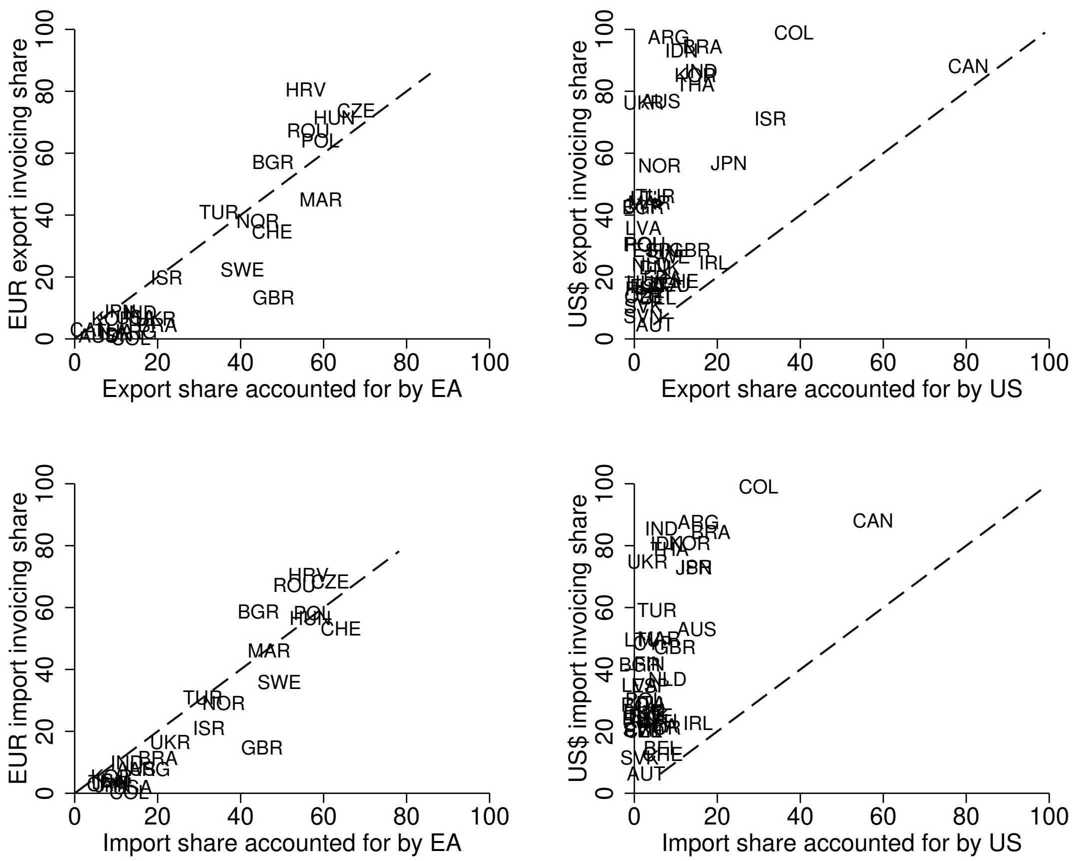
4.3.7 Robustness checks

Identification using external instruments

Columns (2), (4) and (6) in Table 4.8 report results from specifications in which we use spillover estimates obtained using external instruments instead of sign restrictions for identification in the VAR models (Stock and Watson 2012; Mertens and Ravn 2013). Specifically, we consider the estimated demand shock from the medium-scale DSGE model for the US studied in Kulish et al. (2017) as an external instrument for the structural US

[82] Third, in the context of the empirical exercise explored in this paper there is greater measurement error in the euro invoicing share differentials. In particular, unlike in case of the US, for the euro area it is not plausible to assume that if no data is available all of its exports and imports are invoiced in euro, see for example Figure 4.4. For this reason, in the regressions below we use the overall share of economies' exports and imports invoiced in euro—which includes trade invoiced in euro that involves the euro area—instead of the non-euro area export-import euro invoicing share differential. The variable that represents the importance of partial DCP in the regressions is thus not perfectly aligned with the relevant parameter in the model in Section 4.2.

Figure 4.9: Euro and US dollar invoicing shares compared with export and import shares



Note: The figure provides information about the relationship between economies' US dollar and Euro invoicing shares in exports and imports as well as the shares of trade accounted for the US and the euro area. In particular, the upper left-hand side panel compares economies' shares of exports invoiced in Euro and the share of exports destined to the euro area. The right-hand side panel depicts the corresponding comparison for imports, and the panels in the bottom row depict the corresponding comparisons for the US dollar invoicing shares and the share of trade accounted for by the US.

Table 4.9: Results of regressions for real GDP spillovers from a positive euro area demand shock

	All		Only Europe⁺	
	(1)	(2)	(3)	(4)
Exports EUR invoicing share	-0.009		-0.001	
	(0.22)		(0.82)	
Imports EUR invoicing share	0.011		0.008	
	(0.19)		(0.32)	
X/M EUR invoicing share differential		-0.009		-0.001
		(0.21)		(0.88)
Non-EA trade rel. to GDP	0.001	0.000	0.016**	0.013**
	(0.81)	(0.87)	(0.02)	(0.02)
Trade with EA rel. to GDP	0.002	0.003	-0.004	-0.000
	(0.41)	(0.15)	(0.37)	(0.94)
Exchange rate flexibility against EUR	0.135*	0.119**	0.196**	0.131**
	(0.09)	(0.04)	(0.03)	(0.01)
Adjusted R-squared	0.04	0.08	0.61	0.55
Observations	25	25	13	13

demand shocks in the VAR models.[83] For the US monetary policy shock in the VAR models we consider as instruments the intra-day changes in Federal Funds futures over narrow windows around US monetary policy decision meetings used in Gertler and Karadi (2015); we consider monetary policy surprises that are free from central bank information shocks as discussed in Jarocinski and Karadi (2020). And as in Engel and Wu (2018), for the UIP shock in the VAR models we consider as instruments the logarithm of the price of gold, whose price is often viewed as being driven by exogenous changes in demands for safe assets.[84] Our results are robust to these variations in the identification of US demand, monetary policy and UIP shocks.

Controlling for financial spillovers

In our baseline regression specification we control for other spillover transmission channels of US dollar appreciation, in particular bilateral and multilateral demand effects as well

[83]The instruments for the demand shock in the VAR model are given by the estimated time series of the "risk premium" and the "investment-specific technology" shocks; these shocks are labeled as demand shocks by Smets and Wouters (2003). Considering the model of Kulish et al. (2017) is appealing because it accounts for the zero lower bound and is estimated for a substantial overlap with our sample period, namely from 1983 to 2014. We are grateful to Mariano Kulish for sharing these shock series with us.

[84]In particular, Engel and Wu (2018) study the role of the convenience yield differential—which is equivalent to the CIP deviation—for the observed variation in exchange rates. Because the CIP deviation is a component of the UIP residual, exogenous variation in the CIP deviation can be exploited to identify exogenous variation in the UIP residual. In robustness checks Engel and Wu (2018) consider an instrumental variables regression to estimate the effect of changes in CIP deviations on the exchange rate. The results are very similar if we use the residuals from an AR(2) model as in to Stock and Watson (2012) instead of the level of the gold price as in Engel and Wu (2018). Similarly, Piffer and Podstawski (2018) suggest using variation in the real price of gold around specific events as an instrumental variable to identify uncertainty shocks. And Ludvigson et al. (2021) use the change of the gold price with a correlation restriction to identify uncertainty shocks.

as expenditure switching. However, shocks that induce US dollar appreciation may also transmit to the global economy by affecting domestic financial conditions in the US and spill over to financial conditions abroad (Bruno and Shin 2015; Gourinchas et al. 2019). The latter operates on economies' US dollar exposures, which may be related to and may hence correlate with exposure to US dollar trade invoicing (Gopinath and Stein 2018).

We control for the possible role of financial channels in transmitting a US dollar appreciation globally by augmenting $\boldsymbol{z}_i$ in Equation (4.27) to include several variables reflecting economies' financial exposure to the US dollar in particular and to global financial markets more generally. Specifically, we add economies' gross foreign asset and liability position relative to GDP, typically used as a measure of overall financial integration (column (3)); the US dollar-denominated foreign assets (excluding foreign exchange reserves) and liabilities relative to GDP, respectively, reflecting economies' susceptibility to balance sheet exchange rate valuation effects in response to the appreciation of the US dollar and the size of US dollar re-financing needs (columns (4) and (5)); and the share of US dollar foreign liabilities in total foreign-currency liabilities, reflecting the diversification and hence the ease of substituting US dollar liabilities by other currencies (column (6)).[85] For the regressions that include variables reflecting economies' susceptibility to financial spillovers, we exclude financial centers, namely Belgium, Ireland, Netherlands, Switzerland, and the UK. We also exclude Argentina, for which the foreign-currency exposure data might be less reliable due to its sovereign defaults in the early 2000s.[86]

Table 4.10 reports the results from the regressions which include the financial variables. Column (1) replicates the baseline results and column (2) reports the results from the baseline specification for the reduced sample without financial centers and Argentina. And column (8) reports results from a regression in which financial centers and Argentina are included. The results for the non-US export-import US dollar invoicing share differential are unchanged.

Alternative VAR model specifications

Table 4.11 reports results from regressions using spillover estimates obtained from alternative specifications of the VAR models. In particular, column (2) reports results from a regression using spillover estimates from Bayesian VAR models with standard Minnesota-type priors using hyper-parameters conventionally used in the literature. Columns (3) and (4) report results from regressions using spillover estimates obtained from eight and five-variable rather than seven-variable VAR models; in the eight-variable VAR model we add rest-of-the-world real GDP in order to account for higher-order spillovers as discussed in Georgiadis (2017), and in the five-variable VAR models we only include the spillover recipient economy's

[85]The gross foreign asset and liability data are taken from updates of the External Wealth of Nations Database of Lane and Milesi-Ferretti (2007), and the foreign-currency exposure related data from the update of Lane and Shambaugh (2010) constructed by Benetrix et al. (2015).

[86]Data for Bulgaria and Cyprus are not available in Lane and Shambaugh (2010) or Benetrix et al. (2015).

Table 4.10: Results of regressions for real GDP spillovers from a positive US demand shock when controlling for financial spillover channels

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Non-US X/M USD invoicing share differential	-0.003***	-0.003***	-0.003***	-0.004***	-0.004***	-0.004***	-0.005***	-0.003***
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)
Non-US trade rel. to GDP	0.001*	0.001	0.001	0.001*	0.001*	0.000	0.000	0.000
	(0.06)	(0.14)	(0.13)	(0.07)	(0.05)	(0.74)	(0.89)	(0.44)
Trade with US rel. to GDP	0.004**	0.002	0.001	-0.000	0.001	0.002	-0.000	0.005**
	(0.03)	(0.21)	(0.43)	(0.88)	(0.69)	(0.21)	(0.99)	(0.03)
Exchange rate flexibility against USD	0.096***	0.081***	0.064**	0.071***	0.106***	0.059**	0.077**	0.098**
	(0.00)	(0.00)	(0.03)	(0.00)	(0.00)	(0.04)	(0.03)	(0.02)
GFAL/GDP			0.050				-0.014	0.010
			(0.35)				(0.85)	(0.86)
USD foreign assets (excl. res.)/GDP				0.002			0.000	-0.001
				(0.22)			(0.88)	(0.32)
USD foreign liabilities/GDP					0.004**		0.005	0.003
					(0.05)		(0.10)	(0.25)
Share of USD foreign liabilities						-0.001	-0.002	-0.001
						(0.31)	(0.19)	(0.27)
Adjusted R-squared	0.53	0.42	0.42	0.43	0.48	0.41	0.46	0.50
Observations	41	33	33	33	33	33	33	39

real GDP in $\boldsymbol{x}_{it}$ in Equation (4.24) in order to increase the number of degrees of freedom. Column (5) reports results from a regression using spillover estimates obtained from VAR models in which we include dummy variables (and their lags) as exogenous variables in order to remove global, regional or country-specific extreme events from the sample, such as the global, the Asian or the Argentine financial crisis.[87] Column (6) reports the results from a robustness check in which we use spillover estimates obtained from VAR models estimated only up to 2007, ending the sample period just prior to the global financial crisis and the zero-lower bound period.[88] And finally, column (7) reports results from a regression using spillover estimates obtained assuming that the non-US economy is a small open economy, i.e. imposing block exogeneity in the two-country VAR models. In all cases the results are robust.

Alternative spillover definitions

Table 4.12 reports results from regressions in which we consider alternative definitions of the dependent variable $\Delta\widehat{y}_i$ in Equation (4.27). Recall that in the baseline we consider the average real GDP spillover estimate over the first eight quarters after the shock that appreciates the US dollar has occurred. Columns (2) to (4) report results from regressions in which the dependent variable is instead the average real GDP spillover over the first

[87]Due to the large number of parameters to be estimated in some country cases, we consider VAR models with five variables as in column (4) and with only one lag for this exercise.

[88]Due to the reduction of the length of the sample period, we consider VAR models with five variables as in column (4) and with only one lag for this exercise.

Table 4.11: Results of regressions for real GDP spillovers from a positive US demand shock with alternative spillover definitions

	(1) Baseline	(2) 1-year	(3) 3-year	(4) 4-year	(5) Panel
Non-US X/M USD invoicing share differential	-0.003***	-0.002***	-0.003***	-0.003**	-0.003***
	(0.00)	(0.01)	(0.01)	(0.03)	(0.01)
Non-US trade rel. to GDP	0.001*	0.000	0.001**	0.001*	0.001*
	(0.06)	(0.28)	(0.04)	(0.06)	(0.06)
Trade with US rel. to GDP	0.004**	0.004**	0.004*	0.004*	0.005**
	(0.03)	(0.03)	(0.09)	(0.08)	(0.03)
Exchange rate flexibility against USD	0.096***	0.080***	0.112*	0.125*	0.133*
	(0.00)	(0.00)	(0.06)	(0.09)	(0.05)
Adjusted R-squared	0.53	0.50	0.49	0.47	
Observations	41	41	41	41	820

four, twelve or sixteen quarters, respectively. And column (5) reports results from a panel regression in which we include all horizons, namely

$$\Delta\widehat{y}_{ih} = \alpha_h + \beta \cdot \delta_i + \boldsymbol{\gamma}'\boldsymbol{z}_i + \eta_{ih}, \tag{4.30}$$

where h represents the impulse response horizon and α_h are horizon fixed effects. Our results are robust to these alternative definitions of the dependent variable in the cross-section regression.

Table 4.12: Alternative VAR model specifications

	(1) Baseline	(2) BVAR	(3) 8 vars	(4) 5 vars	(5) Dummies	(6) 95-07	(7) SOE
Non-US X/M USD invoicing share differential	-0.003***	-0.002***	-0.003***	-0.004***	-0.002**	-0.004*	-0.003**
	(0.00)	(0.00)	(0.00)	(0.01)	(0.05)	(0.08)	(0.01)
Non-US trade rel. to GDP	0.001*	0.001**	0.000	0.001*	0.001	-0.000	0.001
	(0.06)	(0.02)	(0.24)	(0.09)	(0.10)	(0.77)	(0.11)
Trade with US rel. to GDP	0.004**	0.002	0.003***	0.005**	0.004*	0.008***	0.004*
	(0.03)	(0.19)	(0.01)	(0.03)	(0.10)	(0.00)	(0.07)
Exchange rate flexibility against USD	0.096***	0.048*	0.100***	0.126***	0.087***	0.169***	0.121***
	(0.00)	(0.07)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)
Adjusted R-squared	0.53	0.50	0.47	0.57	0.33	0.31	0.52
Observations	41	41	41	41	41	41	41

Alternative country samples

Table 4.13 reports results from regressions of Equation (4.27) for alternative country samples. In particular, columns (2) reports regression results when we drop Australia and Norway, which stand out in the conditional scatterplot in the right-hand side panel in Figure 4.7. Column (3) reports results from a regression in which we drop financial centers, namely Belgium, Cyprus, Ireland, Netherlands, Switzerland, and the UK. And in column (4) we report results from a regression in which all economies dropped in columns (2) and (3) are dropped together. Finally, in column (5) we report results from a larger sample in which we additionally include Algeria, Iceland, Luxembourg, and Pakistan, which are not included in the baseline because they are either very small countries, do not have quarterly data for real GDP, or are extremely dependent on oil exports. The results hardly change across country samples.

Table 4.13: Results of regressions for real GDP spillovers from a positive US demand shock for alternative country samples

	(1) Baseline	(2) -AUS/NOR	(3) -Fin. centers	(4) Small N	(5) Large N
Non-US X/M USD invoicing share differential	-0.003***	-0.005**	-0.004***	-0.007***	-0.005***
	(0.00)	(0.01)	(0.00)	(0.01)	(0.01)
Non-US trade rel. to GDP	0.001*	0.001**	0.001*	0.001**	0.000
	(0.06)	(0.03)	(0.10)	(0.04)	(0.48)
Trade with US rel. to GDP	0.004**	0.003*	0.002*	0.001	0.002
	(0.03)	(0.07)	(0.09)	(0.30)	(0.44)
Exchange rate flexibility against USD	0.096***	0.078**	0.095***	0.068**	0.063
	(0.00)	(0.01)	(0.00)	(0.02)	(0.11)
Adjusted R-squared	0.53	0.54	0.53	0.57	0.35
Observations	41	39	35	33	45

Alternative cross-section regression specifications

Table 4.14 reports results from alternative cross-section regression specifications. Column (2) reports results from a median regression that is robust to outliers, and column (3) from a weighted least squares regression in which the weights are given by global GDP shares. Again, the results are robust across specifications.

Table 4.14: Alternative regression specifications

	(1) Baseline	(2) rreg	(3) GDP weights
Non-US X/M USD invoicing share differential	-0.003***	-0.003*	-0.003**
	(0.00)	(0.06)	(0.03)
Non-US trade rel. to GDP	0.001*	0.001	0.001
	(0.06)	(0.11)	(0.19)
Trade with US rel. to GDP	0.004**	0.003*	0.003**
	(0.03)	(0.07)	(0.01)
Exchange rate flexibility against USD	0.096***	0.096***	0.084***
	(0.00)	(0.00)	(0.00)
Adjusted R-squared	0.53	0.44	0.64
Observations	41	41	41

4.4 Conclusion

Different assumptions on the pricing of economies' exports have different implications for a host of issues that are critical for policymakers, including optimal monetary policy, optimum currency areas and international monetary policy co-ordination. Against this background, in this paper we contribute to the growing literature that documents the empirical relevance of DCP. Specifically, we first derive a prediction regarding the cross-country heterogeneity in output spillovers from shocks that appreciate the US dollar multilaterally from a structural three-country model: Under partial DCP, the output spillovers decline with an economy's non-US export-import US dollar invoicing share differential. We then estimate the output spillovers from US demand, monetary policy and UIP shocks that all appreciate the US dollar multilaterally, and correlate them with economies' non-US export-import US dollar invoicing share differentials. Our findings are consistent with the hypothesis that a large share of prices in global non-US trade are sticky in US dollar rather than the exporter's or the importer's currency. We contribute to the existing literature in particular by emphasising the role of the export-import US dollar invoicing share differentials, that is the role of asymmetries in the relevance of partial DCP across countries' exports and imports. Moreover, we contribute by exploiting exogenous variation in the US dollar exchange rate in order to evaluate the empirical relevance of DCP.

Appendix A

Appendix for Chapter 1

A.1 Data description

Table A.1: Detailed description of data used in the VAR analysis

Variable	Description	Notes	Source
1-year yield	Germany Government 1 year yield	End of period	Macrobond Financial AB
2-year yield	Germany Government 2 year yield	End of period	Macrobond Financial AB
US/EUR	US-Dollar per Euro, spot rate	Monthly average of daily values	Macrobond Financial AB
Industrial Production	Euro Area Industrial Production excl. Construction		Eurostat
Brent oil price	Brent crude Europe Spot price FOB, US-Dollar per barrel	Monthly average of daily values	Energy Information Administration
CPI (headline)	Euro Area Harmonized Index of Consumer Prices	Seasonally adjusted using X13	Eurostat
HICP housing	Euro Area, HICP, Housing, Water & Electricity & Gas & Other Fuels	Seasonally adjusted using X13	Eurostat
HICP transport	Euro Area, HICP, Transport	Seasonally adjusted using X13	Eurostat
HICP heating	Euro Area, HICP, Housing, Water, Electricity, Fuel, Electricity, Gas	Seasonally adjusted using X13	Eurostat
HICP fuels	Euro Area, HICP, Fuels & Lubricants for Personal Transport Equipment	Seasonally adjusted using X13	Eurostat
HICP energy	Euro Area, HICP, Energy	Seasonally adjusted using X13	Eurostat
Credit spread	ICE BofA Euro High Yield Index Option-Adjusted Spread	Monthly average of daily values	FRED
Euro Area monetary policy proxy	3 month (monetary event window) OIS surprise	Calculated based on data and methodology by Jarocinski and Karadi (2020)	Jarocinski and Karadi (2020) and authors' calculations
Global oil production	Global oil production (million barrels/day)		Baumeister and Hamilton (2019)
Oil inventories	Change in global oil inventories		Baumeister and Hamilton (2019)
Global IP	Global industrial production		Baumeister and Hamilton (2019)
Consensus 1-year ahead inflation expectations	(GDP) Weighted average of Germany, France, Italy, and Spain	I use the big 4 euro area countries since the euro area aggregate is only available starting in 2002.	
Oil supply news proxy	Suprise in oil futures prices around OPEC announcements	Monthly sum of daily values	Känzig (2021)

Note that, as in Born and Pfeifer (2021), I demean the variables to avoid numerical problems arising from under/overflow during the posterior computations that involve the sum of squares in Sections 1.2 and 1.3. In Section 1.4 I no longer do so because adding the 5-year yield to the baseline specification solves the issue.

A.2 Details on the Bayesian Proxy SVAR model

In this appendix, I briefly lay out the BPSVAR model for the general case with $k \geq 1$ proxy variables and k structural shocks of interest. For convenience, I reproduce the BSPVAR model equations from section 1.2.1.

Following the notation of Rubio-Ramirez et al. (2010), consider without loss of generality the structural VAR model with one lag and without deterministic terms

$$\boldsymbol{y}_t'\boldsymbol{A}_0 = \boldsymbol{y}_{t-1}'\boldsymbol{A}_1 + \boldsymbol{\epsilon}_t', \qquad \boldsymbol{\epsilon} \sim N(\boldsymbol{0}, \boldsymbol{I}_n), \tag{A.1}$$

where $\boldsymbol{y}_t$ is an $n \times 1$ vector of endogenous variables and $\boldsymbol{\epsilon}_t$ an $n \times 1$ vector of structural shocks. The BPSVAR framework builds on the following assumptions in order to identify k structural shocks of interest: There exists a $k \times 1$ vector of proxy variables $\boldsymbol{m}_t$ that are (i) correlated with the k structural shocks of interest $\boldsymbol{\epsilon}_t^*$ and (ii) orthogonal to the remaining structural shocks $\boldsymbol{\epsilon}_t^o$. Formally, the identifying assumptions are

$$E[\boldsymbol{\epsilon}_t^*\boldsymbol{m}_t'] = \underset{(k\times k)}{\boldsymbol{V}}, \tag{A.2a}$$

$$E[\boldsymbol{\epsilon}_t^o\boldsymbol{m}_t'] = \underset{((n-k)\times k)}{\boldsymbol{0}}, \tag{A.2b}$$

and represent the relevance and the exogeneity condition, respectively.

Denote by $\tilde{\boldsymbol{y}}_t' \equiv (\boldsymbol{y}_t', \boldsymbol{m}_t')$, by $\tilde{\boldsymbol{A}}_\ell$ the corresponding $\tilde{n} \times \tilde{n}$ coefficient matrices with $\tilde{n} = n + k$, by $\tilde{\boldsymbol{\epsilon}} \equiv (\boldsymbol{\epsilon}_t', \boldsymbol{v}_t')' \sim N(\boldsymbol{0}, \boldsymbol{I}_{n+k})$, where $\boldsymbol{v}_t$ is a $k \times 1$ vector of measurement errors (see below). The augmented structural VAR model is then given by

$$\tilde{\boldsymbol{y}}_t'\tilde{\boldsymbol{A}}_0 = \tilde{\boldsymbol{y}}_{t-1}'\tilde{\boldsymbol{A}}_1 + \tilde{\boldsymbol{\epsilon}}_t'. \tag{A.3}$$

To ensure that the augmentation with equations for the proxy variables does not affect the dynamics of the endogenous variables, the coefficient matrices $\tilde{\boldsymbol{A}}_\ell$ are specified as

$$\tilde{\boldsymbol{A}}_\ell = \begin{pmatrix} \underset{(n\times n)}{\boldsymbol{A}_\ell} & \underset{(n\times k)}{\boldsymbol{\Gamma}_{\ell,1}} \\ \underset{(k\times n)}{\boldsymbol{0}} & \underset{(k\times k)}{\boldsymbol{\Gamma}_{\ell,2}} \end{pmatrix}, \qquad \ell = 0, 1. \tag{A.4}$$

The zero restrictions on the lower left-hand side block imply that the proxy variables do not enter the equations of the endogenous variables. The reduced form of the model is

$$\tilde{\boldsymbol{y}}_t' = \tilde{\boldsymbol{y}}_{t-1}'\tilde{\boldsymbol{A}}_1\tilde{\boldsymbol{A}}_0^{-1} + \tilde{\boldsymbol{\epsilon}}_t{}'\tilde{\boldsymbol{A}}_0^{-1}. \tag{A.5}$$

Because the inverse of $\tilde{\boldsymbol{A}}_0$ in Equation (A.4) is given by

$$\tilde{\boldsymbol{A}}_0^{-1} = \begin{pmatrix} \boldsymbol{A}_0^{-1} & -\boldsymbol{A}_0^{-1}\boldsymbol{\Gamma}_{0,1}\boldsymbol{\Gamma}_{0,2}^{-1} \\ 0 & \boldsymbol{\Gamma}_{0,2}^{-1} \end{pmatrix}, \tag{A.6}$$

the last k equations of the reduced form of the VAR model in Equation (A.5) read as

$$\boldsymbol{m}_t' = \tilde{\boldsymbol{y}}_{t-1}' \tilde{\boldsymbol{A}}_1 \begin{pmatrix} -\boldsymbol{A}_0^{-1} \boldsymbol{\Gamma}_{0,1} \boldsymbol{\Gamma}_{0,2}^{-1} \\ \boldsymbol{\Gamma}_{0,2}^{-1} \end{pmatrix} - \boldsymbol{\epsilon}_t' \boldsymbol{A}_0^{-1} \boldsymbol{\Gamma}_{0,1} \boldsymbol{\Gamma}_{0,2}^{-1} + \boldsymbol{v}_t' \boldsymbol{\Gamma}_{0,2}^{-1}, \tag{A.7}$$

which shows that in the BPSVAR framework the proxy variables may be serially correlated and affected by past values of the endogenous variables and measurement error.

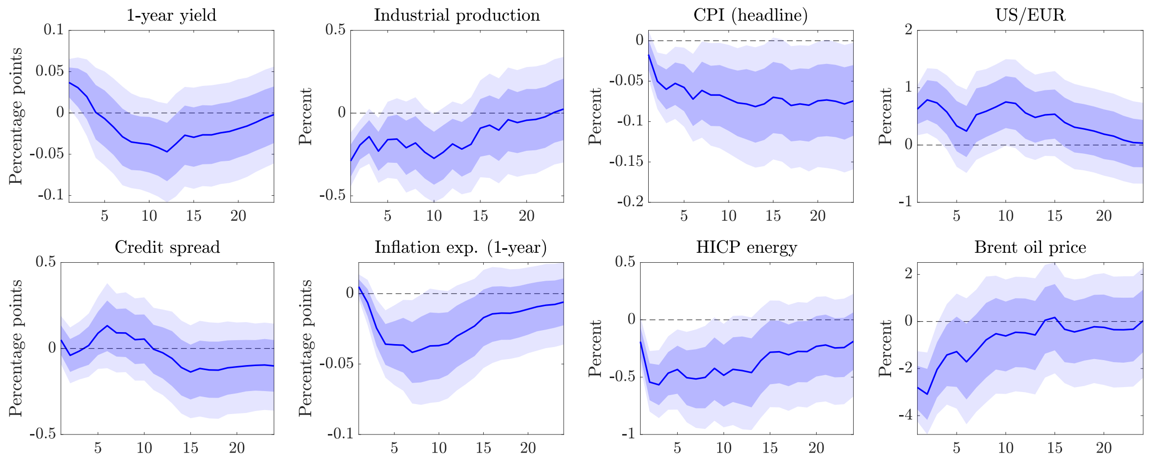
Ordering the structural shocks so that $\boldsymbol{\epsilon}_t = (\boldsymbol{\epsilon}_t^{o\prime}, \boldsymbol{\epsilon}_t^{*\prime})'$ yields

$$E\left[\boldsymbol{\epsilon}_t \boldsymbol{m}_t'\right] = -\boldsymbol{A}_0^{-1} \boldsymbol{\Gamma}_{0,1} \boldsymbol{\Gamma}_{0,2}^{-1} = \begin{pmatrix} \underset{((n-k)\times k)}{\boldsymbol{0}} \\ \underset{(k\times k)}{\boldsymbol{V}} \end{pmatrix}. \tag{A.8}$$

The first equality is obtained using Equation (A.7) and because the structural shocks $\boldsymbol{\epsilon}_t$ are by assumption orthogonal to $\boldsymbol{y}_{t-1}$ and $\boldsymbol{v}_t$. The second equality is due to the exogeneity and relevance conditions in Equations (A.2a) and (A.2b). Equation (A.8) shows that the identifying assumptions imply restrictions on the last k columns of the contemporaneous structural impact coefficients in $\tilde{\boldsymbol{A}}_0^{-1}$. In particular, if the exogeneity condition in Equation (A.2b) holds, the first $n-k$ rows of the upper right-hand side sub-matrix $\boldsymbol{A}_0^{-1} \boldsymbol{\Gamma}_{0,1} \boldsymbol{\Gamma}_{0,2}^{-1}$ of $\tilde{\boldsymbol{A}}_0^{-1}$ in Equation (A.6) are zero. From the reduced form in Equation (A.5) it can be seen that this implies that the first $n-k$ structural shocks do not impact contemporaneously the proxy variables. In turn, if the relevance condition in Equation (A.2a) holds, the last k rows of the upper right-hand side sub-matrix $\boldsymbol{A}_0^{-1} \boldsymbol{\Gamma}_{0,1} \boldsymbol{\Gamma}_{0,2}^{-1}$ of $\tilde{\boldsymbol{A}}_0^{-1}$ are different from zero. From the reduced form in Equation (A.5) it can be seen that this implies that the last k structural shocks impact the proxy variables contemporaneously. The Bayesian estimation algorithm of Arias et al. (2021) determines the estimates of $\boldsymbol{A}_0$ and $\boldsymbol{\Gamma}_{0,\ell}$ such that the restrictions on $\tilde{\boldsymbol{A}}_0^{-1}$ implied by Equations (A.2a) and (A.2b) as well as on $\tilde{\boldsymbol{A}}_\ell$ in Equation (A.4) are simultaneously satisfied, and hence the estimation identifies the structural shocks $\boldsymbol{\epsilon}_t^*$.

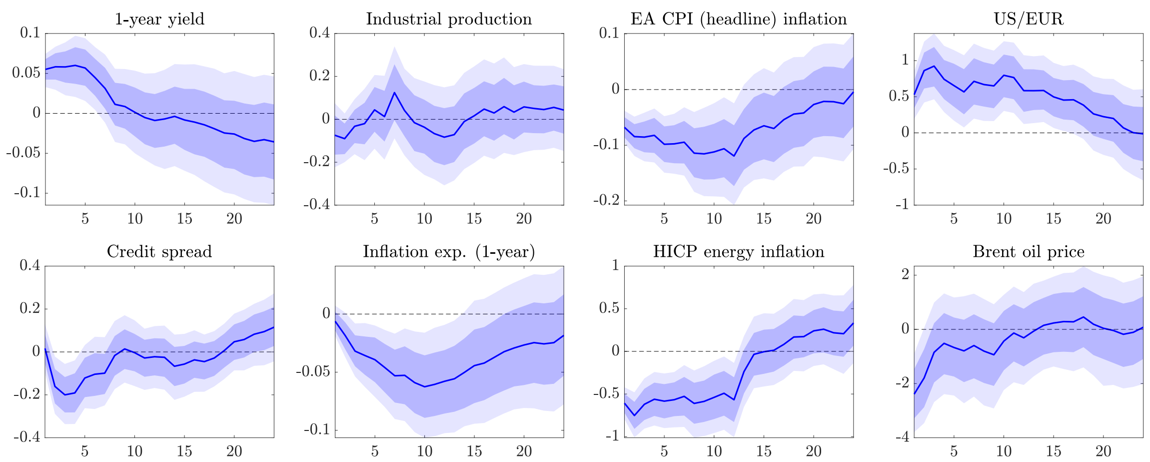
A.3 Robustness for the Baseline BPSVAR model

Figure A.1: Euro Area SVAR model, zero proxy relevance prior threshold



Notes: Euro Area model, with the prior on the relevance of the shock for the proxy set to 0.1%. Impulse response functions to a one standard deviation monetary policy shock. Point-wise posterior means along with 68% and 90% point-wise probability bands. Horizon in months.

Figure A.2: IRFs to a contemporaneous EA Monetary policy shock when estimating the BPSVAR using data until October 2023



Notes: Euro Area SVAR model including the Pandemic (see text for details). Impulse response functions to a one standard deviation monetary policy shock. Point-wise posterior means along with 68% and 90% point-wise credible sets. Horizon in months.

A.4 Details on the hybrid approximation of the structural counterfactual

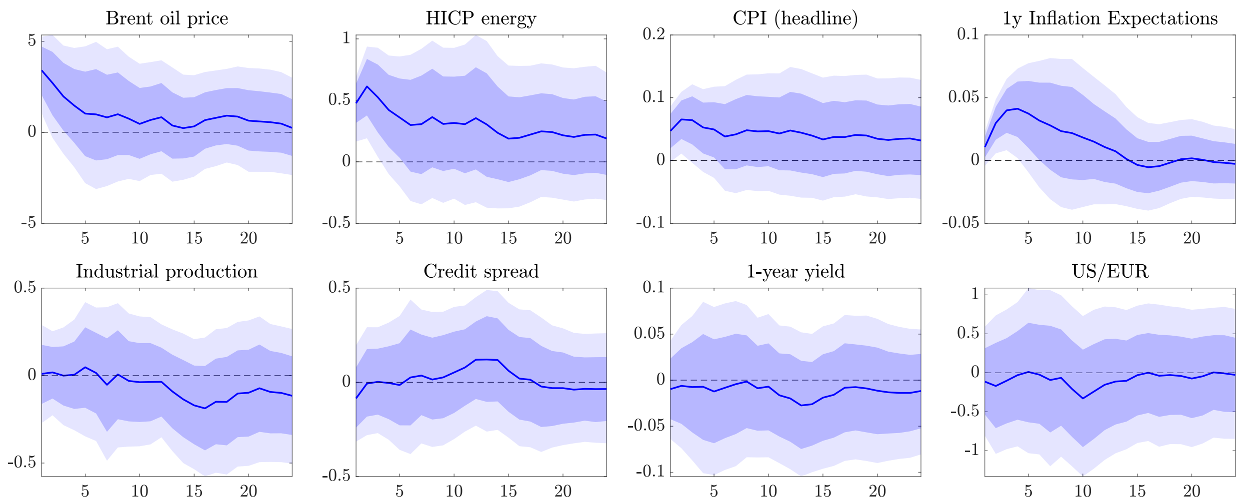
As with the SZ approximation its possible to parse the sequential procedure for the "hybrid counterfactual" of MW directly into the representation in Equation (1.10) and thereby uncover the assumptions on the underlying structural model as described by the matrices $\tilde{\boldsymbol{\mathcal{A}}}, \boldsymbol{\Theta}_{\nu,\mathcal{A}}, \boldsymbol{\Theta}_{\epsilon,\mathcal{A}}$, under which the resulting counterfactual path of the endogenous variables and the policy instrument actually correspond to the true counterfactual. The intuition for the underlying assumed structural model is that agents are myopic with respect to announced changes to the future level of the policy instrument. In this economy, there exists a policymaker who can credibly announce changes to the policy instruments materializing today and furthermore controls the *time t expectations* about the n_e periods ahead level of the policy instruments. In other words, the time t expectations about the future path of the policy instrument for n_e periods themselves become a policy instrument with its own policy rule and crucially with its own distinct shocks to those expectations which already enter the information set of the agents at time t. For example, in each period t there not only exists a shock to the current (0 periods ahead) level of the policy instrument ν_0 but also a shock to the *time t expectations* about the one period ahead level $\nu_{\mathcal{E}_t,1}$. The latter already enters the information set of agents at time t, which stands in contrast to an announced change to the level of the instrument ν_1. For the matrix notation of the hybrid counterfactual in each period there are $n_e + 1$ policy rules which have to be embedded into Equation (1.10) by appropriately changing the matrix $\tilde{\boldsymbol{\mathcal{A}}}$. Furthermore all n_e period ahead, model consistent, expectations of all the n_p variables that are necessary to describe the expected policy rule need to be appended to the matrices of impulse responses $\boldsymbol{\Theta}_{\nu,\mathcal{A}}$ and $\boldsymbol{\Theta}_{\epsilon,\mathcal{A}}$. This implies that the number of variables in the system of Equation 1.10 increases to $n^{\text{hybrid}} = n + (n_e \times n_p)$. The second crucial assumption is that for each period t, the policymaker has access to a set of n_e distinct policy shocks which are shocks to the n_e period ahead of expectations about the policy instrument which implies that the number of n_ν policy (news) shocks in $\boldsymbol{\nu}$ increases to $n_\nu^{\text{hybrid}} = n_h \times (n_e + 1)$. Under these assumptions its possible to show that for each distinct policy shock ν_i, i.e. for each distinct shock to the level of the policy rate or its time t expectations, and each variable j, the matrix $\boldsymbol{\Theta}_{\nu_i,j,\mathcal{A}}^{\text{hybrid}}$is lower triangular. The intuition is that, for each period t agents only observe/care about the shock of that type that materialized in that period. To be more precise, each column c of matrix $\boldsymbol{\Theta}_{\nu_i,j,\mathcal{A}}^{\text{hybrid}}$, which describes the n_h responses of variable j to a news shock of type ν_i realized in c periods, is given by $\boldsymbol{\Theta}_{\nu_i,j,\mathcal{A},c} = [\mathbf{0}'_{1\times(c-1)}, \boldsymbol{\Theta}'_{\nu_{i,t},j,\mathcal{A},0:(n_h-c)}]'$ where the vector $\boldsymbol{\Theta}_{\nu_{i,t},\mathcal{A},j,0:(n_h-c)}$ describes the impulse responses of variable j to a shock announcing a change to the $0 \leq i \leq n_h$ period ahead expectation of the policy instrument at time t for periods 0 to $n_h - i$. Thus with the estimated effects to a contemporaneous policy shock $\nu_{0,t}$ and estimated effects for n_e forward guidance shocks at hand, an econometrician can recover each of the $n_\nu^{\text{hybrid}} \times n_h$ matrices $\boldsymbol{\Theta}_{\nu_i,j,\mathcal{A}}^{\text{hybrid}}$ for all n^{hybrid} variables. She can then , for the contemporaneous shock and all forward guidance shocks, $v_0 ... v_{n_e}$, stack the matrices across variables to form $\mathbf{\Theta}^{\text{hybrid}}_{\nu_i,\mathcal{A}} = [\mathbf{\Theta}^{\text{hybrid}'}_{\nu_i,1,\mathcal{A}}, \mathbf{\Theta}^{\text{hybrid}'}_{\nu_i,2,\mathcal{A}}, ..., \mathbf{\Theta}^{\text{hybrid}'}_{\nu_i,n^{\text{hybrid}},\mathcal{A}}]'$ and then stack all those matrices to arrive at $\mathbf{\Theta}^{\text{hybrid}}_{\nu,\mathcal{A}} = [\mathbf{\Theta}^{\text{hybrid}'}_{\nu_1,\mathcal{A}}, \mathbf{\Theta}^{\text{hybrid}'}_{\nu_2,\mathcal{A}}, ..., \mathbf{\Theta}^{\text{hybrid}'}_{\nu_{n_e},\mathcal{A}}]'$ which under the assumptions outlined above corresponds to $\mathbf{\Theta}^{\text{true}}_{\nu,\mathcal{A}}$. With this structure the matrix $\left(\tilde{\boldsymbol{\mathcal{A}}}[\mathbf{\Theta}^{\text{hybrid}}_{\nu,\mathcal{A}}]\right)$ is invertible and the solution to the problem in Equation 1.10 is given by $\tilde{\boldsymbol{\nu}}_{\text{hybrid}} = -\left(\tilde{\boldsymbol{\mathcal{A}}}[\mathbf{\Theta}^{\text{hybrid}}_{\nu,\mathcal{A}}]\right)^{-1} \tilde{\boldsymbol{\mathcal{A}}}[\mathbf{\Theta}_{A,\epsilon}]$.

Intuitively, given the assumed structure of $\mathbf{\Theta}^{\text{hybrid}}_{\nu,\mathcal{A}}$ implied by the "hybrid counterfactual" it becomes apparent that the hybrid solution will correspond to the true solution in Equation (1.10) if expectations about the policy instrument more than n_e periods ahead expectations do not enter the decision problem of agents in the underlying data generating structural model. This makes the expected counterfactual path of the policy instrument for these periods irrelevant and therefore the approximation error from neglecting those expectations vanishes. This counterfactual approach could for instance be a good approximation if, in the underlying true model, agents follow some form of exponential discounting, so that the importance of the expectations decays quickly with the horizon. To summarize: Although the hybrid approach perfectly enforces the policy rule in each period and in n_e period ahead expectations, the accuracy of the approximation depends on the degree of myopia of agents in the non-policy block and/or the information set of those agents in the (unknown) structural model underlying the data. [89]

[89] Despite these stark assumptions on the degree of myopia in the true underlying model, it's possible to show as the number of explicit policy rules $n_e + 1$ approaches n_h, the resulting counterfactual from this approach converges to the true counterfactual independent of the true degree of myopia in the underlying structural model. The intuition is that, by announcing the policy rule and corresponding deviations for each period already at time t, agents at time t directly observe those rules and therefore form the correct expectations about the future path of the policy instrument. Even for a full information rational expectations HANK model, the estimated counterfactual where the counterfactual policy rule is only enforced at point time t and in $n_e = 1$ period ahead expectations, this approximation already comes very close to the true counterfactual obtained from the underlying model (see McKay and Wolf (2022a)).

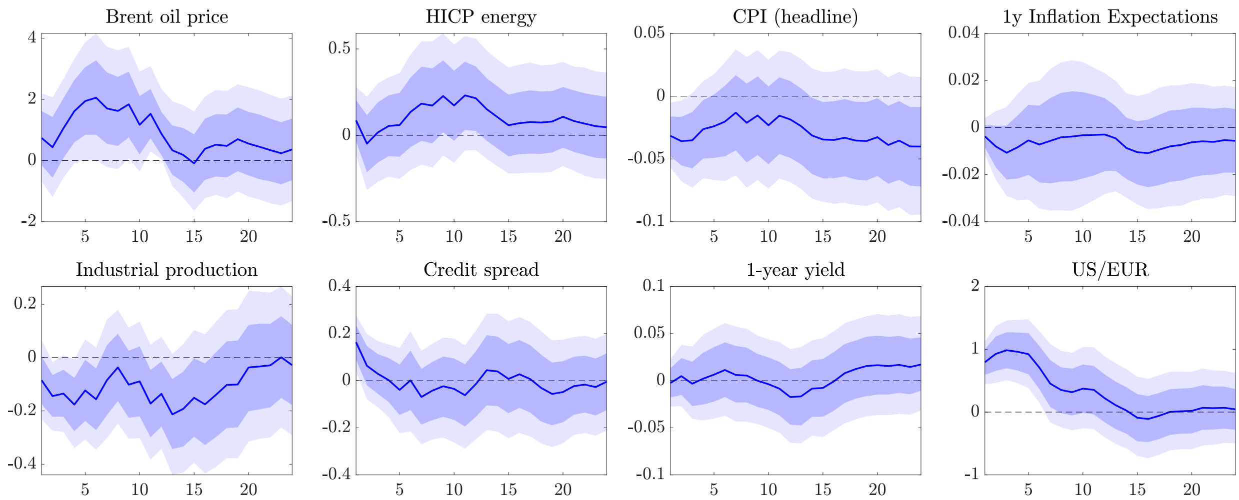
A.5 Additional information for the OPEC policy rule counterfactual

Figure A.3: IRFs to short-run oil supply news shocks



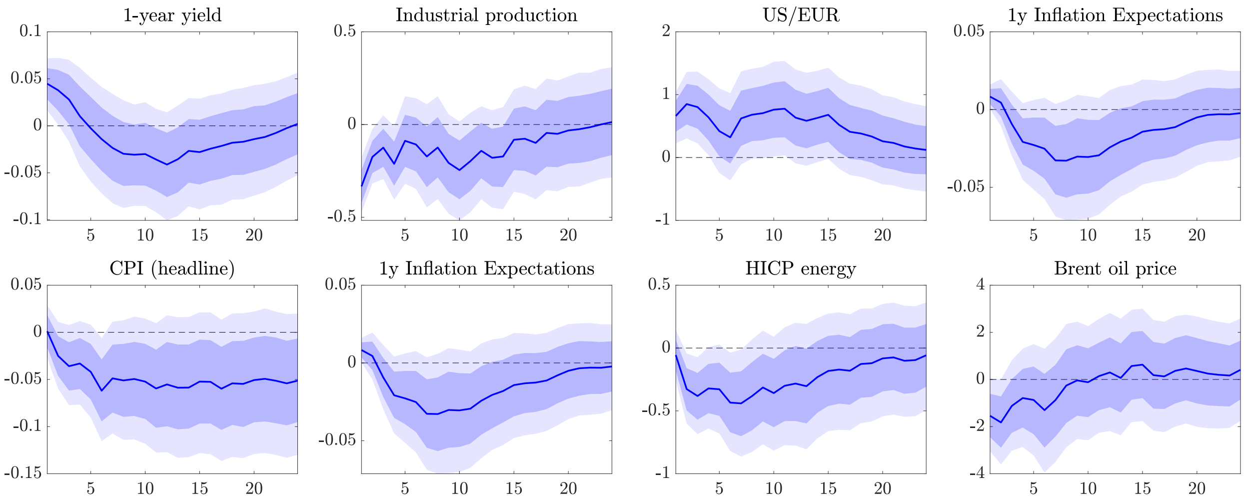
Notes: Impulse response functions to a one standard deviation short-run oil supply news shock showing the point-wise posterior means along with 68% and 90% point-wise credible sets in blue.

Figure A.4: IRFs to medium-run oil supply news shocks



Notes: Impulse response functions to a one standard deviation medium-run oil supply news shock showing the point-wise posterior means along with 68% and 90% point-wise credible sets in blue.

Figure A.5: IRFs to a euro area monetary policy shock when jointly identified alongside the oil supply shocks



Notes: Impulse response functions to a one standard deviation monetary policy shock. Point-wise posterior means along with 68% and 90% point-wise credible sets. Horizon in months.

A.6 Details on estimating the impulse responses under optimal policy

A.6.1 Estimating the baseline optimal policy counterfactual

As laid out in the section 1.4 I construct the estimate of the counterfactual impulse responses under optimal policy using the same multistep-step procedure as in McKay and Wolf (2023), Wolf (2023) and Caravello et al. (2024). I describe the approach in more detail below.

First, I estimate impulse responses to an identified oil price shock. To do so I use the same endogenous variables as in the baseline BPSVAR model and include the 5-year German Bund yield. All variables enter the estimation in log levels if they are not already expressed in percentage terms. I estimate the impulse responses starting the sample in 1999.

Second, I identify the euro area conventional monetary policy and forward guidance shocks by combining the high-frequency proxies with the zero, magnitude, and sign restrictions described in the text. Again I use the same variables and transformations as in step 1. As stated in the main text for this estimation I start the sample in 2002 to take into account the liquidity concerns raised by Altavilla et al. (2019), which are particularly severe for the high-frequency changes in the longer run maturities.

Third, I condition on the impulse responses from the first step and compute the optimal policy counterfactual for each draw from the posterior distribution of the second step.

Lastly, I plot the point-wise mean which can be interpreted as summarizing the posterior distribution of impulse responses under the optimal policy response conditional on the data and the impulse responses from the first step.

A.6.2 Estimating the impulse responses under counterfactual optimal policy

The procedure for the counterfactual optimal policy is very similar to the one sketched in Appendix A.6.1 but involves two additional steps.

First, I estimate impulse responses to a short-run oil supply news shock, which in the main text I refer to as oil price or oil supply shock for short. I use the same endogenous variables as in the baseline BPSVAR model and include the 5-year German Bund yield. All variables enter the estimation in log levels if they are not already expressed in percentage terms. I estimate the impulse responses starting the sample in 1999 as I do in section 1.2.

Second, I identify the euro area conventional monetary policy and forward guidance shocks by combining the high-frequency proxies with the zero, magnitude, and sign restrictions described in the text. Again I use the same variables and transformations as in step 1. As stated in the main text for this estimation I start the sample in 2002 to take into account the liquidity concerns raised by Altavilla et al. (2019), which are particularly severe for the high-frequency changes in the longer run maturities.

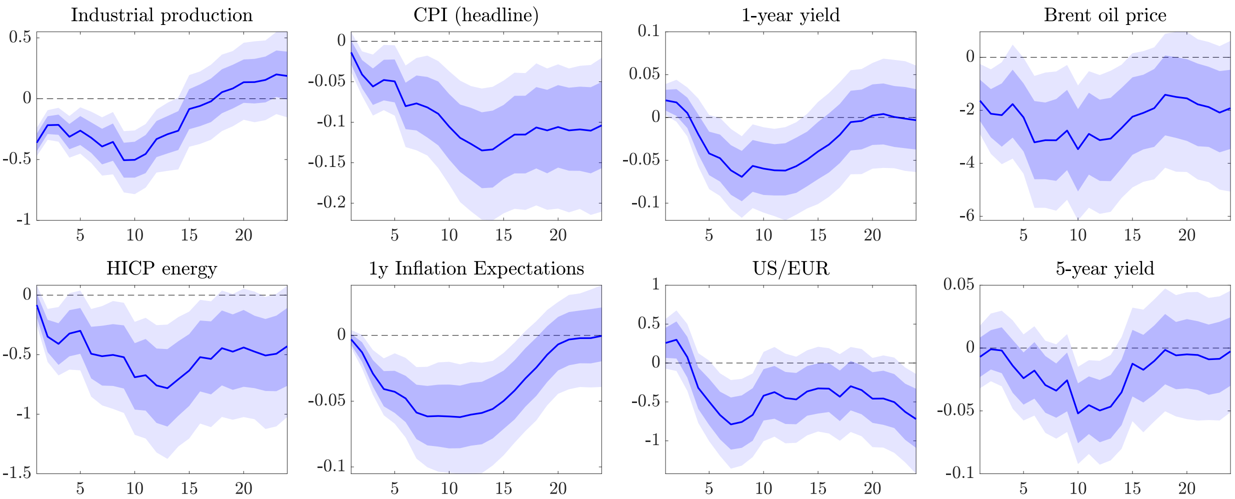
Third, and in contrast to the procedure in Appendix A.6, I use the same endogenous variables and sample as in the first step to estimate the impulse responses to a short- and long-run oil supply news shock in line with the description in Section 1.3.2. Again I start the sample in 1999 as I do not need to take into account the aforementioned liquidity concerns for the interest rate futures. The impulse responses to the identified shocks are plotted in Figures A.8 and A.9. The differences of the estimated responses to those depicted in Figures A.3 and A.4 arise because I additionally included the 5-year yield.

Fourth, I compute the posterior distribution of each of the counterfactual impulse responses, where the euro area monetary policy shocks from the second step do not affect the oil price. I do so by applying the procedure of McKay and Wolf (2023) to each draw from the posterior distribution of the second alongside the corresponding draws for the oil price shocks from the third step.

Fifth, I condition the impulse responses from the first step and compute the optimal policy counterfactual for each draw from the posterior distribution of the fourth step.

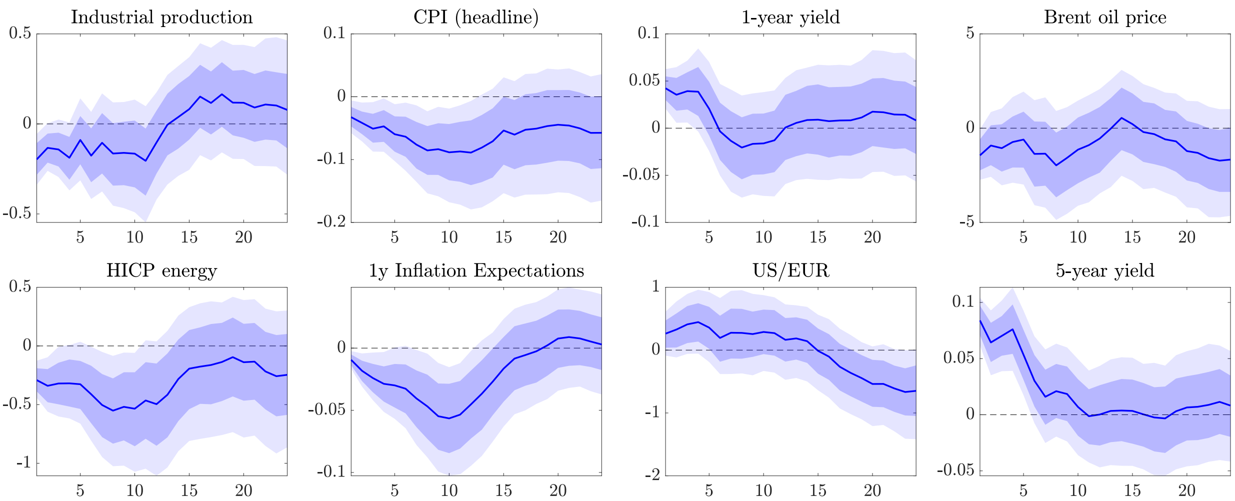
Lastly, I plot the point-wise mean which can be interpreted as summarizing the posterior distribution of impulse responses under the optimal (counterfactual) policy response conditional on the data and the impulse responses from step 1.

Figure A.6: IRFs to a contemporaneous EA Monetary policy shock when jointly identified alongside a forward guidance shock



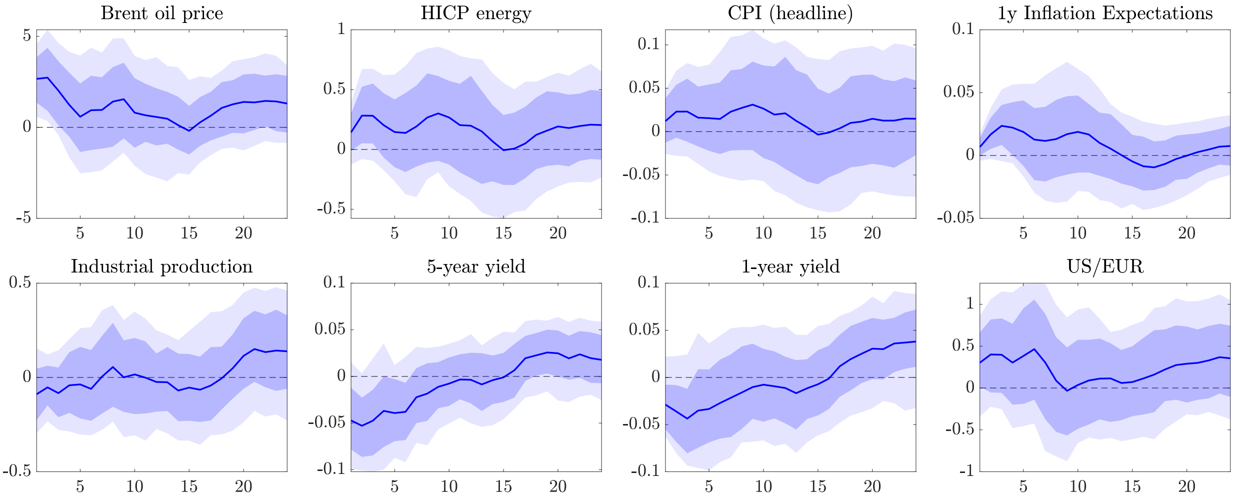
Notes: Impulse response functions to a one standard deviation conventional monetary policy shock showing the point-wise posterior means along with 68% and 90% point-wise credible sets in blue. Because I add the 5-year yield to the SVAR model, start the sample in 2002 and jointly identify a contemporaneous monetary policy and a forward guidance shock, the estimated IRFs differ somewhat from the ones presented in Figure 1.1, where I only identify a contemporaneous monetary policy shock.

Figure A.7: IRFs to an EA Forward Guidance Shock when jointly identified alongside a monetary policy shock



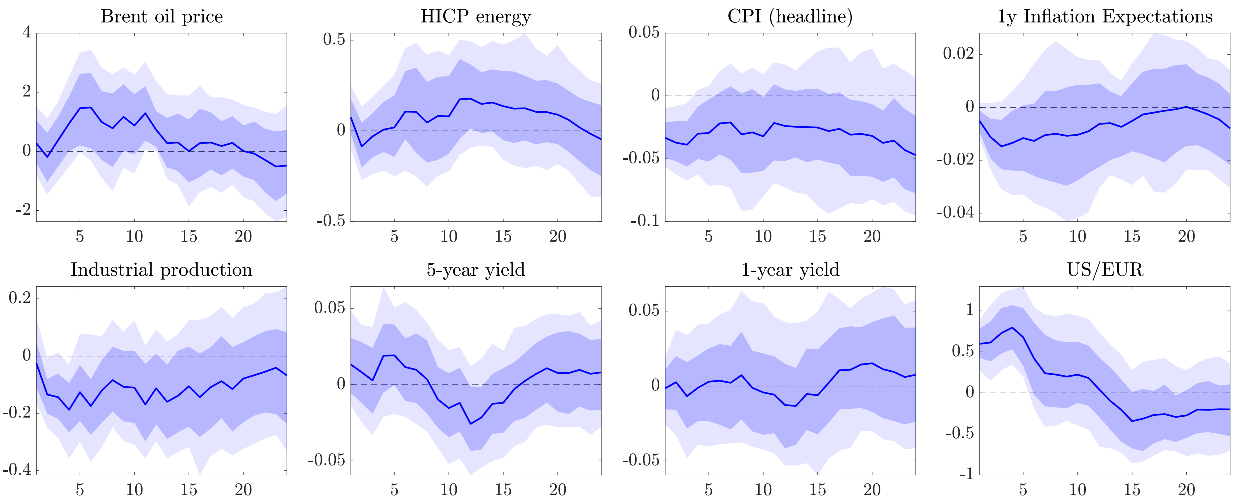
Notes: Impulse response functions to a one standard deviation forward guidance shock showing the point-wise posterior means along with 68% and 90% point-wise credible sets in blue.

Figure A.8: IRFs to a short-run oil supply news shock when including the 5-year yield



Notes: Impulse response functions to a one standard deviation short-term oil supply news shock showing the point-wise posterior means along with 68% and 90% point-wise credible sets in blue.

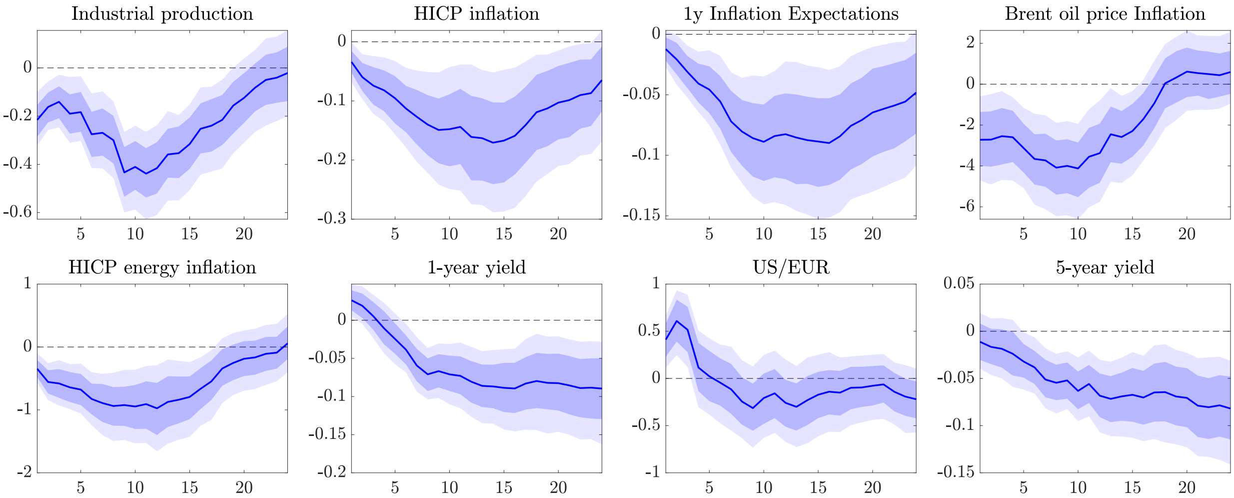
Figure A.9: IRFs to a medium-run oil supply news shock when including the 5-year yield



Notes: Impulse response functions to a one standard deviation medium-term oil supply news shock showing the point-wise posterior means along with 68% and 90% point-wise credible sets in blue.

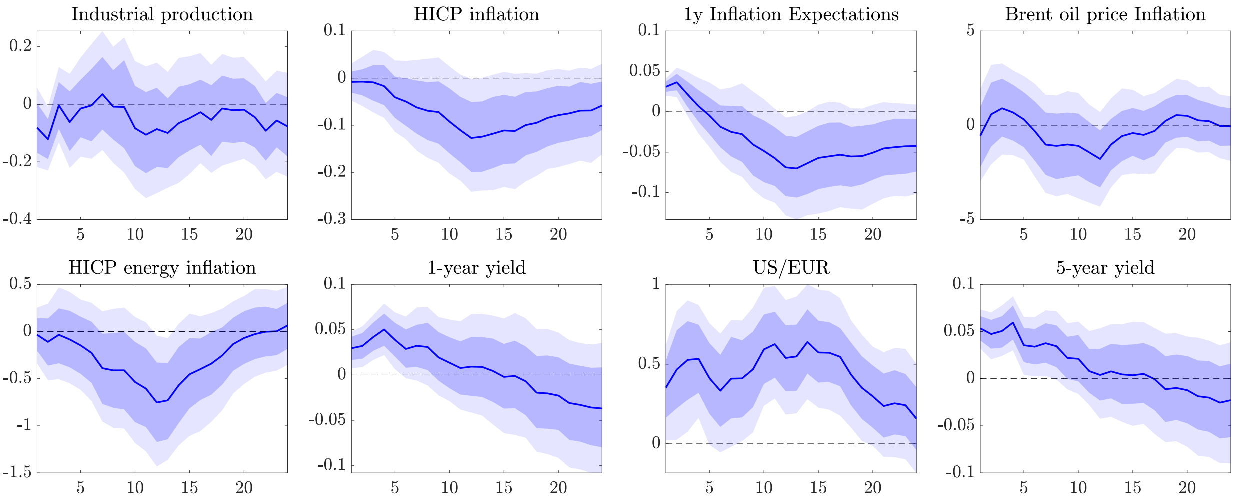
A.7 IRFs to monetary policy shocks when including the pandemic into the estimation

Figure A.10: IRFs to a contemporaneous EA Monetary policy shock when estimating the BPSVAR using data October 2023



Notes: Impulse response functions to a one standard deviation conventional monetary policy shock showing the point-wise posterior means along with 68% and 90% point-wise credible sets in blue. I extended the dataset for the estimation until 2024, which includes the pandemic. I explicitly model the Pandemic using the Pandemic Priors approach of Cascaldi-Garcia (2022). Furthermore, relative to the baseline, I add the 5-year yield to the SVAR model, start the sample in 2002 and jointly identify a contemporaneous monetary policy and a forward guidance shock. Lastly I incorporate y-o-y inflation rates for prices instead of their level to ensure stationarity.

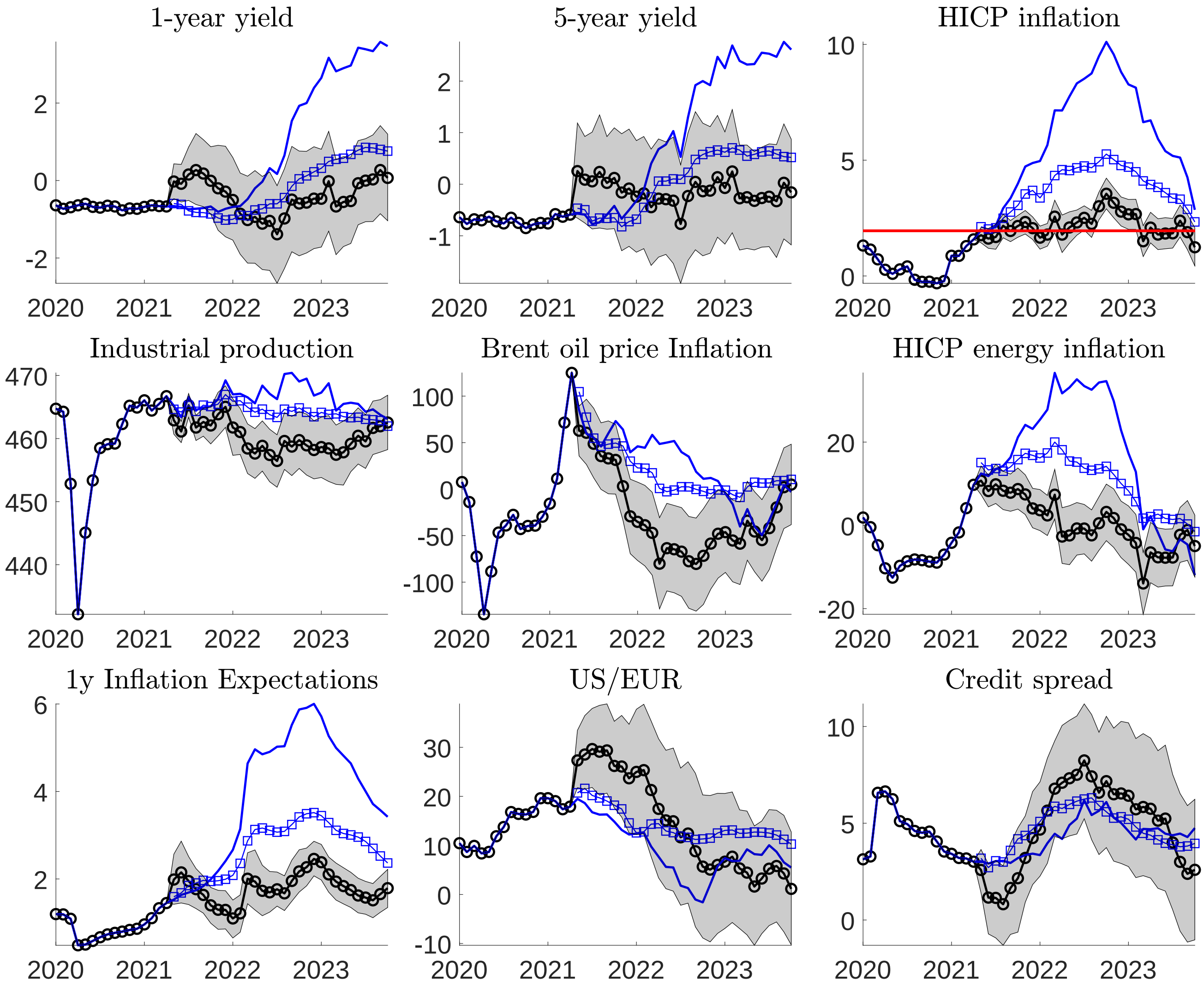
Figure A.11: IRFs to a EA forward guidance when estimating the BPSVAR using data until October 2023



Notes: See notes to Figure A.10

A.8 Further material for the counterfactual evolution of the economy under optimal policy

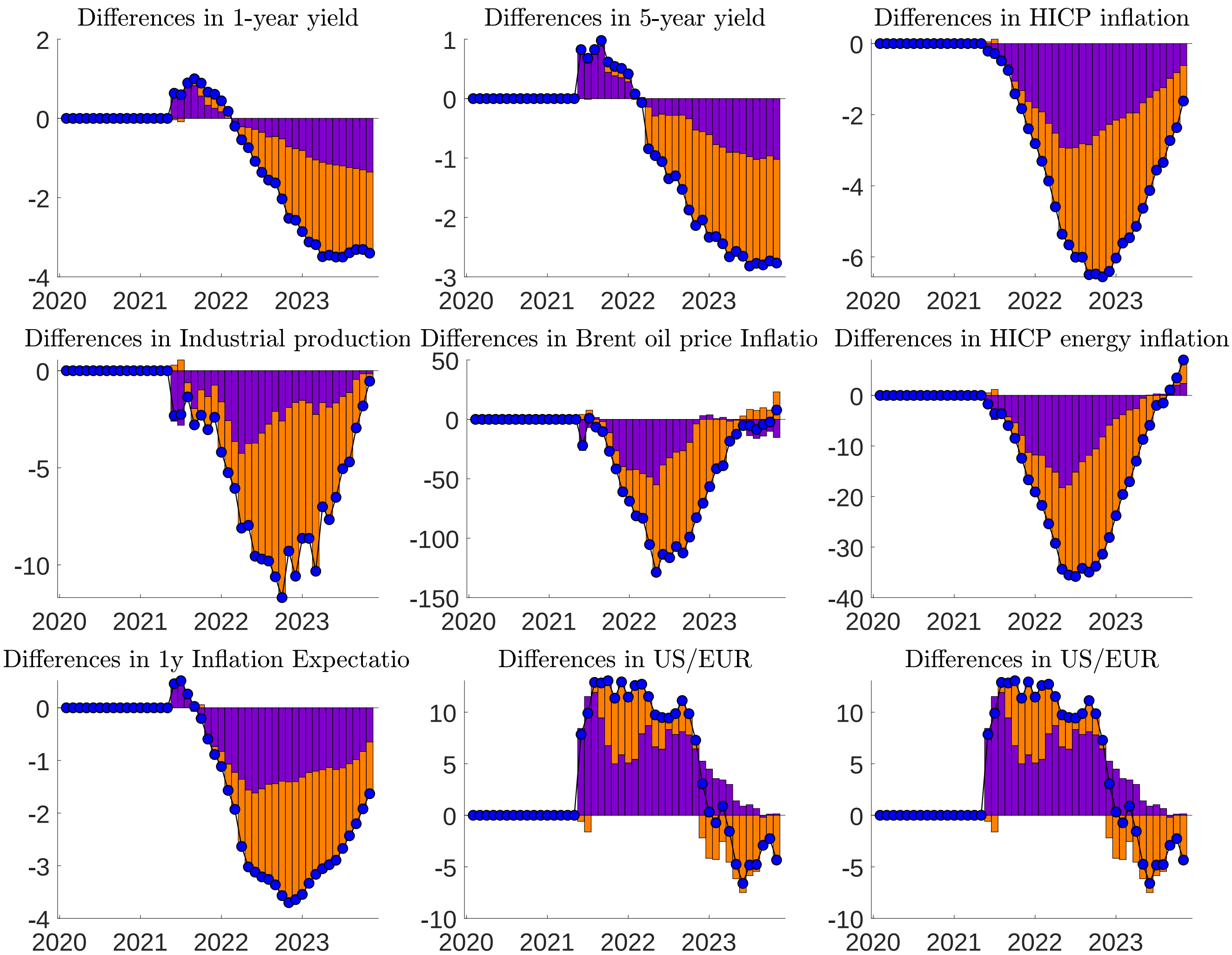
Figure A.12: Counterfactual and expected evolution of the endogenous variables under optimal ECB policy



Notes: The figure plots the time series of the endogenous variables (blue solid line) alongside the estimated counterfactual evolution (black circled line) of these variables under the assumption that, from April 2021 onwards, the ECB would have conducted optimal monetary policy as described in Equation 1.19 and thereby minimizes the weighted squared deviations of "medium-term" inflation from the target. Furthermore, the blue solid line with squares represents the expected evolution (forecast) of the economy under the assumption that the ECB follows the baseline policy rule also from April 2021 onwards. I refer to this as the deterministic or expected component. The difference between the blue lines is therefore driven by surprises and represents the stochastic component.

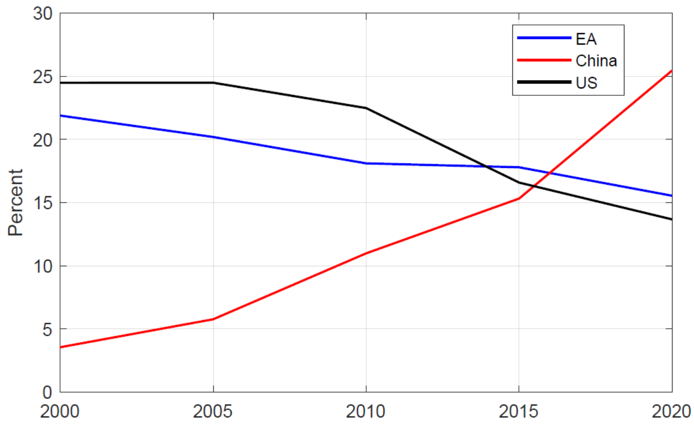
A.9 Oil import shares

Figure A.13: Counterfactual and expected evolution of the endogenous variables under optimal ECB policy



Notes: The Figure decomposes the difference between the observed evolution of the economy (blue lines in Figure 1.8) and the counterfactual evolution of the economy (black lines in Figure 1.8). These differences arise because of the change in the policy rule from the baseline to the optimal policy rule, which I assume to happen in April 2021. They can be decomposed into differences in deterministic components (i.e. the expected path of the endogenous variables without any further shocks) which are depicted in purple and differences in the stochastic component (i.e. the way that new, unexpected shocks propagate) which are depicted in orange.

Figure A.14: Oil import shares



Notes: Figure represents the global share of crude oil imports of the euro area, China and the United States at five-year intervals from 2000 to 2020. The global share is computed as a percentage of total crude oil imports in the world using data from the International Energy Agency (IEA).

Appendix B

Appendix for Chapter 2

B.1 Discussion of the Bayesian approach to identification with external instruments

The BPSVAR framework has several appealing features relative to traditional frequentist external instrument SVAR models that render it particularly well-suited for the purpose of estimating the effects of US monetary policy and global uncertainty shocks on the world economy. First, it requires relatively weak additional identifying assumptions when more than one structural shock is to be identified by proxy variables. In this case, the shocks are only set identified as rotations of the structural shocks $\boldsymbol{Q}\boldsymbol{\epsilon}_t^*$ with orthonormal matrices $\boldsymbol{Q}$ also satisfy the relevance and exogeneity conditions in Equations (2a) and (2b) in the paper Therefore, additional restrictions are needed in order to point-identify the structural shocks in $\boldsymbol{\epsilon}_t^*$. In the frequentist external instruments VAR model these additional restrictions are imposed on the contemporaneous relationships between the *endogenous variables* $\boldsymbol{y}_t$ reflected in $\boldsymbol{A}_0^{-1}$ (Mertens and Ravn 2013; Lakdawala 2019). However, Arias et al. (2021) show that relaxing this type of additional identifying assumptions can change the results profoundly. Instead, the BPSVAR framework allows us to impose the additional identifying assumptions on the contemporaneous relationships between the *structural shocks* $\boldsymbol{\epsilon}_t^*$ *and proxy variables* $\boldsymbol{m}_t$ reflected in $\boldsymbol{V}$ in the relevance condition in Equation (2a). For example, we can impose the restriction that a particular structural shock does not affect a particular proxy variable. Restrictions on the contemporaneous relationships are arguably weaker for structural shocks and proxy variables in $\boldsymbol{V}$ than for the endogenous variables in $\boldsymbol{A}_0^{-1}$.

Second, the BPSVAR framework allows coherent and exact finite sample inference, even in settings in which the proxy variables are weak instruments and only set rather than point identification is achieved with a combination of sign, magnitude and zero restrictions (see Moon and Schorfheide 2012; Caldara and Herbst 2019; Arias et al. 2021). In particular, frequentist external instruments VAR models are estimated in a two-step procedure (Mertens and Ravn 2013; Gertler and Karadi 2015): (i) estimate the reduced-form VAR model; (ii) regress the reduced-form residuals on the proxy variable to obtain the structural parameters.

This two-step procedure is inefficient, as the estimation of the reduced-form VAR model in (i) is not informed by the proxy variable. In contrast, the BPSVAR model considers the joint likelihood of the endogenous variables and the proxy variables based on Equation (3) in the paper, so that the proxy variables inform the estimation of both reduced-form and structural parameters. The BPSVAR framework also facilitates inference, as the joint estimation captures all sources of uncertainty. Furthermore, as long as the prior distribution is proper, in a Bayesian setting inference is straightforward even when the instruments are weak (Poirier 1998). By contrast, frequentist external instruments VAR models require an explicit theory to accommodate weak instruments (Montiel Olea et al. 2021), either to derive the asymptotic distributions of the estimators or to ensure satisfactory coverage in bootstrap algorithms.[90]

Third, from Equation (7) it can be seen that the BPSVAR framework is relatively flexible in that it allows for the proxy variables to be serially correlated and to be affected by lags of the endogenous variables as well as by measurement error. This is a useful feature as it has been shown that some widely-used proxy variables are serially correlated and/or contaminated by measurement error (Miranda-Agrippino and Ricco 2021). In these cases, it is typically proposed to cleanse the proxy variables in an additional step preceding the analysis in the VAR model, exacerbating issues regarding efficiency and coherent inference.

And fourth, the BPSVAR model allows us to incorporate a prior belief about the strength of the proxy variables as instruments based on the notion that "researchers construct proxies to be relevant" (Caldara and Herbst 2019, p. 165). In particular, consider the 'reliability matrix' $\boldsymbol{R}$ derived in Mertens and Ravn (2013) given by

$$\boldsymbol{R} = \left(\boldsymbol{\Gamma}_{0,2}^{-1\prime}\boldsymbol{\Gamma}_{0,2} + \boldsymbol{V}\boldsymbol{V}'\right)^{-1}\boldsymbol{V}\boldsymbol{V}'. \tag{B.1}$$

Intuitively, $\boldsymbol{R}$ indicates the share of the total variance of the proxy variables that is accounted for by the structural shocks $\boldsymbol{\epsilon}_t^*$ (see Equation (7)). Specifically, the minimum eigenvalues of $\boldsymbol{R}$ can be interpreted as the share of the variance of (any linear combination of) the proxy variables explained by the structural shocks $\boldsymbol{\epsilon}_t^*$ (Gleser 1992).

B.2 Additional Tables

[90] To the best of our knowledge, there is no consensus yet on how to conduct inference in frequentist external instruments VAR models, even in a setting with only a single proxy variable (Jentsch and Lunsford 2019).

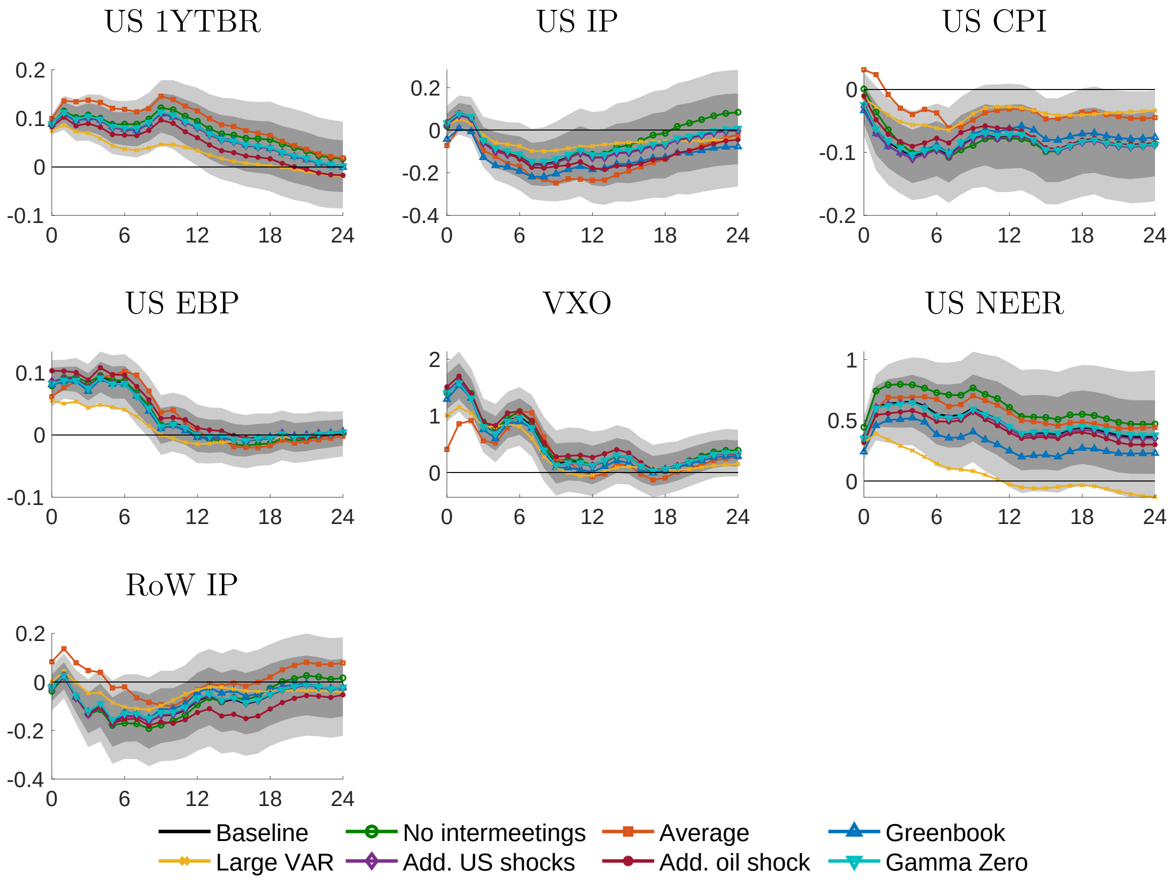
Table B.1: Data description

Variable	Description	Source	Coverage
US 1-year TB rate	1-year Treasury Bill yield at constant maturity	US Treasury/Haver	1990m1 - 2019m6
US IP	Industrial production excl. construction	FRB/Haver	1990m1 - 2019m6
US CPI	Consumer price index	BLS/Haver	1990m1 - 2019m6
US EBP	Excess bond premium	See Favara et al. (2016)	1990m1 - 2019m6
VXO	CBOE market volatility index VXO	Wall Street Journal/Haver	1990m1 - 2019m6
S&P 500	S&P 500 Composite	S&P/Haver	1990m1 - 2019m6
US PPI	PPI finished goods	BLS/Haver	1990m1 - 2019m6
US import prices	Import price index: All imports	BLS/Haver	1990m1 - 2019m6
US import prices excl. petroleum	Import price index: Non-petroleum imports	BLS/Haver	1990m1 - 2019m6
US consumption	Real personal consumption expenditures (chnd. 2012$)	BEA/Haver	1990m1-2019m6
US investment	Gross private domestic investment (chnd. 2012$)	BEA/Haver	1990q1-2019q2, interpolated to monthly frequency
US exports	Exports of goods and services (chnd. 2012$)	BEA/Census Bureau/Haver	1990m1-2019m6
US imports	Imports of goods and services (chnd. 2012$)	BEA/Census Bureau/Haver	1990m1-2019m6
US dollar NEER	Nominal broad trade-weighted dollar index	FRB/Haver	1990m1-2019m6
CFED-1Y real interest rate	See Haubrich et al. (2012)	Cleveland Fed	1990m1 - 2019m6
Oil prices	European Brent spot price ($ per barrel)	EIA/Haver	1990m1 - 2019m6
Dow Jones World	Dow Jones Global Index	Dow Jones/Haver	1992m1 - 2019m6
Dow Jones excl. US	Dow Jones Global Index excl. US	Dow Jones/Haver	1992m1 - 2019m6
MSCI AEs	MSCI AEFE Index: Developed markets in Europe, Australasia, Israel and the Far East	MSCI/Bloomberg	1990m1 - 2019m6
MSCI EMEs	MSCI MXEF Index: Emerging markets with mid to large cap	MSCI/Bloomberg	1990m1 - 2019m6
S&P 500 earnings expectations	S&P 500 Composite 12-months forward earnings per share	S&P/Bloomberg	1990m1 - 2019m6
S&P 500 low/high RoW exposure	Based on sectoral S&P 500 indices	S&P/Haver	1990m1 - 2019m6
RoW, AE, EME IP	Industrial production, see Martínez-García et al. (2015)	Dallas Fed Global Economic Indicators/Haver	1990m1 - 2019m6
RoW, AE, EME CPI	Consumer price index	Dallas Fed Global Economic Indicators/Haver (Martínez-García et al. 2015)	1990m1 - 2019m6
RoW, AE, EME policy rate	Short-term official/policy rate, see Martínez-García et al. (2015)	Dallas Fed Global Economic Indicators/Haver	1990m1 - 2019m6
Singapore IP	Industrial production: Manufacturing (excl. rubber processing)	Department of Statistics/Haver	1990m1 - 2019m6
Taiwan IP	Industrial production: Manufacturing	Ministry of Economic Affairs/Haver	1990m1 - 2019m6
Israel IP	Industrial production: Manufacturing	Central Bureau of Statistics/Haver	1990m1 - 2019m6

Notes: BLS stands for Bureau of Labour Statistics, FRB for Federal Reserve Board, BEA for Bureau of Economic Analysis, and EIA for Energy Information Administration.

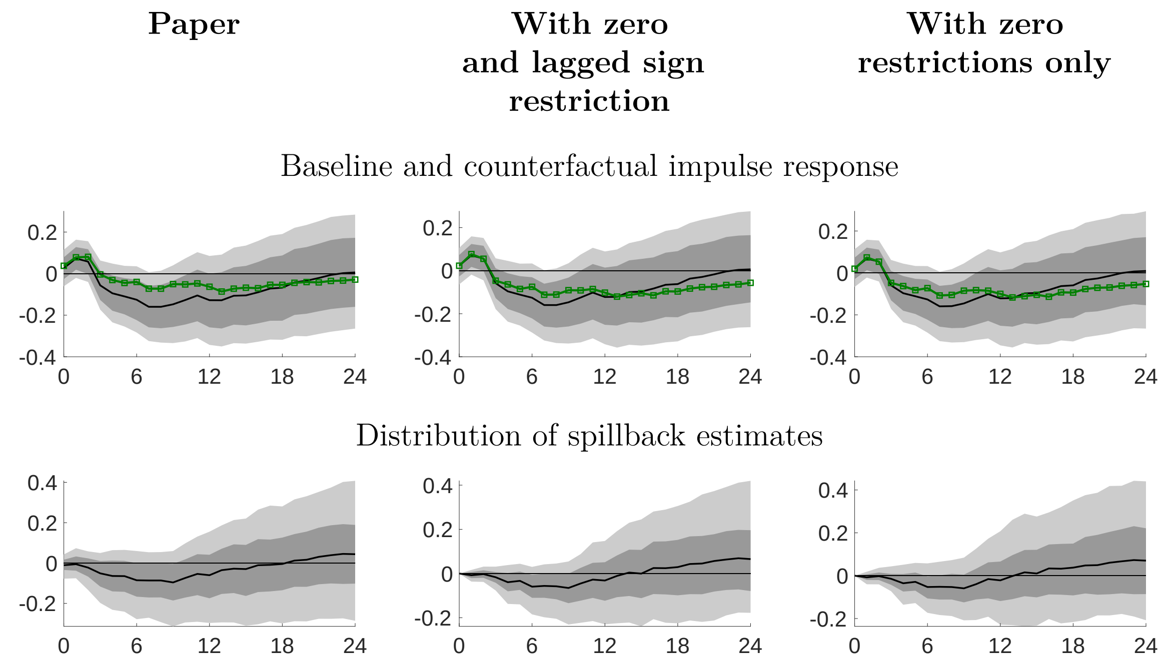
B.3 Additional Figures

Figure B.1: Impulse responses to US monetary policy shocks from alternative BPSVAR specifications



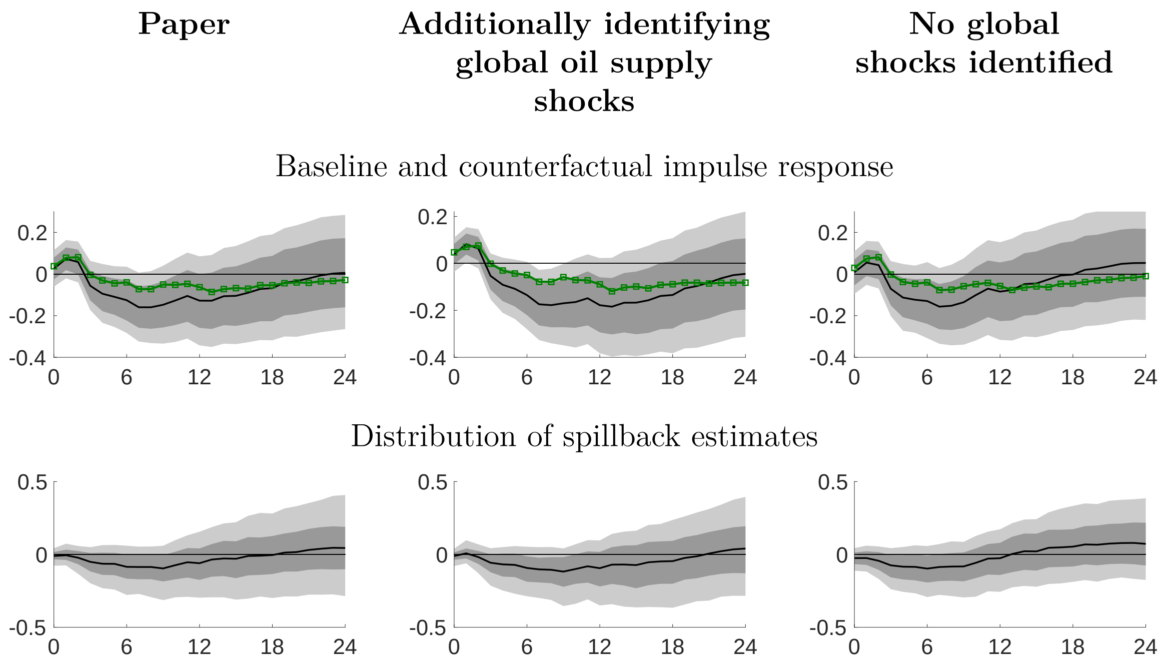
Notes: The figure shows point-wise posterior mean impulse responses from our baseline specification (black solid lines) together with 68% and 90% centered point-wise probability bands (grey areas) and various robustness checks (coloured solid lines with markers). '1YTBR' stands for the one-year Treasury Bill rate, 'IP' for industrial production, 'CPI' for consumer-price index, 'EBP' for excess bond premium, 'VXO' is the S&P 500 stock market volatility index, and 'NEER' the nominal effective exchange rate. In the specification labelled 'No intermeetings' we exclude FOMC announcements on unscheduled, 'inter-meeting' dates; in 'Average' we temporally aggregate to monthly frequency the interest rate and gold price surprises originally available at daily frequency by taking monthly averages as in Jarocinski and Karadi (2020) rather than as in Gertler and Karadi (2015); in 'Greenbook' we purge the interest rate surprises from Green Book forecast as suggested in Miranda-Agrippino and Ricco (2021) instead of applying the 'poor-man's' approach of Jarocinski and Karadi (2020); in 'Large VAR' we add RoW consumer prices, AEs policy rate, US imports and exports, and the MSCI world stock price index to the vector of endogenous variables in the VAR model; in 'Add. US shocks' we additionally identify US demand and supply shocks with standard sign restrictions (we impose on impact: negative demand shocks affect US IP, CPI, NEER, and 1YTBR negatively and US EBP positively; negative supply shocks affect US IP negatively and US CPI and EBP positively); in 'Add. oil shock' we additionally identify global oil supply shocks using the proxy variable of Känzig (2021); and in 'Gamma Zero' we set γ—the 'relevance threshold' of the proxy variables—to zero.

Figure B.2: Comparison of 'RoW shocks SSA' with alternative identification assumptions on real activity spillovers to US from RoW shocks



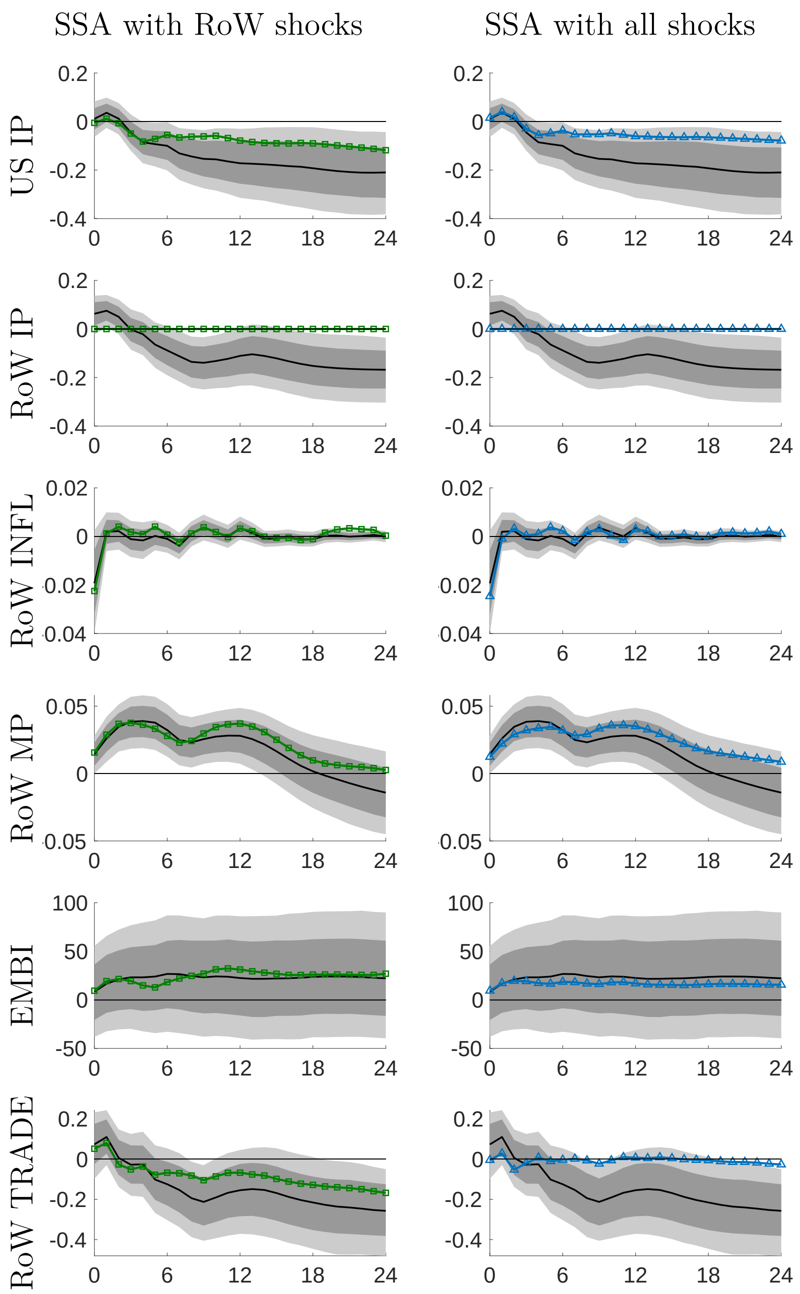
Notes: In the first row, the black solid lines depict point-wise posterior means of the baseline impulse responses of US industrial production to a US monetary policy shock and the green lines with squares the point-wise posterior means of the counterfactual impulse responses based on SSA with RoW shocks. The grey shaded areas represent 68% and 90% centered point-wise probability bands for the baseline impulse responses. In the second row, the black solid lines depict point-wise posterior means of the differences between the baseline and the counterfactual effects of US monetary policy on US industrial production together with 68% and 90% centered point-wise probability bands. The first column presents the results from the specification in the paper. In the second column, we report results from an alternative specification in which we use RoW shocks identified assuming they have no contemporaneous effect on US real activity and imposing the corresponding sign restrictions from Table 1 in the manuscript only with a one-month lag. In the third column, we report results from an alternative specification in which we use RoW shocks identified assuming they have no contemporaneous effect on US real activity and in which we impose no sign restriction at all.

Figure B.3: Comparison of SSA estimates with RoW shocks imposing alternative identification restrictions on the global shocks



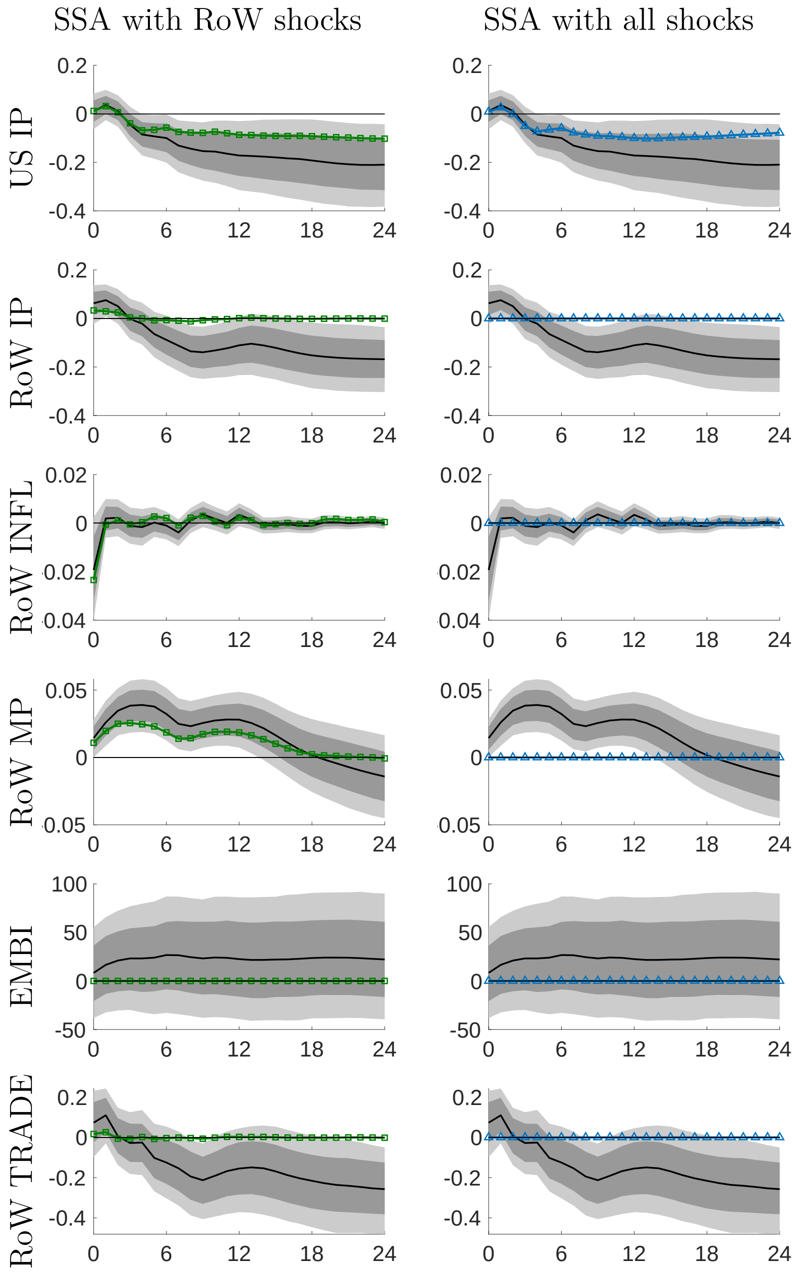
Notes: In the first row, the black solid lines depict point-wise posterior means of the baseline impulse responses of US industrial production to a US monetary policy shock and the green lines with squares the point-wise posterior means of the counterfactual impulse responses based on SSA with RoW shocks. The grey shaded areas represent 68% and 90% centered point-wise probability bands for the baseline impulse responses. In the second row, the black solid lines depict point-wise posterior means of the differences between the baseline and the counterfactual effects of US monetary policy on US industrial production together with 68% and 90% centered point-wise probability bands. The first column presents the results from the specification in the paper. In the second column, we report results from an alternative specification in which we additionally identify global oil supply shocks using high-frequency oil futures surprises around OPEC announcements constructed by Känzig (2021), and in the third column from an alternative specification in which we identify neither the global uncertainty nor the oil price shock.

Figure B.4: Baseline and counterfactual impulse responses to a US monetary policy shock from a VAR specification with additional RoW variables



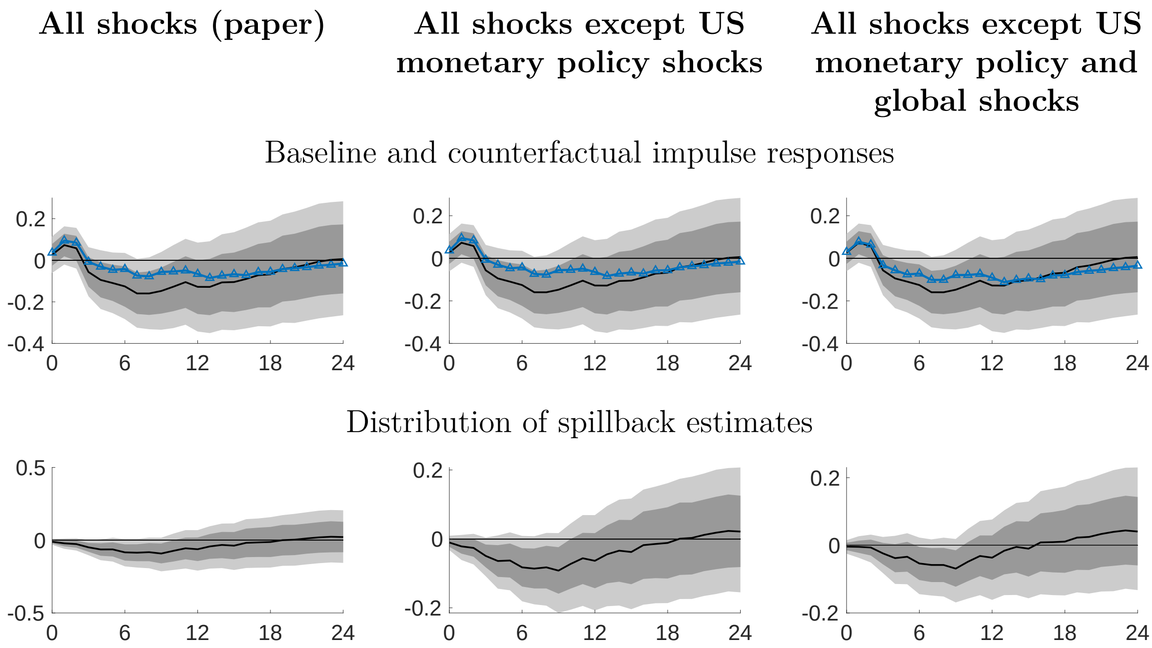
Notes: The black solid lines depict the baseline impulse responses of US industrial production to a US monetary policy shock and the coloured solid lines with symbol markers the counterfactual impulse responses based on SSA with RoW shocks (left column, green lines with squares), based on SSA with all shocks (middle column, blue lines with triangles), and based on MRE (right column, orange lines with circles). The grey shaded areas represent 68% and 90% centered point-wise probability bands for the baseline impulse responses. 'US IP' stands for US industrial production, 'RoW IP' for RoW industrial production, 'RoW INFL' for RoW inflation rate, 'RoW MP' for RoW monetary policy rate, 'EMBI' for the EMBI spread, and 'RoW TRADE' for RoW gross trade. Due to the significantly larger dimensionality of the VAR we use informative priors in the estimation (Giannone et al. 2015).

Figure B.5: Baseline and counterfactual impulse responses to a US monetary policy shock from a VAR specification with additional RoW variables and additional counterfactual constraints on RoW variables



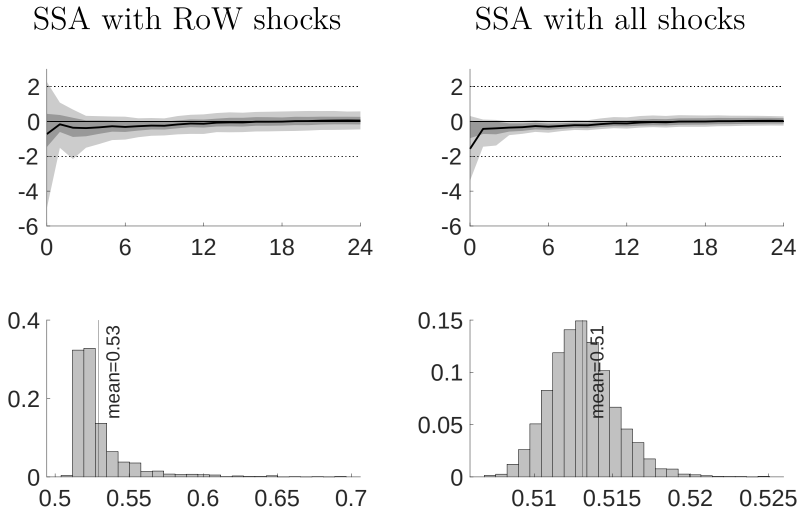
Notes: See the note Figure B.4. As we only use two shocks in case of SSA with RoW shocks in column (1) the no-spillovers constraint cannot be satisfied for all RoW variables in the counterfactual. Instead, when the number of variables that are constrained in the counterfactual is greater than the number of shocks used to enforce the counterfactual, SSA minimises a weighted sum of deviations of the impulse responses of all RoW variables from the counterfactual constraint. As MRE in column (3) is implemented as a numerical optimisation problem and as the number of parameters—the number of variables for which the impulse responses are constrained to be zero at each horizon—is larger in this exercise, at some horizons the MRE solution is unreliable (see the kinks in case of, for example, the counterfactual response of US industrial production in the third column).

Figure B.6: SSA with different sets of offsetting shocks



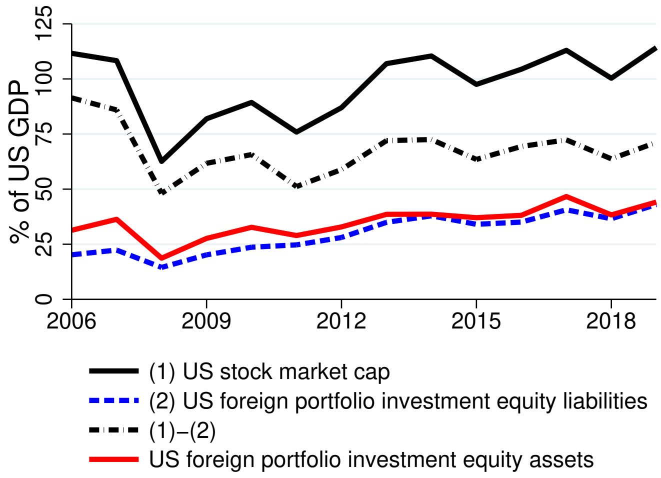
Notes: The first row shows the baseline (black solid lines) and counterfactual impulse responses (blue lines with triangles) based on SSA using as offsetting shocks all shocks as in the paper (first column), all shocks except US monetary policy shocks (second column), all shocks except US monetary policy and global shocks (third column). The grey shaded areas represent 68% and 90% centered point-wise probability bands for the baseline impulse responses. In the second row, the black solid lines show the point-wise mean of the differences between the baseline and the counterfactual effects together with 68% and 90% centered point-wise probability bands.

Figure B.7: Modesty statistic of Leeper and Zha (2003) and distribution of the q-divergence of Antolin-Diaz et al. (2021) for SSA counterfactuals



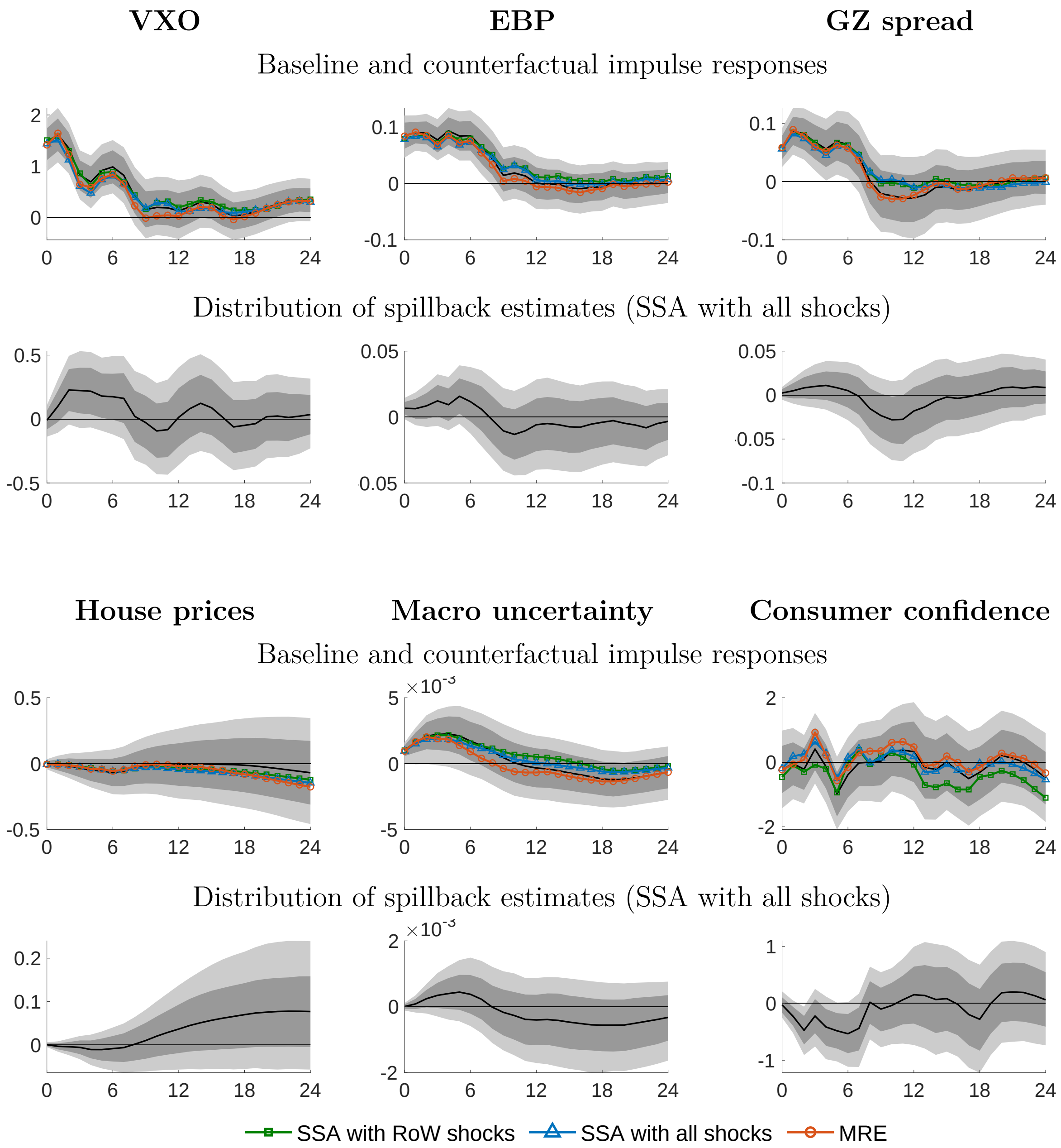
Notes: The top panels show the 'modesty statistic' of Leeper and Zha (2003) for the implied offsetting shocks that impose the counterfactual constraint for RoW industrial production; the black solid line depicts the point-wise mean and the grey shaded areas the 68% and 90% centered point-wise probability bands. The offsetting shocks are 'modest'—meaning their materialisation is unlikely to induce agents to adjust their expectation formation and beliefs about the structure of the economy—if the statistic is smaller than two in absolute value. The bottom panels show the distribution of the q-divergence of Antolin-Diaz et al. (2021) for the SSA; the left-hand side panel presents results for the case in which only the RoW shocks are used as offsetting shocks, while the right-hand side panel for the case in which all shocks are used. The q-divergence indicates how unlikely a conditional forecast is in terms of comparing the implied distributions of shocks with their unconditional distributions, translated into a comparison of the binomial distributions of a fair and a biased coin. See Appendix B.4.1 for a description how we implement the q-divergence. We drop SSA counterfactuals when the offsetting shocks increase over time ($\tilde{\epsilon}_{T+1} < \tilde{\epsilon}_{T+h}$) or if the offsetting shocks are particularly large on impact ($\tilde{\epsilon}_{T+1}$ lies above the 99^{th} percentile in absolute terms).

Figure B.8: US stock market capitalisation and US foreign portfolio investment equity assets and liabilities



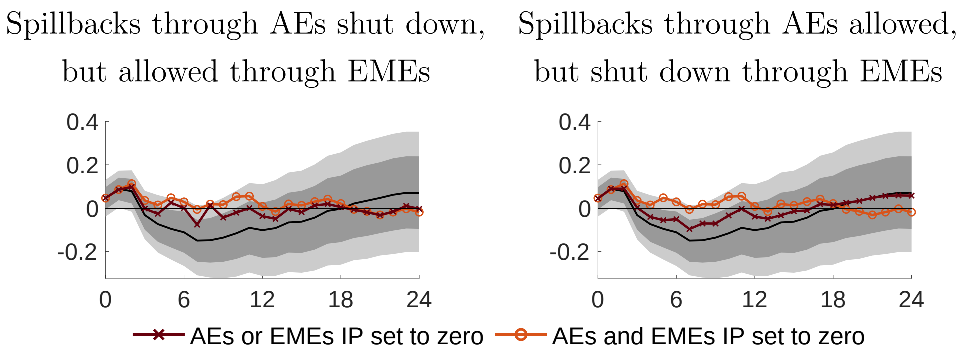
Notes: The figure shows the evolution of US (New York Stock Exchange) stock market capitalisation depicted by the black solid line as well as US foreign portfolio investment equity assets and liabilities depicted by the solid red and dashed blue lines, respectively; the black dashed line depicts US stock market capitalisation less US foreign portfolio investment equity liabilities. All variables are depicted as a percentage of US nominal annual GDP. The data on stock market capitalisation are obtained from the World Federation of Exchanges, those on US foreign equity assets and liabilities from the Bureau of Economic Analysis.

Figure B.9: Channels of transmission for spillbacks from US monetary policy: Additional variables



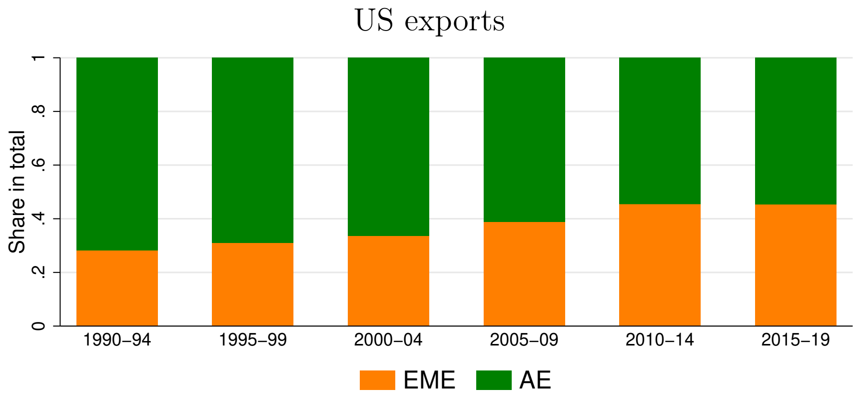
Notes: In the first and third row the figure shows the baseline and counterfactual impulse responses based on SSA with RoW shocks (green lines with squares), SSA with all shocks (blue lines with triangles) and MRE (orange lines with circles) for the VXO, the excess bond premium, the Gilchrist-Zakrajsek (GZ) spread, the S&P CoreLogic Case-Shiller home price index, the macroeconomic uncertainty index of Jurado et al. (2015), and the Conference Board consumer confidence index. The grey shaded areas represent 68% and 90% centered point-wise probability bands for the baseline impulse responses. In the second and fourth row the figure shows the point-wise mean of the differences between the baseline and the counterfactual effects together with 68% and 90% centered point-wise probability bands.

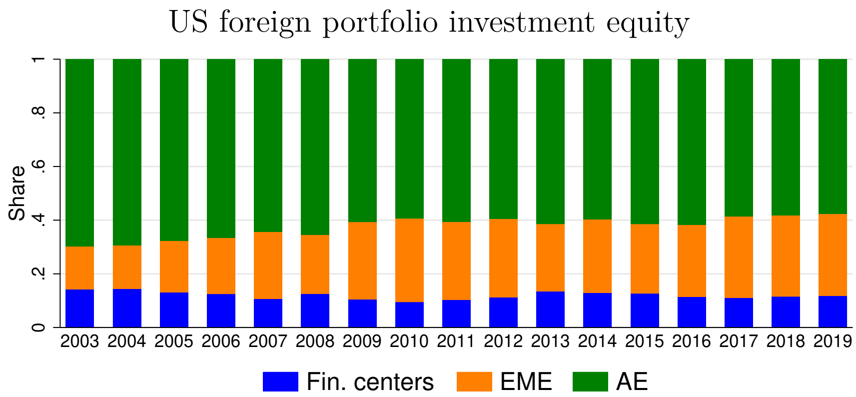
Figure B.10: Spillbacks from US monetary policy through AEs and EMEs (MRE counterfactuals)



Notes: The black solid lines depict the response of US industrial production to a US monetary policy shock from VAR models in which RoW industrial production is replaced with separate measures for AEs and EMEs industrial production and their impulse responses to a US monetary policy shock are not constrained; the grey shaded areas represent the associated 68% and 90% centered point-wise probability bands. The light orange lines with circles depict counterfactual impulse responses in which spillovers to the RoW are precluded, meaning that both AEs and EMEs industrial production are constrained to not respond to a US monetary policy shock. In the left-hand side panel the dark red line with crosses depicts the counterfactual impulse response of US industrial production when AEs industrial production is constrained to not respond while EMEs industrial production is set to respond as in the unconstrained case. In the right-hand side panel the dark red line with crosses depicts the counterfactual impulse response when AEs industrial production is constrained to respond as in the unconstrained case while EMEs industrial production is constrained to not respond to a US monetary policy shock. All counterfactual impulse responses shown are based on MRE. We cannot carry out this exercise for SSA with RoW shocks as in this specification we do not identify shocks that are common to both AEs and EMEs industrial production.

Figure B.11: Country composition of US foreign portfolio investment equity holdings and US exports





Notes: The top panel shows the country composition of US exports of goods obtained from the IMF Direction of Trade Statistics as a proxy for the composition of US firms' sales to the RoW. The bottom panel shows the country composition of US foreign portfolio equity holdings through common stocks and mutual funds based on the analysis in Bertaut et al. (2019), which accounts for measurement problems related to multinationals, financial centres and mutual funds. The total underlying the shares does not include cross-border equity of firms that primarily operate in the US (for details see Bertaut et al. 2019). The list of financial centres is taken from Bertaut et al. (2019).

B.4 The SSA framework of Antolin-Diaz et al. (2021)

Building on the work of Waggoner and Zha (1999), the SSA framework of Antolin-Diaz et al. (2021, henceforth ADPRR) provides a rigorous and general treatment on how to impose specific paths on observables in a VAR model as conditional forecasts with and without constraints on the set of offsetting—or 'driving'—shocks. Denoting by $\boldsymbol{y}'_{T+1,T+h} \equiv [\boldsymbol{y}'_{T+1}, \boldsymbol{y}'_{T+2}, \ldots, \boldsymbol{y}'_{T+h}]$ the $1 \times nh$ vector that stacks the future values of the observables over an horizon of h periods, the SSA framework of ADPRR consists of obtaining the distribution

$$\widetilde{\boldsymbol{y}}_{T+1,T+h} \sim N(\boldsymbol{\mu}_y, \boldsymbol{\Sigma}_y), \tag{B.2}$$

where the $nh \times 1$ vector vector $\widetilde{\boldsymbol{y}}_{T+1,T+h}$ contains the values of all observables—i.e. both those whose paths are constrained and those whose paths are unconstrained—under the conditional forecast. The $nh \times 1$ vector $\boldsymbol{\mu}_y$ contains the corresponding means of the distribution of the observables in $\widetilde{\boldsymbol{y}}_{T+1,T+h}$ under the conditional forecast, and the $nh \times nh$ matrix $\boldsymbol{\Sigma}_y$ the associated uncertainty.

In the framework of ADPRR, structural scenarios involve

(i) 'conditional-on-observables forecasting', i.e. specifying paths for a subset of observables in $\boldsymbol{y}_{T+1,T+h}$ that depart from their unconditional forecast, and/or

(ii) 'conditional-on-shocks forecasting', i.e. specifying the subset of (and potentially a path for) the structural shocks $\boldsymbol{\epsilon}_{T+1,T+h}$ that are allowed to depart from their unconditional distribution to produce the specified path of the observables in (i);

Both the case in which the path of observables under (i) and the case in which the path of structural shocks under (ii) is constrained can be laid out based on Equation (B.2). The goal is to determine $\boldsymbol{\mu}_y$ and $\boldsymbol{\Sigma}_y$ such that the constraints under (i) and (ii) are satisfied simultaneously.

Assume the structural parameters of the VAR model are known. The future values of the observables are given by

$$\boldsymbol{y}_{T+1,T+h} = \boldsymbol{b}_{T+1,T+h} + \boldsymbol{M}'\boldsymbol{\epsilon}_{T+1,T+h}, \tag{B.3}$$

where the $nh \times 1$ vector $\boldsymbol{b}_{T+1,T+h}$ represents the deterministic component due to initial conditions and the autoregressive dynamics of the VAR model, and the $nh \times nh$ matrix $\boldsymbol{M}'$ the impact of future structural shocks.

Under (i), 'conditional-on-observables forecasting' can be written as

$$\overline{\boldsymbol{C}}\widetilde{\boldsymbol{y}}_{T+1,T+h} = \overline{\boldsymbol{C}}\boldsymbol{b}_{T+1,T+h} + \overline{\boldsymbol{C}}\boldsymbol{M}'\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h} \sim N(\overline{\boldsymbol{f}}_{T+1,T+h}, \overline{\boldsymbol{\Omega}}_f). \tag{B.4}$$

where $\overline{\boldsymbol{C}}$ is a $k_o \times nh$ selection matrix, the $k_o \times 1$ vector $\overline{\boldsymbol{f}}_{T+1,T+h}$ is the mean of the distribution of the observables constrained under the conditional forecast and the $k_o \times k_o$ matrix $\overline{\boldsymbol{\Omega}}_f$ the associated uncertainty. In turn, under (ii), 'conditional-on-shocks forecasting' can be written as

$$\boldsymbol{\Xi}\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h} \sim N(\boldsymbol{g}_{T+1,T+h}, \boldsymbol{\Omega}_g), \tag{B.5}$$

where $\boldsymbol{\Xi}$ is a $k_s \times nh$ selection matrix, the $k_s \times 1$ vector $\boldsymbol{g}_{T+1,T+h}$ the mean of the distribution of the shocks constrained under the conditional forecast and the $k_s \times k_s$ matrix $\boldsymbol{\Omega}_g$ the associated uncertainty.[91] Under invertibility we have

$$\begin{aligned} \boldsymbol{M}'^{-1}\widetilde{\boldsymbol{y}}_{T+1,T+h} &= \boldsymbol{M}'^{-1}\boldsymbol{b}_{T+1,T+h} + \widetilde{\boldsymbol{\epsilon}}_{T+1,T+h}, \\ \Xi\boldsymbol{M}'^{-1}\widetilde{\boldsymbol{y}}_{T+1,T+h} &= \Xi\boldsymbol{M}'^{-1}\boldsymbol{b}_{T+1,T+h} + \Xi\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h}, && \text{(B.6)} \\ \underline{\boldsymbol{C}}\widetilde{\boldsymbol{y}}_{T+1,T+h} &= \underline{\boldsymbol{C}}\boldsymbol{b}_{T+1,T+h} + \Xi\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h}, && \text{(B.7)} \end{aligned}$$

and hence

$$\underline{\boldsymbol{C}}\widetilde{\boldsymbol{y}}_{T+1,T+h} = \underline{\boldsymbol{C}}\boldsymbol{b}_{T+1,T+h} + \Xi\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h} \sim N(\underline{\boldsymbol{f}}_{T+1,T+h}, \underline{\boldsymbol{\Omega}}_f), \tag{B.8}$$

with $\underline{\boldsymbol{\Omega}}_f = \boldsymbol{\Omega}_g$.

Based on Equations (B.4) and (B.8), we can combine the k_o constraints on the observables under 'conditional-on-observables forecasting' and the k_s constraints on the structural shocks under 'conditional-on-shocks forecasting' by defining the $k \times nh$, $k = k_o + k_s$, matrices $\boldsymbol{C} \equiv [\overline{\boldsymbol{C}}', \underline{\boldsymbol{C}}']'$ and $\boldsymbol{D} \equiv [\boldsymbol{M}\overline{\boldsymbol{C}}', \Xi']'$ to write

$$\boldsymbol{C}\widetilde{\boldsymbol{y}}_{T+1,T+h} \;=\; \boldsymbol{C}\boldsymbol{b}_{T+1,T+h} + \boldsymbol{D}\widetilde{\epsilon}_{T+1,T+h} \sim N(\boldsymbol{f}_{T+1,T+h}, \boldsymbol{\Omega}_f), \tag{B.9}$$

where the $k \times 1$ vector $\boldsymbol{f}_{T+1,T+h} \equiv [\overline{\boldsymbol{f}}'_{T+1,T+h}, \underline{\boldsymbol{f}}'_{T+1,T+h}]'$ stacks the means of the distributions under the 'conditional-on-observables forecasting' ($\overline{\boldsymbol{f}}_{T+1,T+h} = \overline{\boldsymbol{C}}\boldsymbol{b}_{T+1,T+h}$) and the 'conditional-on-shocks forecasting' ($\underline{\boldsymbol{f}}_{T+1,T+h} = \underline{\boldsymbol{C}}\boldsymbol{b}_{T+1,T+h} + \boldsymbol{g}_{T+1,T+h}$), and the $k \times k$ matrix $\boldsymbol{\Omega}_f \equiv diag(\overline{\boldsymbol{\Omega}}_f, \underline{\boldsymbol{\Omega}}_f)$.[92]

Based on the combination of 'conditional-on-observables forecasting' and 'conditional-on-shocks forecasting' in Equation (B.9), we can derive the solutions for $\boldsymbol{\mu}_y$ and $\boldsymbol{\Sigma}_y$. Define

$$\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h} \sim N(\boldsymbol{\mu}_\epsilon, \boldsymbol{\Sigma}_\epsilon), \quad \boldsymbol{\Sigma}_\epsilon = \boldsymbol{I} + \boldsymbol{\Psi}_\epsilon, \tag{B.10}$$

so that the $nh \times 1$ vector $\boldsymbol{\mu}_\epsilon$ and the $nh \times nh$ matrix $\boldsymbol{\Psi}_\epsilon$ represent the deviation of the mean and the variance of the structural shocks under the conditional forecast from their values in the unconditional forecast. Given Equations (B.9) and (B.10), we have

$$\begin{aligned} \boldsymbol{f}_{T+1,T+h} &= \boldsymbol{C}\boldsymbol{b}_{T+1,T+h} + \boldsymbol{D}\boldsymbol{\mu}_\epsilon, && \text{(B.11)} \\ \boldsymbol{\Omega}_f &= \boldsymbol{D}(\boldsymbol{I} + \Psi_\epsilon)\boldsymbol{D}'. && \text{(B.12)} \end{aligned}$$

[91] For the conditional forecast that underlies an impulse response function to the i-th shock in period $T+1$ we have

$$\Xi = \boldsymbol{I}_{nh}, \quad \boldsymbol{g}_{T+1,T+h} = [\boldsymbol{e}'_i, \boldsymbol{0}'_{n(h-1)\times 1}]'_{nh\times 1}, \quad \boldsymbol{\Omega}_g = \boldsymbol{0}_{nh\times nh},$$

where $\boldsymbol{e}_i$ is an $n \times 1$ vector of zeros with unity at the i-th position.

[92] Note that $\underline{\boldsymbol{f}}_{T+1,T+h}$ refers to the mean of $\underline{\boldsymbol{C}}\widetilde{\boldsymbol{y}}_{T+1,T+h} = \Xi\boldsymbol{M}'^{-1}\widetilde{\boldsymbol{y}}_{T+1,T+h}$ and hence not just of a path of some observable(s). Instead, $\Xi\boldsymbol{M}'^{-1}\widetilde{\boldsymbol{y}}_{T+1,T+h}$ are the values of the observables that are implied by a specific path of the structural shocks assumed under 'conditional-on-shocks forecasting'.

The solutions for $\boldsymbol{\mu}_\epsilon$ and $\boldsymbol{\Sigma}_\epsilon$ are given by

$$\begin{aligned} \boldsymbol{\mu}_\epsilon &= \boldsymbol{D}^*(\boldsymbol{f}_{T+1,T+h} - \boldsymbol{C}\boldsymbol{b}_{T+1,T+h}), && \text{(B.13)} \\ \boldsymbol{\Sigma}_\epsilon &= \boldsymbol{D}^*\boldsymbol{\Omega}_f\boldsymbol{D}^{*\prime} + (\boldsymbol{I} - \boldsymbol{D}^*\boldsymbol{D}\boldsymbol{D}'\boldsymbol{D}^{*\prime}), && \text{(B.14)} \end{aligned}$$

where the $nh \times k$ matrix $\boldsymbol{D}^*$ is the Moore-Penrose inverse of $\boldsymbol{D}$.[93] Equation (B.13) shows that the path of the implied future structural shocks under the conditional forecast depends on its deviation from the unconditional forecast. In turn, Equation (B.14) shows that the variance of the implied future structural shocks depends on the uncertainty the researcher attaches to the conditional forecast; if the uncertainty is zero, then $\boldsymbol{\Omega}_f = \boldsymbol{0}$ as $\overline{\boldsymbol{\Omega}}_f = \underline{\boldsymbol{\Omega}}_f = \boldsymbol{\Omega}_g = \boldsymbol{0}$, and hence $\boldsymbol{\Sigma}_\epsilon = 0$, meaning that a unique, certain path $\boldsymbol{\mu}_\epsilon$ for the structural shocks is implied by the conditional forecast.[94]

Finally, as

$$\widetilde{\boldsymbol{y}}_{T+1,T+h} = \boldsymbol{b}_{T+1,T+h} + \boldsymbol{M}'\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h}, \tag{B.15}$$

and given Equations (B.13) and (B.14) we have that

$$\begin{aligned} \boldsymbol{\mu}_y &= \boldsymbol{b}_{T+1,T+h} + \boldsymbol{M}'\boldsymbol{D}^*(\boldsymbol{f}_{T+1,T+h} - \boldsymbol{C}\boldsymbol{b}_{T+1,T+h}), && \text{(B.16)} \\ \boldsymbol{\Sigma}_y &= \boldsymbol{M}'\boldsymbol{M} - \boldsymbol{M}'\boldsymbol{D}^*(\boldsymbol{\Omega}_f - \boldsymbol{D}\boldsymbol{D}')\boldsymbol{D}^{*\prime}\boldsymbol{M}. && \text{(B.17)} \end{aligned}$$

Again, when $\boldsymbol{\Omega}_f = \boldsymbol{0}$ then $\boldsymbol{\Sigma}_y = \boldsymbol{0}$, and there is no uncertainty about the path of the observables under the conditional forecast.

It is useful to discuss how the framework of ADPRR is parsed in the context of our paper. Recall that we constrain the effect of a US monetary policy shock on rest-of-the world real activity to be zero, and we assume this occurs due to two offsetting rest-of-the world shocks. Ordering RoW output last in $\boldsymbol{y}_t$, the US monetary policy shock first and the two RoW shocks last in $\boldsymbol{\epsilon}_t$, and denoting by $\boldsymbol{e}_i$ a $n \times 1$ vector of zeros with unity at the i-th position, for 'conditioning-on-observables forecasting' we have

$$\begin{aligned} \overline{\boldsymbol{C}} &= \boldsymbol{I}_h \otimes \boldsymbol{e}_n', && \text{(B.18)} \\ \overline{\boldsymbol{f}}_{T+1,T+h} &= \boldsymbol{0}_{h\times 1}, && \text{(B.19)} \\ \overline{\boldsymbol{\Omega}}_f &= \boldsymbol{0}_{h\times h}. && \text{(B.20)} \end{aligned}$$

The intuition underlying Equations (B.18) and (B.19) is that in the conditional forecast that underlies the impulse response we constrain RoW output (ordered at the n-th position in $\boldsymbol{y}_t$) to be zero over all horizons $T+1, T+2, \ldots, T+h$, and Equation (B.20) indicates that we do not allow for any uncertainty. In turn, for 'conditioning-on-shocks forecasting'

[93] ADPRR discuss the properties of the solutions under different values for k relative to nh.

[94] As discussed in ADPRR, the researcher could impose that the uncertainty under the conditional forecast is identical to that of the unconditional forecast, i.e. set $\boldsymbol{\Omega}_f = \boldsymbol{D}\boldsymbol{D}'$.

we have

$$
\begin{aligned}
\boldsymbol{\Xi} &= \begin{bmatrix} \boldsymbol{e}_1' & \mathbf{0}_{1\times n(h-1)} \\ (\mathbf{0}_{n-3\times 1}, \boldsymbol{I}_{n-3}, \mathbf{0}_{n-3\times 2}) & \mathbf{0}_{n-3\times n(h-1)} \\ \mathbf{0}_{(h-1)(n-2)\times n} & \boldsymbol{I}_{h-1} \otimes (\boldsymbol{I}_{n-2}, \mathbf{0}_{n-2\times 2}) \end{bmatrix}_{h(n-2)\times nh} && \text{(B.21)} \\
\underline{\boldsymbol{f}}_{T+1,T+h} &= \boldsymbol{g}_{T+1,T+h} = [1, \mathbf{0}_{1\times n-3}, \mathbf{0}_{1\times(n-2)(h-1)}]', && \text{(B.22)} \\
\underline{\boldsymbol{\Omega}}_f &= \boldsymbol{\Omega}_g = \mathbf{0}_{h(n-2)\times h(n-2)}. && \text{(B.23)}
\end{aligned}
$$

The first row in Equation (B.21) selects the US monetary policy shock ordered first in $\boldsymbol{\epsilon}_t$ and the first row in Equation (B.22) constrains it to be unity in the impact period $T+1$; the second row in Equation (B.21) selects the non-US monetary policy and the non-RoW shocks ordered from position 2 to $n-3$ in $\boldsymbol{\epsilon}_t$ and the second entry in Equation (B.22) constrains them to be zero in the impact period $T+1$; the third row in Equation (B.21) selects the US monetary policy and the non-RoW shocks and Equation (B.22) constrains them to be zero over horizons $T+2, T+3, \ldots, T+h$. It is furthermore interesting to consider—recalling that $\underline{\boldsymbol{C}} \equiv \boldsymbol{\Xi}\boldsymbol{M}'^{-1}$—the stacked matrices $\boldsymbol{C}$ and $\boldsymbol{D}$ in Equation (B.9)

$$
\boldsymbol{C} = \begin{bmatrix} \overline{\boldsymbol{C}}_{h\times hn} \\ \underline{\boldsymbol{C}}_{h(n-2)\times hn} \end{bmatrix}_{h(n-1)\times hn}, \qquad \boldsymbol{D} = \begin{bmatrix} \overline{\boldsymbol{C}}\boldsymbol{M}' \\ \boldsymbol{\Xi} \end{bmatrix}_{h(n-1)\times nh}. \tag{B.24}
$$

Note that the fact that $\boldsymbol{C}$ and $\boldsymbol{D}$ are not square and full rank reflects that at every horizon we have two RoW shocks to impose one constraint (the absence of a RoW real activity response to a US monetary policy shock), implying a multiplicity of solutions. ADPRR show that the solution chosen in this case—obtained using the Moore-Penrose inverse of $\boldsymbol{D}$—minimises the Frobenius norm of the deviation of the distribution of the structural shocks under the conditional forecast from the baseline, i.e. $\boldsymbol{\mu}_\epsilon$ from $\mathbf{0}$ and $\boldsymbol{\Sigma}_\epsilon$ from $\boldsymbol{I}$. Note that $\boldsymbol{C}$ and $\boldsymbol{D}$ become square and full rank if h additional constraints are imposed. For example, we could impose that the two RoW shocks we use for the offsetting of the effects of the US monetary policy shock on RoW output are of equal size. To do so, we would stack below $\boldsymbol{\Xi}$ in Equation (B.21) an $h \times nh$ matrix

$$
\boldsymbol{\Xi}^{add} = \boldsymbol{I}_h \otimes [0, 0, \ldots, 0, 1, -1]_{1\times n}, \tag{B.25}
$$

and below $\underline{\boldsymbol{f}}_{T+1,T+h}$ in Equation (B.22) an $h \times 1$ vector

$$
\underline{\boldsymbol{f}}^{add}_{T+1,T+h} = \mathbf{0}_{h\times 1}. \tag{B.26}
$$

B.4.1 How plausible is the counterfactual?

When constructing a counterfactual using SSA it might be that the implied shocks are so ‘unusual’ that the analysis becomes subject to the Lucas critique. Against this background, ADPRR propose to use the Kullback-Leibler (KL) divergence $\mathcal{D}(F_{bl}||F_{cf})$ between the distributions of the implied shocks in the conditional forecasts in the baseline F_{bl} and in the counterfactual F_{cf}. While it is straightforward to compute $\mathcal{D}(F_{bl}||F_{cf})$, it is difficult to grasp intuitively whether a given numerical value for the KL divergence is large or small. In other words, the KL divergence can be easily used to rank scenarios, but it is hard to understand how far away they are from the unconditional forecast. To allow an intuitive interpretation of the KL divergence, ADPRR 'calibrate' it based on two generic distributions Q to P, using the KL divergence between two easily interpretable distributions. In particular, ADPRR suggest comparing $\mathcal{D}(F_{bl}||F_{cf})$ with the KL divergence between two binomial distributions Q and P, one with probability q and the other with probability $p = 0.5$. ADPRR suggest calibrating the KL divergence from Q to P to a parameter q that would solve $\mathcal{D}(B(nh; 0.5)||B(nh; q)) = \mathcal{D}(F_{bl}||F_{cf})$. The solution is $q = 0.5(1 + \sqrt{1 - exp(-2z/nh)})$, where $z = \mathcal{D}(F_{bl}||F_{cf})$. Intuitively, the value for the KL divergence $\mathcal{D}(F_{bl}||F_{cf})$ is translated into a comparison between the flip of a fair and a biased coin. For example, a value of $q = 0.501$ suggests that the distribution of the shocks under the counterfactual is hardly different from the distribution under the baseline, so that the counterfactual can be considered as quite realistic relative to the baseline.

It is worthwhile noting that this measure of plausibility is similar in spirit to the concept of 'modest' policy interventions proposed by Leeper and Zha (2003). In particular, the measure proposed by Leeper and Zha (2003) evaluates how unusual a sequence of policy shocks needed to achieve some conditional forecast is. For example, if the counterfactual implies a sequence of policy shocks close to their unconditional mean, the policy intervention is considered 'modest', in the sense that these policy shocks are unlikely to induce agents to revise their beliefs about policy rules. Instead, if the counterfactual involves an unlikely sequence of shocks the analysis is likely to be subject to the Lucas critique. In contrast to the 'modesty' statistic of Leeper and Zha (2003) the q-divergence of ADPRR compares the entire distribution rather than only the path of the shocks and generalises to counterfactuals that use more than a single offsetting shock.

ADPRR propose the KL divergence to assess the plausibility of a conditional forecast relative to an unconditional forecast. In the context of our paper, we need to slightly adjust their proposed KL divergence. In particular, while in the case of ADPRR the baseline is given by an unconditional forecast and the counterfactual by a conditional forecast both afflicted by uncertainty, in our case the baseline and the counterfactual are given by conditional forecasts both of which are *not* subject to any uncertainty. Obviously, the KL divergence is not defined in case the baseline and the counterfactual do not feature any uncertainty. For the purpose of assessing the plausibility of the shocks that materialise to produce our counterfactual, we therefore consider the following exercise. As baseline we consider a conditional forecast in which we assume that a US monetary policy shock of size 1 occurs in period $T + 1$ with certainty, while all other non-US monetary policy shocks in period $T+1$ as well as all shocks in periods $T+2, T+3, \ldots, T+h$ follow their unconditional distributions. For the conditional forecast under the counterfactual, we impose the mean constraint from our main exercise (i.e. that RoW output stays at zero and that there is a US monetary policy shock in period $T+1$) but we also allow for uncertainty.

Formally, this exercise involves setting for the baseline and the counterfactual $\ell \in \{bl, cf\}$ $\overline{\boldsymbol{C}}_\ell = \boldsymbol{I}$ and

$$\begin{aligned} \boldsymbol{f}_{bl} &= \boldsymbol{\mu}_{y,bl} = \boldsymbol{M}'(\boldsymbol{e}_i', \boldsymbol{0}_{n(h-1)\times 1})', && \text{(B.27)} \\ \boldsymbol{f}_{cf} &= \boldsymbol{\mu}_{y,cf}, && \text{(B.28)} \end{aligned}$$

where i is the position of the US monetary policy shock, Equation (B.27) states that the observables on average shall follow the impulse response to a US monetary policy shock in period $T+1$ under the baseline, and Equation (B.27) that they shall follow the path we obtained in the SSA counterfactual. Moreover, we set $\boldsymbol{\Xi}_\ell = \boldsymbol{0}$ and $\underline{\boldsymbol{C}}_\ell = \boldsymbol{0}$ so that $\boldsymbol{D}_\ell = \boldsymbol{M}'$, $\boldsymbol{\Psi}_{\epsilon,\ell} = \boldsymbol{0}$ as we allow the shocks to have their unconditional variance. Hence we have

$$\begin{aligned} \boldsymbol{\Omega}_{f,\ell} &= \boldsymbol{D}_\ell \boldsymbol{D}_\ell' = \boldsymbol{M}'\boldsymbol{M}, && \text{(B.29)} \\ \boldsymbol{\Sigma}_{\epsilon,\ell} &= \boldsymbol{D}_\ell^* \boldsymbol{\Omega}_{f,\ell} \boldsymbol{D}_\ell^{*\prime} = \boldsymbol{D}_\ell^{-1} \boldsymbol{\Omega}_{f,\ell} \boldsymbol{D}_\ell^{-1\prime} = \boldsymbol{M}'^{-1}\boldsymbol{M}'\boldsymbol{M}\boldsymbol{M}^{-1} = \boldsymbol{I}, && \text{(B.30)} \\ \boldsymbol{\mu}_{\epsilon,\ell} &= \boldsymbol{D}_\ell^* \boldsymbol{f}_\ell = \boldsymbol{D}_\ell^{-1} \boldsymbol{f}_\ell = \boldsymbol{M}'^{-1} \boldsymbol{f}_\ell, && \text{(B.31)} \\ \boldsymbol{\Sigma}_{y,\ell} &= \boldsymbol{M}'\boldsymbol{M} - \boldsymbol{M}'\boldsymbol{D}_\ell^*(\boldsymbol{\Omega}_{f,\ell} - \boldsymbol{D}_\ell \boldsymbol{D}_\ell')\boldsymbol{D}_\ell^{*\prime}\boldsymbol{M} \\ &= \boldsymbol{M}'\boldsymbol{M} - \boldsymbol{M}'\boldsymbol{D}_\ell^{-1}(\boldsymbol{\Omega}_{f,\ell} - \boldsymbol{D}_\ell \boldsymbol{D}_\ell')\boldsymbol{D}_\ell^{-1\prime}\boldsymbol{M} \\ &= \boldsymbol{M}'\boldsymbol{M} - \boldsymbol{M}'\boldsymbol{M}'^{-1}(\boldsymbol{\Omega}_{f,\ell} - \boldsymbol{M}'\boldsymbol{M})\boldsymbol{M}^{-1}\boldsymbol{M} \\ &= \boldsymbol{M}'\boldsymbol{M} - (\boldsymbol{\Omega}_{f,\ell} - \boldsymbol{M}'\boldsymbol{M}) \\ &= \boldsymbol{M}'\boldsymbol{M}. && \text{(B.32)} \end{aligned}$$

Note that $\boldsymbol{\mu}_{\epsilon,bl}$ equals a vector of zeros with unity at the i^{th} position, where i is the position of the US monetary policy shock. The KL divergence between the distribution of the shocks under the baseline $\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h,bl}$ and the counterfactual $\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h,cf}$ is then given by

$$\mathcal{D}(F_{bl}||F_{cf}) = \frac{1}{2}\left[tr\left(\boldsymbol{\Sigma}_{\epsilon,cf}^{-1}\boldsymbol{\Sigma}_{\epsilon,bl}\right) + \left(\boldsymbol{\mu}_{\epsilon,cf} - \boldsymbol{\mu}_{\epsilon,bl}\right)' \boldsymbol{\Sigma}_{\epsilon,cf}^{-1} \left(\boldsymbol{\mu}_{\epsilon,cf} - \boldsymbol{\mu}_{\epsilon,bl}\right) - nh + log\left(\frac{|\boldsymbol{\Sigma}_{\epsilon,cf}|}{|\boldsymbol{\Sigma}_{\epsilon,bl}|}\right)\right]. \quad \text{(B.33)}$$

B.5 Implementation of the MRE approach

The posterior distribution of the impulse responses $f(\cdot)$ is approximated by N draws obtained from a Bayesian estimation algorithm. Following the importance sampling procedure of Arias et al. (2018,0), the re-sampled draws from the BPSVAR for $\boldsymbol{y}_{T+1,T+h}$ constitute an unweighted and independent sample from the posterior distribution $f(\cdot)$, and as such are assigned a weight of $w_i = 1/N$, $i = 1, 2, \dots, N$. The counterfactual posterior distribution $f^\star(\cdot)$ can be approximated by assigning different weights w_i^* to the draws from the baseline posterior.

The relative entropy (or distance) between the approximated posterior distributions is measured by

$$\mathcal{D}(f^*, f) = \sum_{i=1}^{N} w_i^\star log\left(\frac{w_i^\star}{w_i}\right). \tag{B.34}$$

The goal of the MRE approach is to determine the counterfactual weights $\boldsymbol{w}^*$ that minimise $\mathcal{D}(\cdot)$ subject to

$$w_i^\star \geq 0, \qquad \forall\, i = 1, 2, ..., N, \tag{B.35}$$

$$\sum_{i=1}^{N} w_i^* = 1, \tag{B.36}$$

$$\sum_{i=1}^{N} w_i^* g(\boldsymbol{y}_{T+1,T+h}^{(i)}) = \bar{\boldsymbol{g}}, \tag{B.37}$$

where $\boldsymbol{y}_{T+1,T+h}^{(i)}$ are the impulse responses to a US monetary policy shock as defined in Section 3 in the paper. Equations (B.35) and (B.36) reflect that the weights are probabilities, and Equation (B.37) that the counterfactual posterior distribution shall satisfy some constraint.

In particular, in our application for Equation (B.37) we have

$$\sum_{i=1}^{N} y_{ip^*,T+h}^{(i)} w_{i,h}^* = 0, \tag{B.38}$$

where $y_{ip^*,T+h}^{(i)}$ denotes the impulse response of RoW real activity to a US monetary policy shock at horizon h associated with the i-th draw. Notice that—consistent with the baseline posterior for which we report point-wise means in Figure 1 and elsewhere in the paper as well as in line with Giacomini and Ragusa (2014)—we apply the MRE approach separately at each impulse response horizon $T+1, T+2, \dots, T+h$.

As shown by Robertson et al. (2005) and Giacomini and Ragusa (2014), the weights of the counterfactual posterior distribution $\boldsymbol{w}_h^*$ can be obtained numerically by tilting the weights of the baseline posterior distribution $\boldsymbol{w}_h$ using the method of Lagrange. In particular, the weights of the counterfactual posterior distribution are given by

$$w_{i,h}^* = \frac{w_{i,h} \exp\left[\lambda_h g(y_{ip^*,T+h}^{(i)})\right]}{\sum_{i=1}^{N} w_{i,h} \exp\left[\lambda_h g(y_{ip^*,T+h}^{(i)})\right]}, \quad i = 1, 2, \dots, N, \tag{B.39}$$

where λ_h is the Lagrange multiplier associated with the constraint $g(y_{ip^*,T+h}^{(i)}) = y_{ip^*,T+h}^{(i)} = 0$. It can be shown that the Lagrange multiplier can be obtained numerically as

$$\lambda_h = \arg\min_{\tilde{\lambda}_h} \sum_{i=1}^{N} w_{i,h} \exp\left\{\tilde{\lambda}_h \left[g(y_{ip^*,T+h}^{(i)})\right]\right\}. \tag{B.40}$$

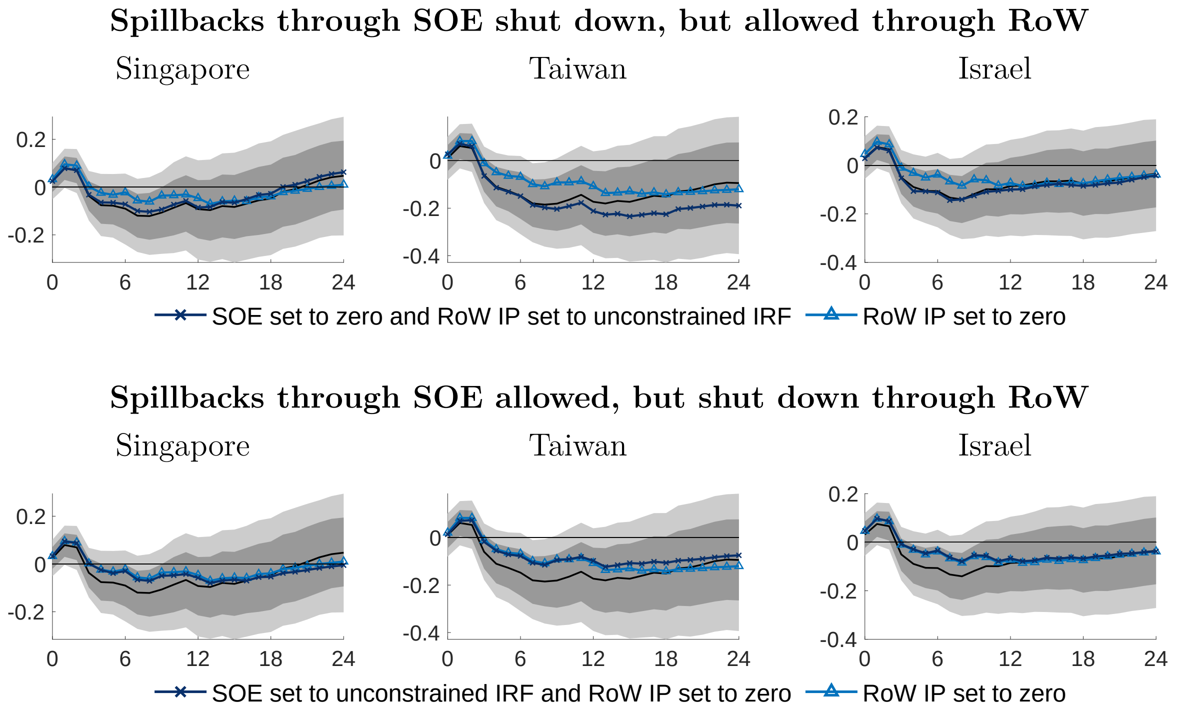
B.6 Placebo tests

The top row in Figure B.12 presents results for estimations in which we constrain real activity spillovers from US monetary policy to some individual SOEs to be nil while those to the RoW to be identical to our baseline (dark blue crossed lines). The impulse responses of US industrial production from this alternative counterfactual are very different from the counterfactual in which real activity spillovers to the entire RoW are constrained to be nil (light blue lines with triangles). In fact, the impulse responses of US industrial production from this alternative counterfactual in which only spillovers to individual SOEs are precluded are very similar to the unconstrained baseline (black solid lines). SSA and MRE thus indicate that the spillbacks from US monetary policy that materialise through individual SOEs alone are essentially zero. This is plausible.

A related check for the plausibility of SSA and MRE counterfactuals is to explore how large spillbacks from US monetary policy are assessed if we constrain spillovers to the RoW to be nil but leave those to individual SOEs unconstrained. The impulse responses of US industrial production for this specification (dark blue crossed lines) in the bottom row in Figure B.12 are hardly distinguishable from the counterfactual in which spillovers to the entire RoW are precluded (light blue lines with triangles. This suggests that the contribution of individual SOEs to the overall spillbacks from US monetary policy are negligible. Again, this is plausible.

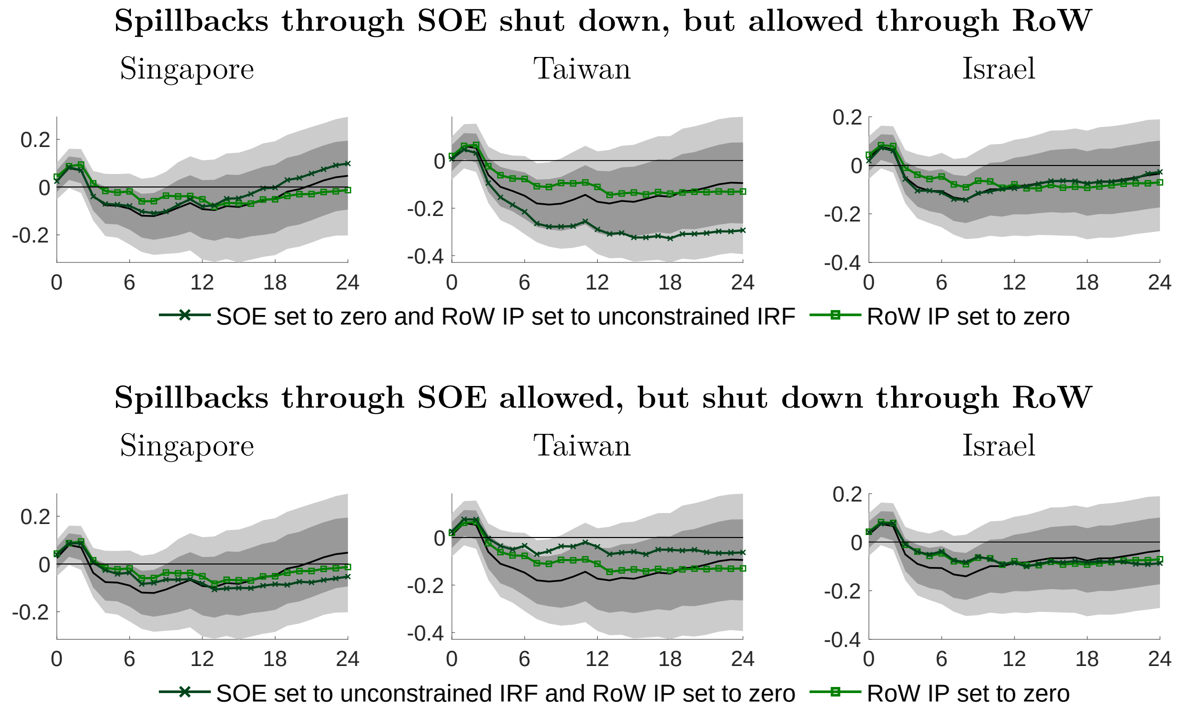
Overall, the results from these exercises bolster the plausibility of our counterfactual analysis based on SSA and MRE. Figures B.13 and B.14 provide results for SSA with RoW shocks and for MRE.

Figure B.12: SSA with all shocks placebo-test impulse responses of US industrial production



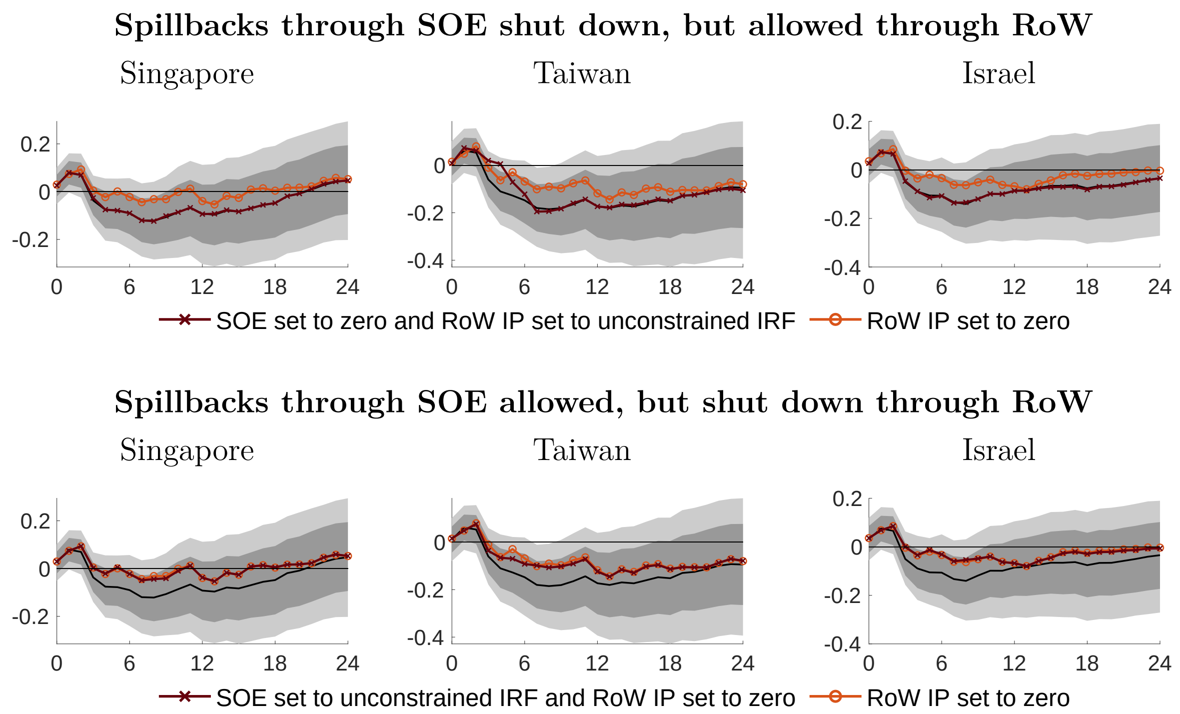
Notes: The black solid lines depict the response of US industrial production from VAR models in which SOE industrial production is added to the vector of observables and their impulse responses to a US monetary policy shock are not constrained, and the grey shaded areas represent the associated 68% centered point-wise probability bands. The light blue lines with triangles depict the counterfactual impulse response in which spillovers to RoW and SOE industrial production are precluded. In the top row, the dark blue lines with crosses depict the counterfactual response of US industrial production when SOE industrial production is constrained to not respond to a US monetary policy shock, and RoW industrial production is constrained to respond as in the unconstrained case. In the bottom row, the dark blue lines with crosses depict the response of US industrial production when the response of SOE industrial production is constrained to respond as in the unconstrained case, and RoW industrial production is constrained to not respond to a US monetary policy shock. Figure B.13 and Figure B.14 document the placebo tests based on SSA with RoW shocks and MRE.

Figure B.13: SSA with RoW shocks placebo-test responses of US industrial production



Notes: See the notes to Figure B.12.

Figure B.14: MRE placebo-test responses of US industrial production



Notes: See the notes to Figure B.12.

Appendix C

Appendix for Chapter 3

C.1 Advantages of the BPSVAR framework over the traditional frequenstist external instruments SVAR framework

The BPSVAR framework has several appealing features relative to traditional frequentist external instrument SVAR models that render it particularly well-suited for the purpose of estimating the effects of global risk and US monetary policy shocks on the world economy. First, it requires relatively weak additional identifying assumptions when more than one structural shock is to be identified by proxy variables. In this case, the shocks are only set identified as rotations of the structural shocks $\boldsymbol{Q}\boldsymbol{\epsilon}_t^*$ with orthonormal matrices $\boldsymbol{Q}$ also satisfy the relevance and exogeneity conditions in Equation (2) in the manuscript. Therefore, additional restrictions are needed in order to point-identify the structural shocks in $\boldsymbol{\epsilon}_t^*$. In the frequentist external instruments VAR model these additional restrictions are imposed on the contemporaneous relationships between the *endogenous variables* $\boldsymbol{y}_t$ reflected in $\boldsymbol{A}_0^{-1}$ (Mertens and Ravn 2013; Lakdawala 2019). However, Arias et al. (2021) show that relaxing this type of additional identifying assumptions can change the results profoundly. Instead, the BPSVAR framework allows us to impose the additional identifying assumptions on the contemporaneous relationships between the *structural shocks* $\boldsymbol{\epsilon}_t^*$ *and proxy variables* $\boldsymbol{m}_t$ reflected in $\boldsymbol{V}$ in the relevance condition in Equation (2) in the manuscript. For example, we can impose the restriction that a particular structural shock does not affect a particular proxy variable. Restrictions on the contemporaneous relationships are arguably weaker for structural shocks and proxy variables in $\boldsymbol{V}$ than for the endogenous variables in $\boldsymbol{A}_0^{-1}$.

Second, the BPSVAR framework allows coherent and exact finite sample inference, even in settings in which the proxy variables are weak instruments and only set rather than point identification is achieved with a combination of sign, magnitude and zero restrictions (see Moon and Schorfheide 2012; Caldara and Herbst 2019; Arias et al. 2021). In particular, frequentist external instruments VAR models are estimated in a two-step procedure (Mertens and Ravn 2013; Gertler and Karadi 2015): (i) estimate the reduced-form VAR model; (ii) regress the reduced-form residuals on the proxy variable to obtain the structural parameters. This two-step procedure is inefficient, as the estimation of the reduced-form VAR model in (i) is not informed by the proxy variable. In contrast, the BPSVAR model considers the joint likelihood of the endogenous variables and the proxy variables, so that the proxy variables inform the estimation of both reduced-form and structural parameters. The BPSVAR framework also facilitates inference, as the joint estimation captures all sources of uncertainty. Furthermore, as long as the prior distribution is proper, in a Bayesian setting inference is straightforward even when the instruments are weak (Poirier 1998). By contrast, frequentist external instruments VAR models require an explicit theory to accommodate weak instruments (Montiel Olea et al. 2021), either to derive the asymptotic distributions of the estimators or to ensure satisfactory coverage in bootstrap algorithms.[95]

Third, from from the BPSVAR model augemnted with equatins for the proxy variables it can be seen that framework is relatively flexible in that it allows for the proxy variables to be serially correlated and to be affected by lags of the endogenous variables as well as by measurement error. This is a useful feature as it has been shown that some widely-used proxy variables are serially correlated and/or contaminated by measurement error (Miranda-Agrippino and Ricco 2021). In these cases, it is typically proposed to cleanse the proxy variables in an additional step preceding the analysis in the VAR model, exacerbating issues regarding efficiency and coherent inference.

And fourth, the BPSVAR model allows us to incorporate a prior belief about the strength of the proxy variables as instruments based on the notion that "researchers construct proxies to be relevant" (Caldara and Herbst 2019, p. 165). In particular, consider the 'reliability matrix' $\boldsymbol{R}$ derived in Mertens and Ravn (2013) given by

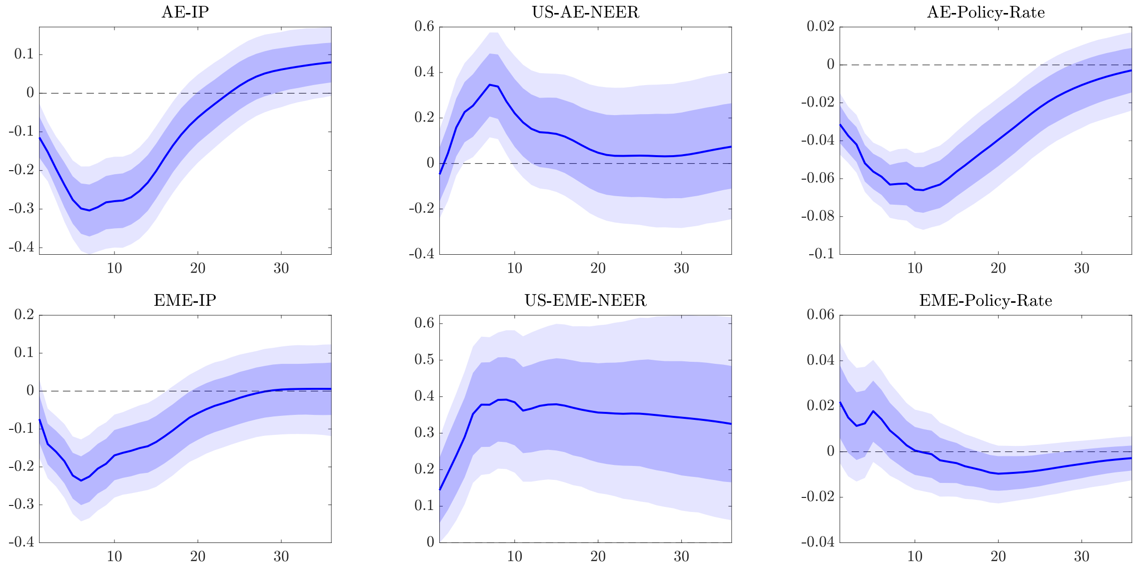
$$\boldsymbol{R} = \left(\boldsymbol{\Gamma}_{0,2}^{-1'}\boldsymbol{\Gamma}_{0,2} + \boldsymbol{V}\boldsymbol{V}'\right)^{-1}\boldsymbol{V}\boldsymbol{V}'. \tag{C.1}$$

Intuitively, $\boldsymbol{R}$ indicates the share of the total variance of the proxy variables that is accounted for by the structural shocks $\boldsymbol{\epsilon}_t^*$. Specifically, the minimum eigenvalues of $\boldsymbol{R}$ can be interpreted as the share of the variance of (any linear combination of) the proxy variables explained by the structural shocks $\boldsymbol{\epsilon}_t^*$ (Gleser 1992).

C.2 Additional figures

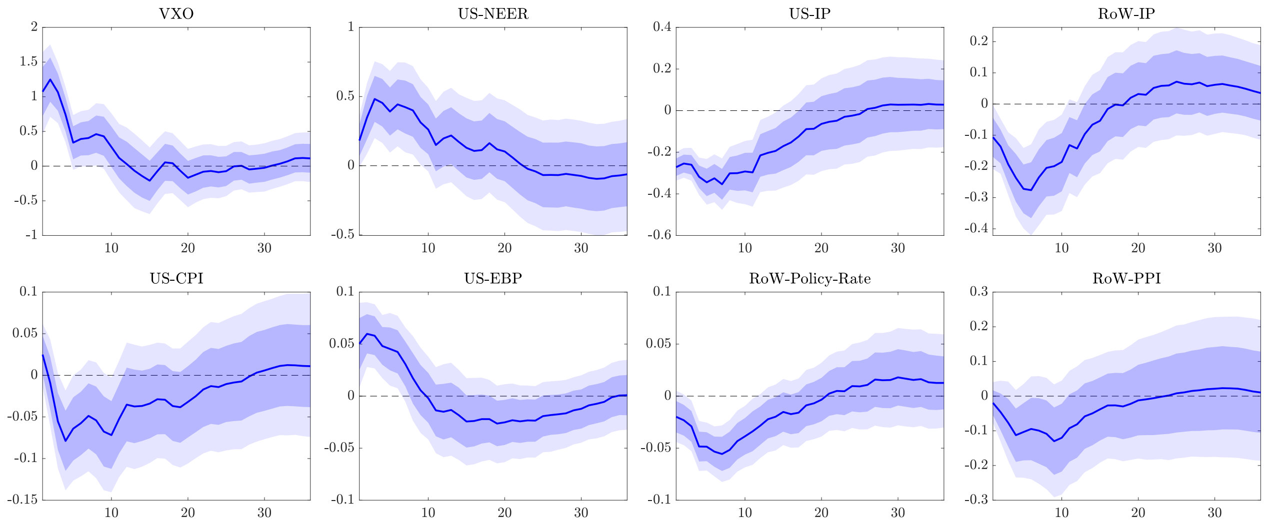
[95]To the best of our knowledge, there is no consensus yet on how to conduct inference in frequentist external instruments VAR models, even in a setting with only a single proxy variable (Jentsch and Lunsford 2019).

Figure C.1: Impulse responses for AEs and EMEs to a global risk shock



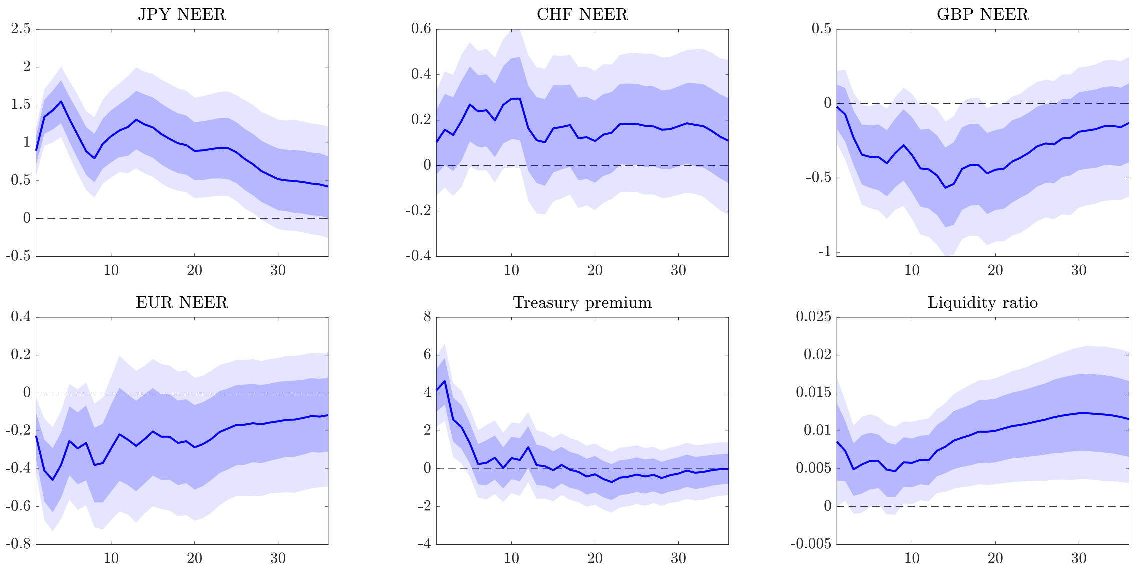
Note: The figure presents the impulse responses to a one-standard deviation global risk shock. Due to the larger dimensionality of the VAR model we use informative Minnesota-type priors and optimal hyperpriors/prior tightness as suggested by Giannone et al. (2015) in the estimation. Horizontal axis measures time in months, vertical axis deviation from pre-shock level. Blue solid line represents point-wise posterior mean and shaded areas 68%/90% equal-tailed, point-wise credible sets.

Figure C.2: Impulse responses to a global risk shock when including RoW PPI



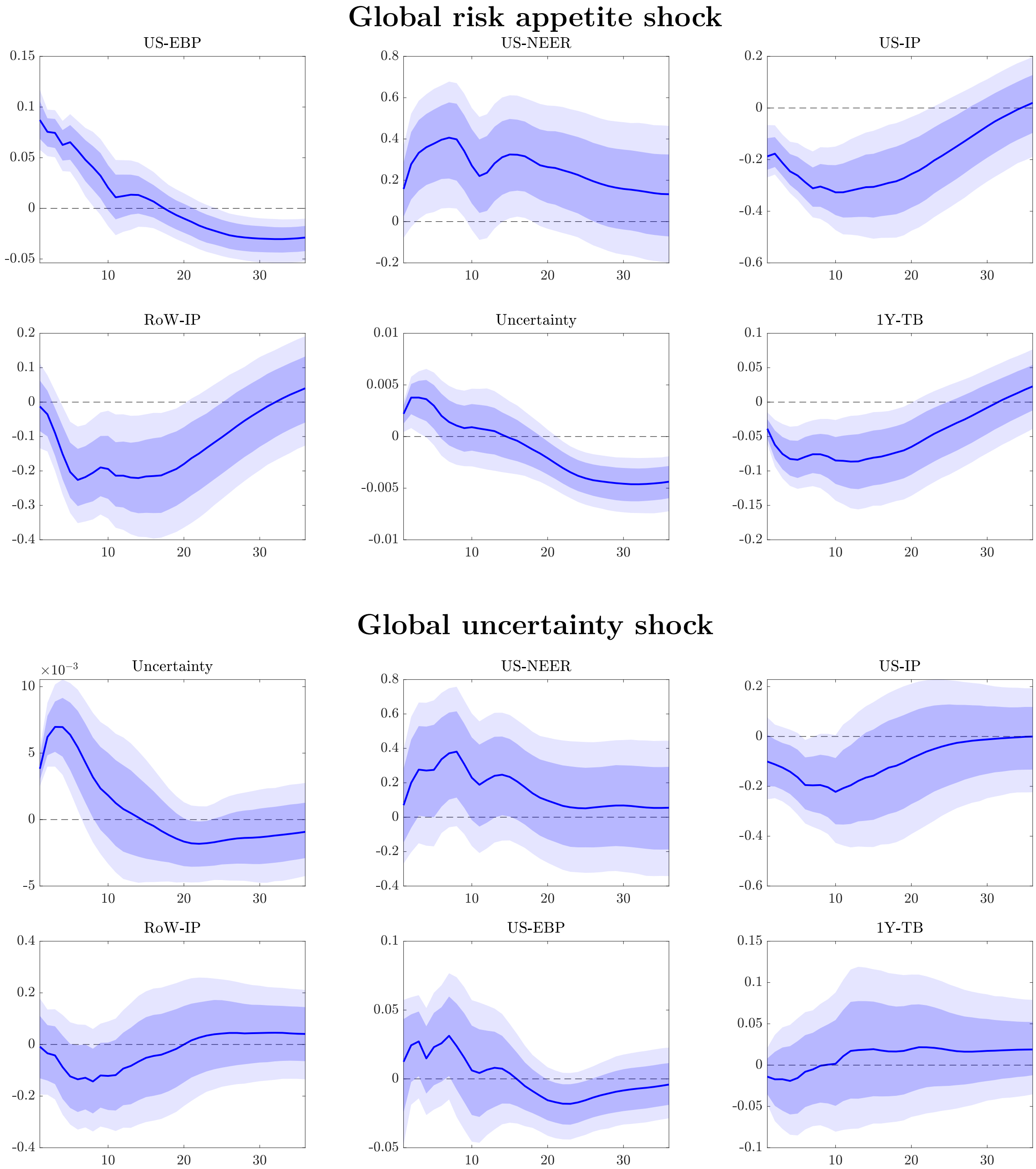
Note: Horizontal axis measures time in months, vertical axis deviation from pre-shock level; size of shock is one standard deviation; blue solid line represents point-wise posterior mean and shaded areas 68%/90% equal-tailed, point-wise credible sets.

Figure C.3: Impulse responses of additional variables



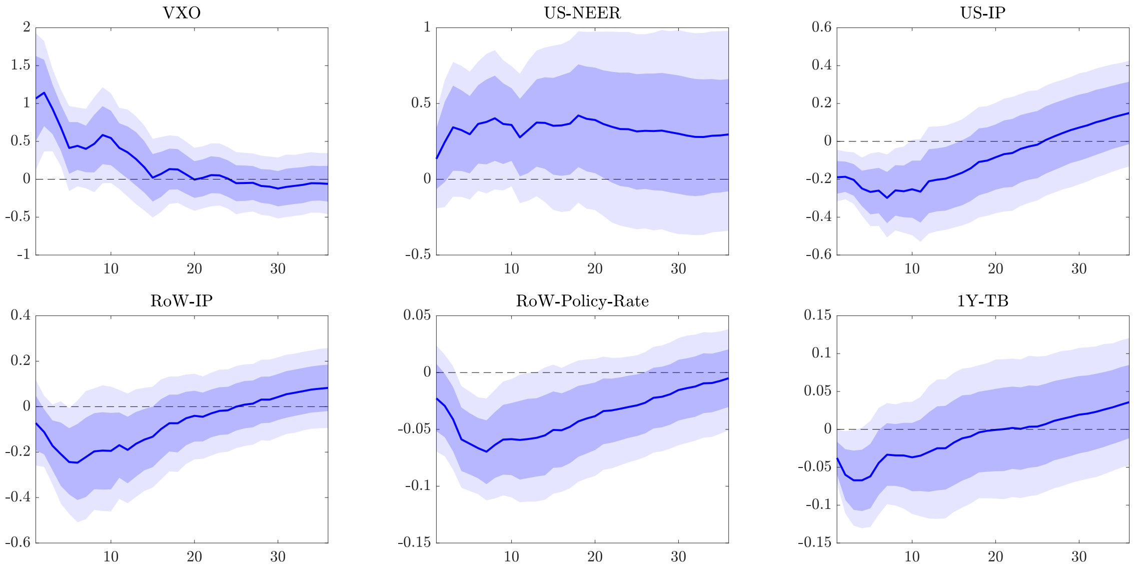
Note: See notes to Figure C.2. Responses are obtained from estimating the baseline BPSVAR model with the vector $\boldsymbol{y}_t$ augmented with one additional variable at a time. Because data on the liquidity ratio is only available from 2001 we use informative priors and optimal hyperpriors/prior tightness as suggested by Giannone et al. (2015).

Figure C.4: Impulse responses to global risk appetite and global uncertainty shocks identified with FEVD restrictions



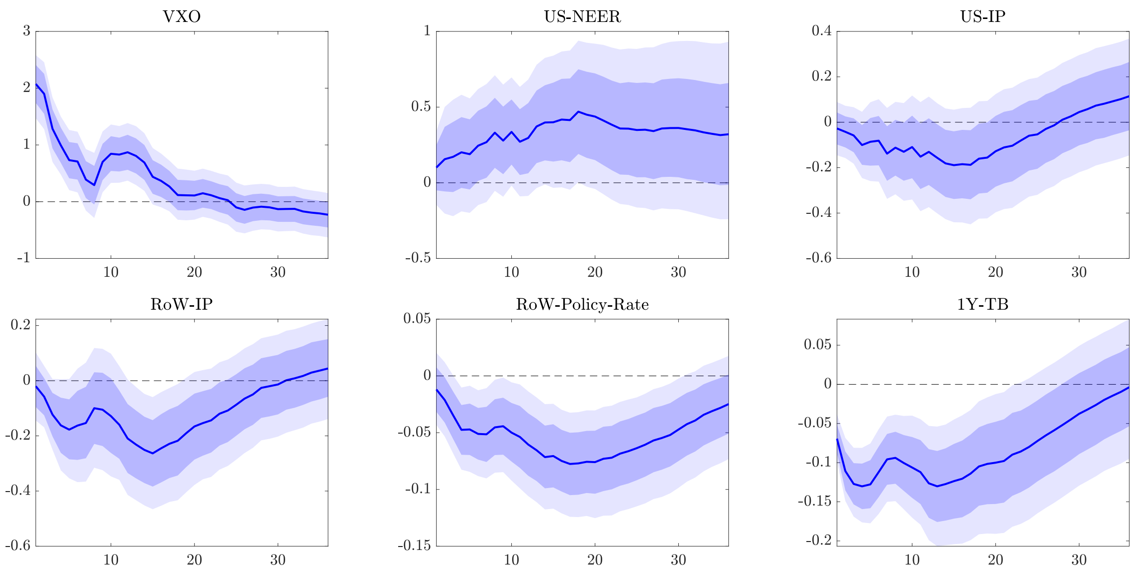
Note: The figure presents the impulse responses to a one-standard deviation global appetite risk and global uncertainty shocks based on an alternative identification scheme in which we (i) allow both shocks to drive the gold price surprises (ii) impose that the global risk appetite (uncertainty) shock explains the largest share of the FEVD of the excess bond premium (macroeconomic uncertainty measure of Jurado et al. 2015). We drop the VXO from the BPSVAR model as it reflects both risk aversion and uncertainty and replace it by the macroeconomic uncertainty measure of Jurado et al. (2015). Impulse responses of RoW Policy Rate and the US CPI are omitted to save space. Horizontal axis measures time in months, vertical axis deviation from pre-shock level. Blue solid line represents point-wise posterior mean and shaded areas 68%/90% equal-tailed, point-wise credible sets.

Figure C.5: Impulse responses to a global risk shock when allowing the gold price surprises to be correlated with all structural shocks



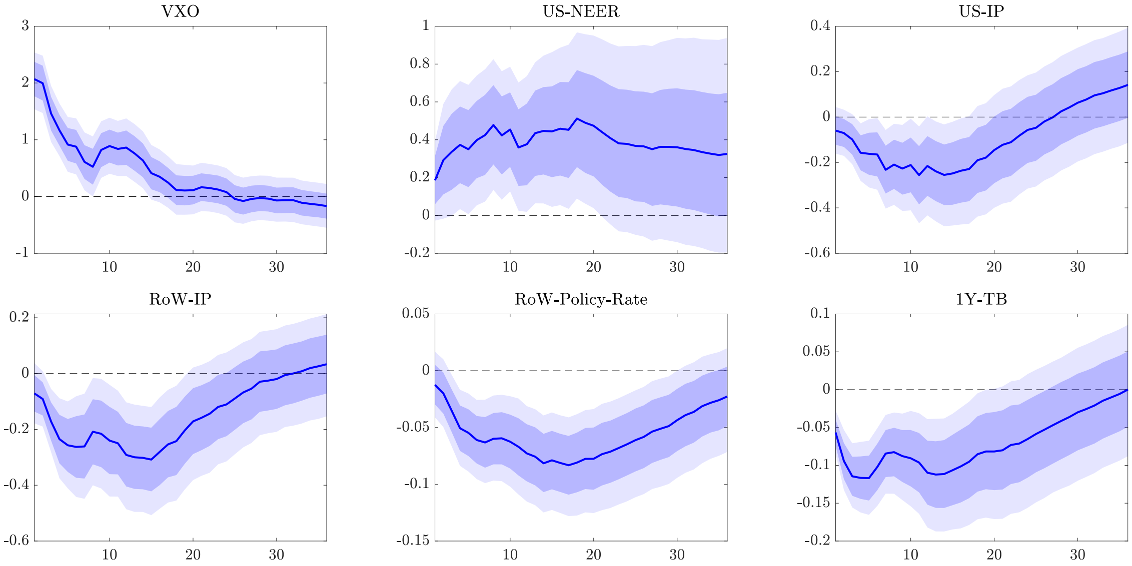
Note: The figure presents the impulse responses to a one-standard deviation global risk shock based on an alternative identification scheme in which the gold price surprises are allowed to be correlated with all structural shocks, imposing only that the correlation is strongest with the global risk shock. Impulse responses of US CPI and the EBP are omitted to save space.

Figure C.6: Impulse responses to a global risk shock when considering intra-daily surprises in 30-year Treasury yields instead of the gold price as proxy variable



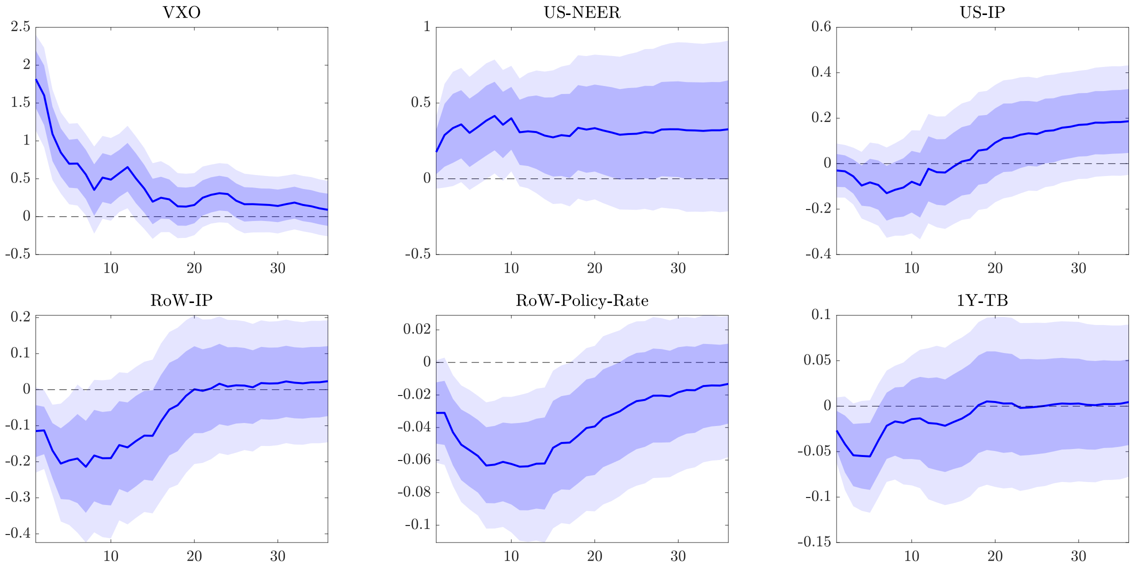
Note: The results are obtained from a BPSVAR model with intra-daily 30-year Treasury yield surprises as proxy variable. We drop the identification of the monetary policy shocks for this specification because we don't compute any counterfactuals using this specification. Impulse responses of US CPI and the EBP are omitted to save space.

Figure C.7: Impulse responses to a global risk shock when considering intra-daily surprises in the US dollar-euro exchange rate instead of the gold price as proxy variable



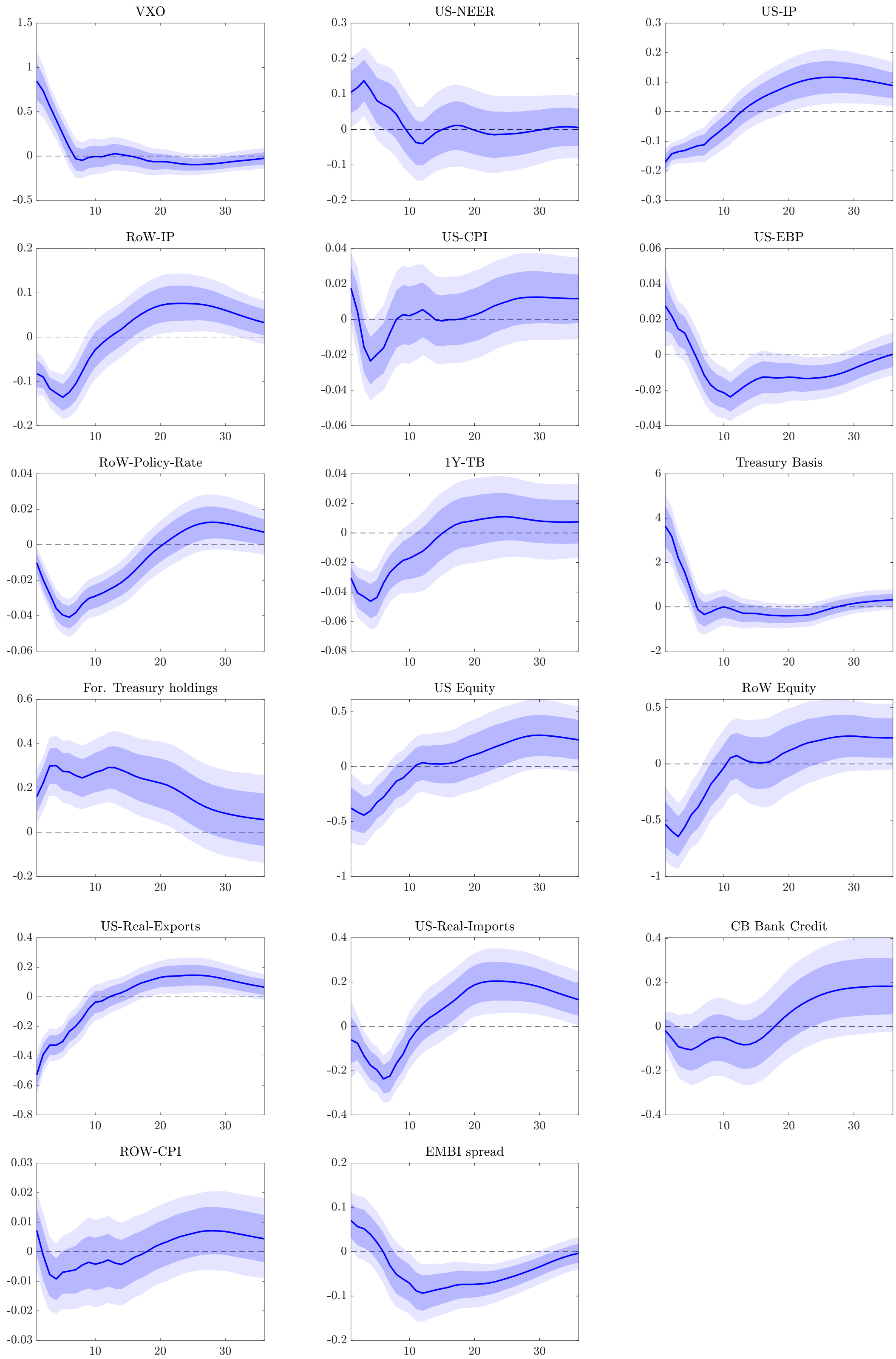
Note: The results are obtained from a BPSVAR model with intra-daily US dollar-euro exchange rate surprises as proxy variable. We drop the identification of the monetary policy shocks for this specification because we don't compute any counterfactuals using this specification. Impulse responses of US CPI and the EBP are omitted to save space.

Figure C.8: Impulse responses to a global risk shock when considering changes in the Geopolitical Risk Index of Caldara and Iacoviello (2022) instead of gold price surprises as proxy variable



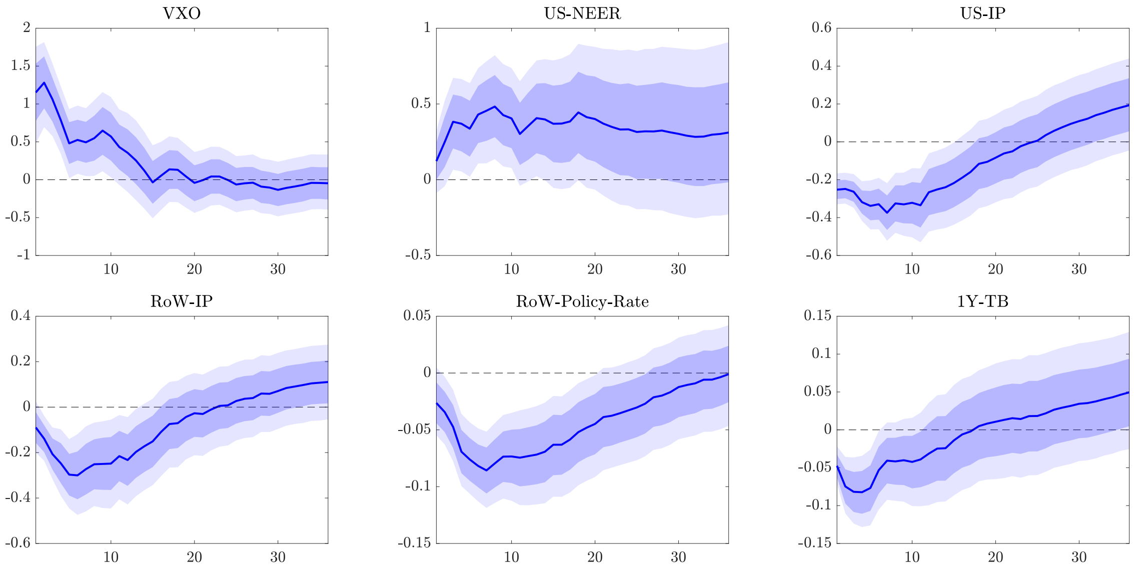
Note: The results are obtained from a BPSVAR model with monthly changes in the Geopolitical Risk index of Caldara and Iacoviello (2022) as proxy variable. We drop the identification of the monetary policy shocks for this specification because we don't compute any counterfactuals using this specification. Impulse responses of US CPI and the EBP are omitted to save space.

Figure C.9: Impulse responses to a global risk shock from a large BPSVAR model



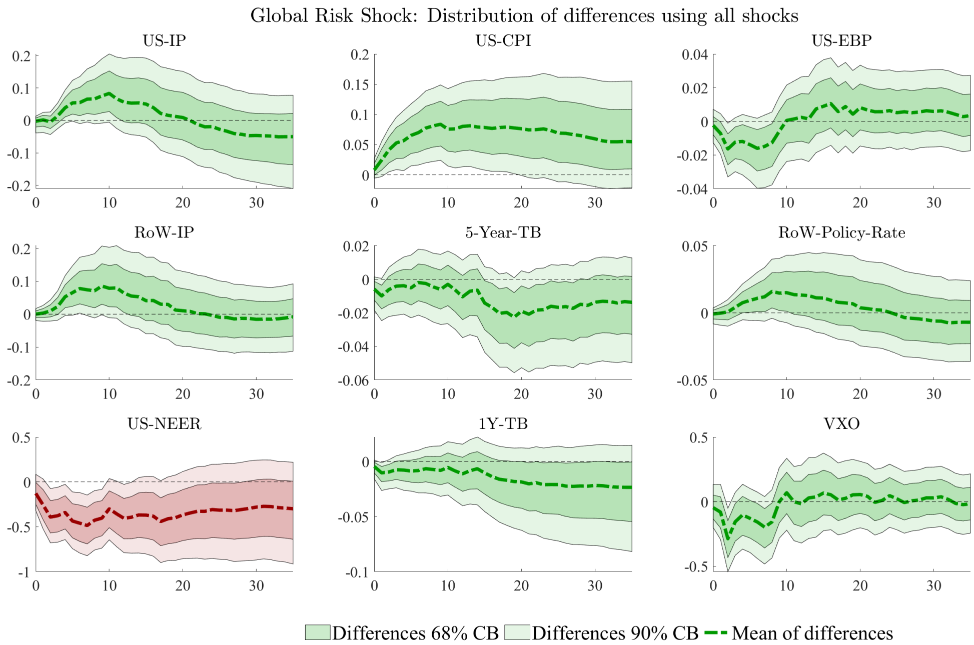
Note: The model is estimated with informative Minnesota-type priors and optimal hyperpriors/prior tightness as in Giannone et al. (2015). We do not include the liquidity ratio in the VAR model because it is only available for a substantially shorter sample period (see Table C.1).

Figure C.10: Impulse responses to global risk shock when no relevance threshold is imposed



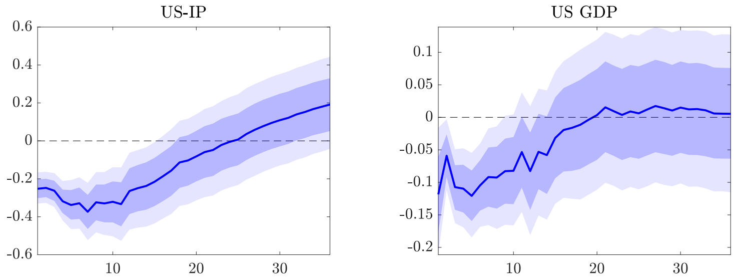
Note: The figure presents the impulse responses to a one-standard deviation global risk shock based on an alternative identification scheme in which we do not impose any relevance threshold. Impulse responses of US CPI and the EBP are omitted to save space. Horizontal axis measures time in months, vertical axis deviation from pre-shock level. Blue solid line represents point-wise posterior mean and shaded areas 68%/90% equal-tailed, point-wise credible sets.

Figure C.11: Distribution of differences for the SSA



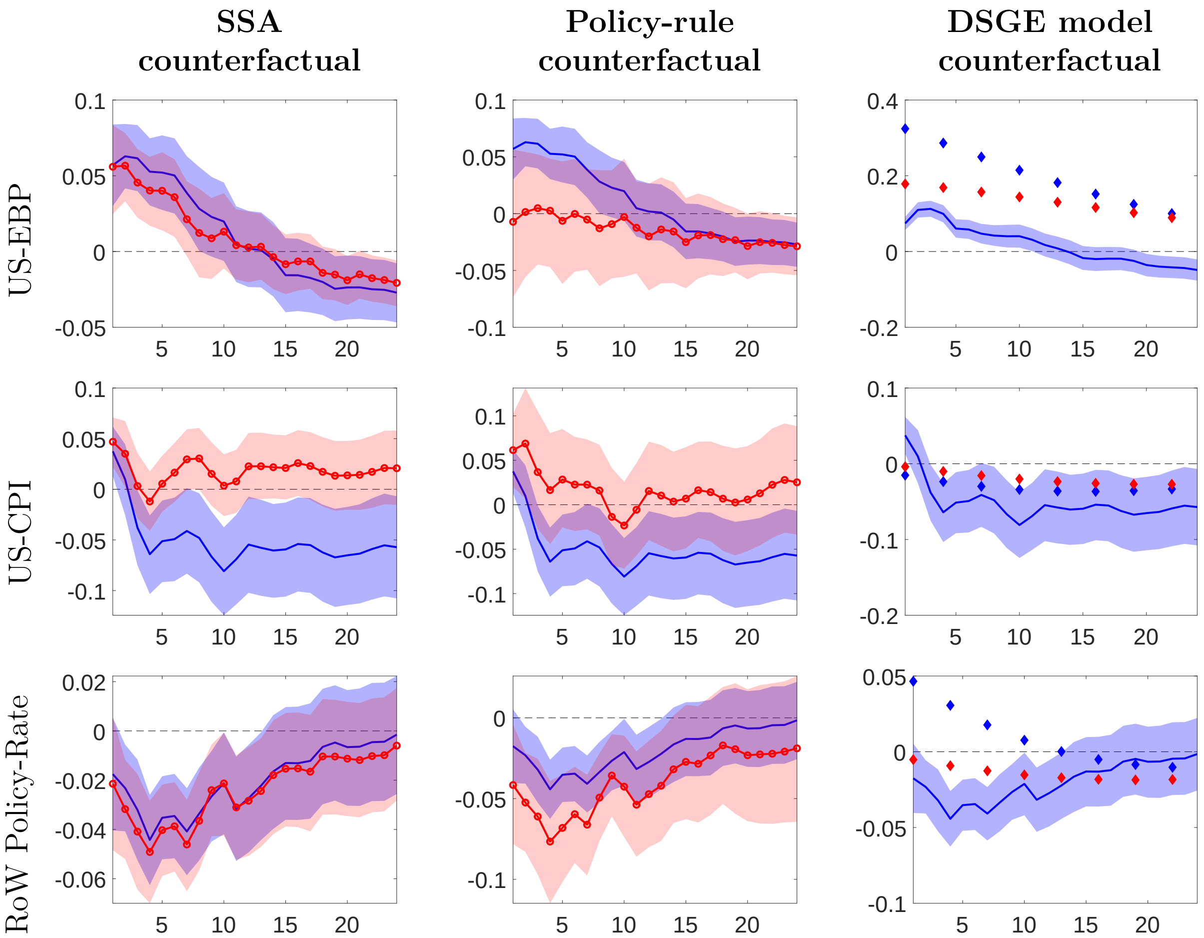
Note: The figures shows the posterior distribution of the differences between the baseline impulse response from the BPSVAR and the counterfactual impulse responses from the SSA. Dark (light) green bands represent 68% (90%) point-wise credible sets. The figure shows that for most of the variables, roughly 90% of the posterior probability mass lies below or above the zero line.

Figure C.12: Impulse responses of US IP and monthly US GDP



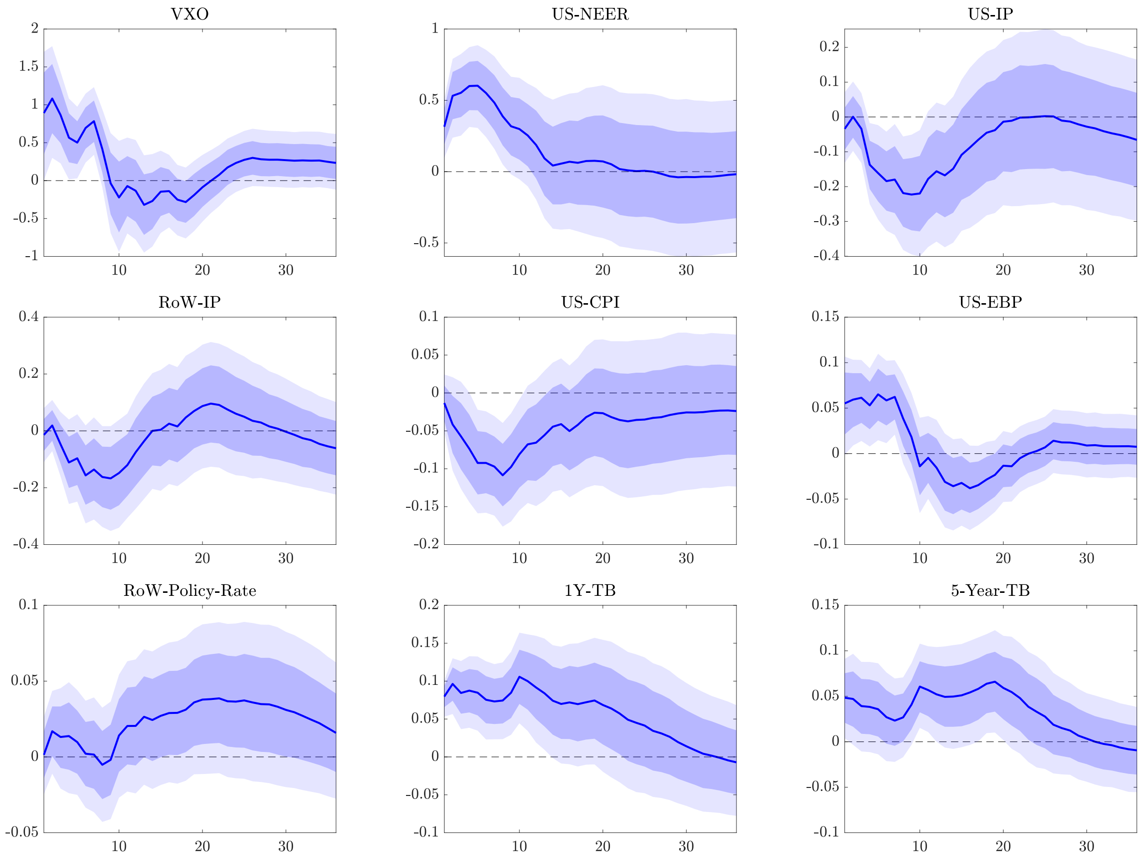
Note: The left-hand side panel depicts the of US IP from the baseline BPSVAR, whereas the right-hand side panel depicts the response of a monthly estimate of US GDP from Standard & Poors. The IRFs confirm that the response of US IP is roughly 2.5 times larger than those of US GDP, which we assume when comparing the DSGE model to the BPSVAR. Horizontal axis measures time in months, vertical axis deviation from pre-shock level. Blue solid line represents point-wise posterior mean and shaded areas 68%/90% equal-tailed, point-wise credible sets.

Figure C.13: Baseline and counterfactual responses of remaining BPSVAR model variables



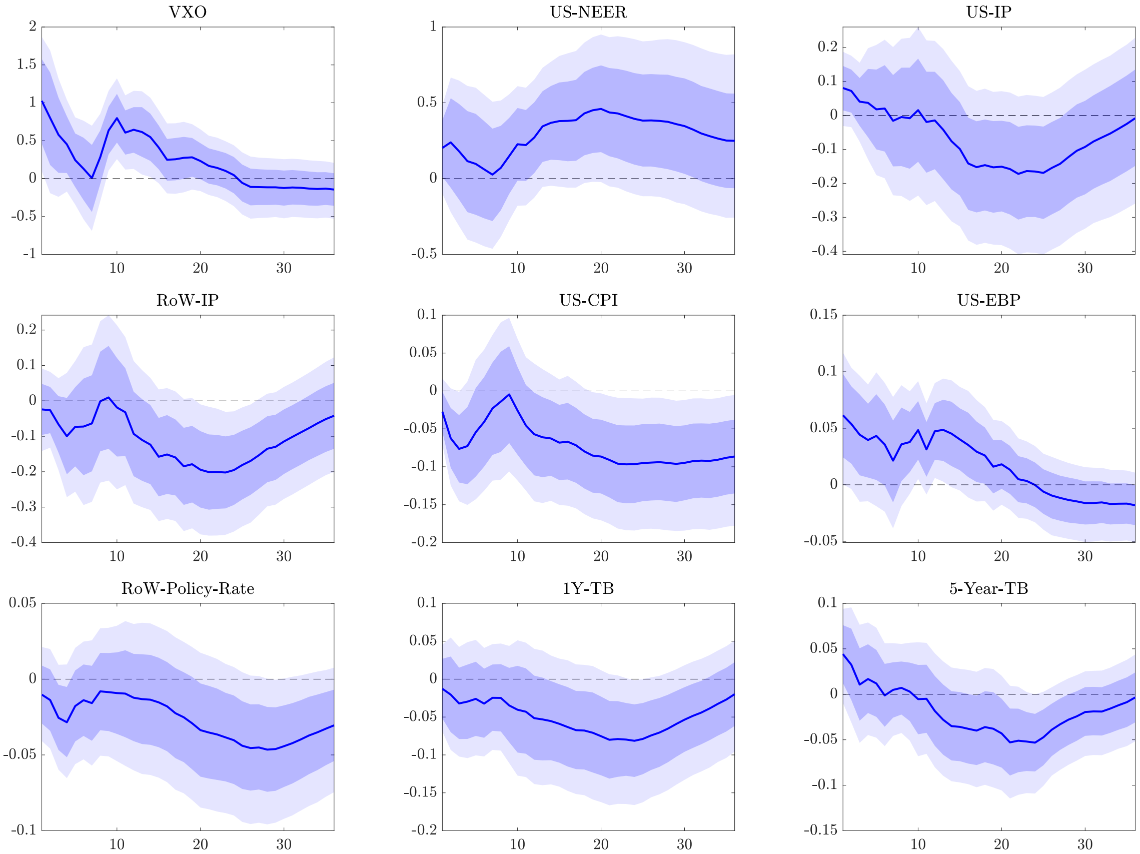
Note: See notes to 4. As the model of Georgiadis et al. (2023) does not include an exact counterpart US EBP we plot the responses of the US credit spread instead. While in the baseline specification of the BPSVAR we included the AE policy rate as our indicator for the RoW policy stance as weighted average of the policy rates of the "entire" RoW as computed in the Dallas Fed Global Economic Indicators data (Grossman et al. 2014) are extremely volatile volatility in the 1990s due to several crises involving major emerging market economies (EMEs), including Mexico, Brazil, Russia, Thailand, Indonesia, Malaysia, South Korea, Philippines, Argentina and Turkey. As in the DSGE model we want to capture the policy stance in the "entire" RoW the blue lines in the last row corresponds to results from a specification with a 'hybrid' RoW policy rate. : Due to the extreme values in the beginning of the sample we impute backwards from 2000 the levels of RoW policy rates in the AE policy rate. Given that the size of EMEs—especially of China—took off only after the late 1990s, this should introduce only mild distortions in the RoW aggregate series. Due to the extreme values in the beginning of the sample we impute backwards from 2000 the levels of RoW policy rates changes in the AE policy rates.

Figure C.14: Responses to a contractionary conventional US monetary policy shock



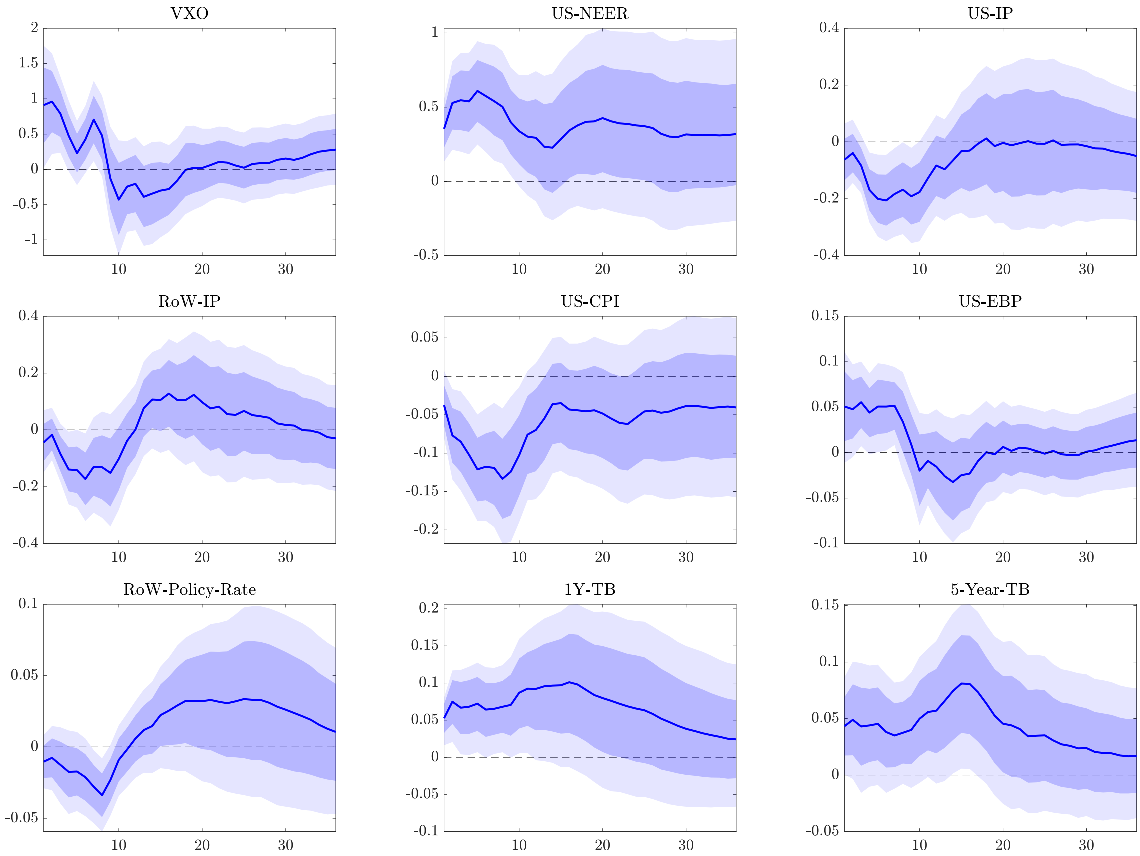
Note: The figure presents the impulse responses to a one-standard deviation US monetary policy shock. Horizontal axis measures time in months, vertical axis deviation from pre-shock level. Blue solid line represents point-wise posterior mean and shaded areas 68%/90% equal-tailed, point-wise credible sets.

Figure C.15: Responses to a contractionary US forward guidance shock



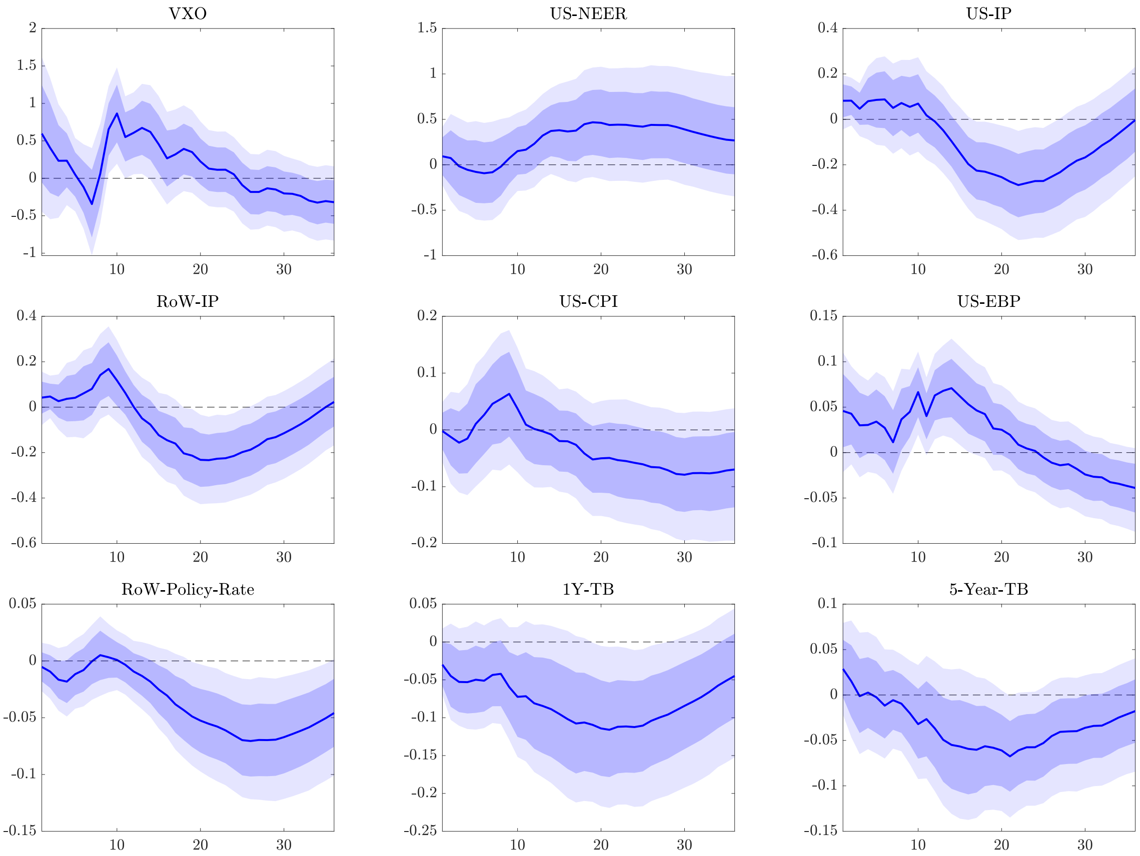
Note: The figure presents the impulse responses to a one-standard deviation US forward guidance shock. Horizontal axis measures time in months, vertical axis deviation from pre-shock level. Blue solid line represents point-wise posterior mean and shaded areas 68%/90% equal-tailed, point-wise credible sets.

Figure C.16: Responses to a contractionary conventional US monetary policy shock when estimation starts in 1996



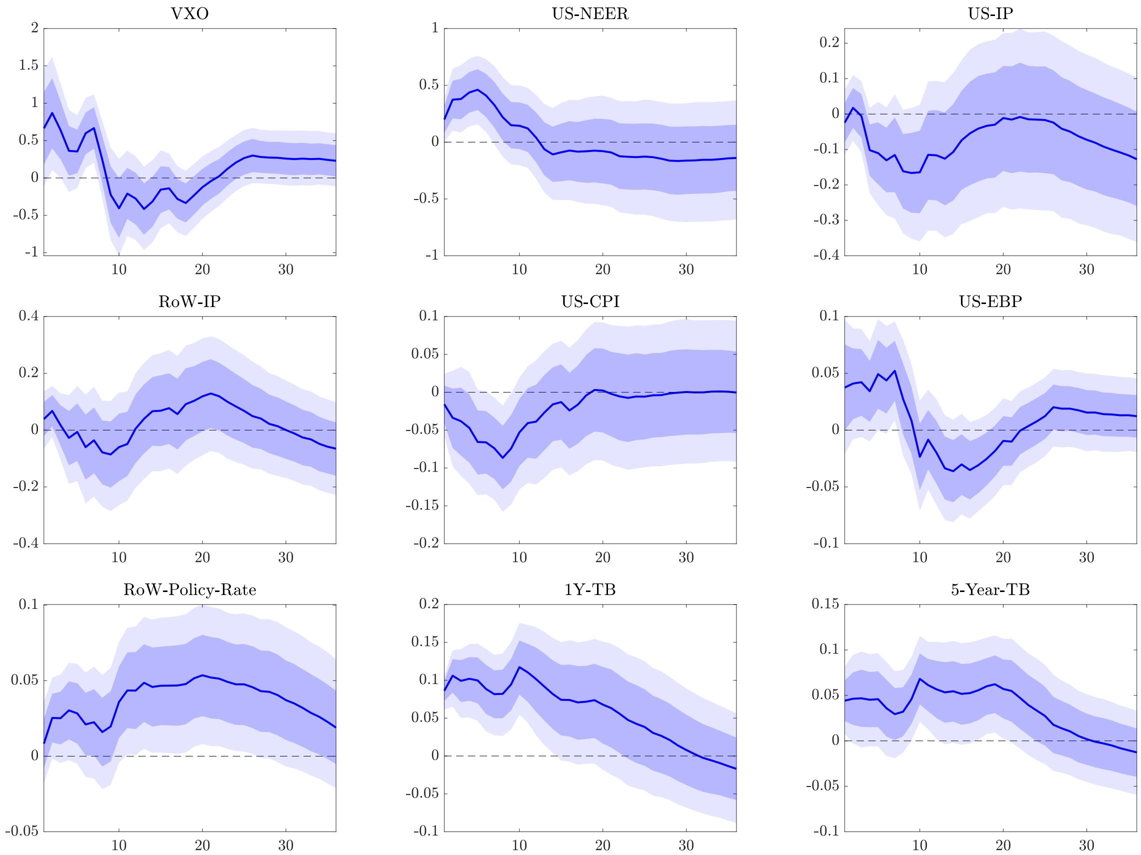
Note: The figure presents the impulse responses to a one-standard deviation US monetary policy shock when we start the estimation from 1996 and don't replace the pre 1996 missing values of the 5 year Treasury Bill futures with zeros. Horizontal axis measures time in months, vertical axis deviation from pre-shock level. Blue solid line represents point-wise posterior mean and shaded areas 68%/90% equal-tailed, point-wise credible sets.

Figure C.17: Responses to a contractionary US forward guidance shock when estimation starts in 1996



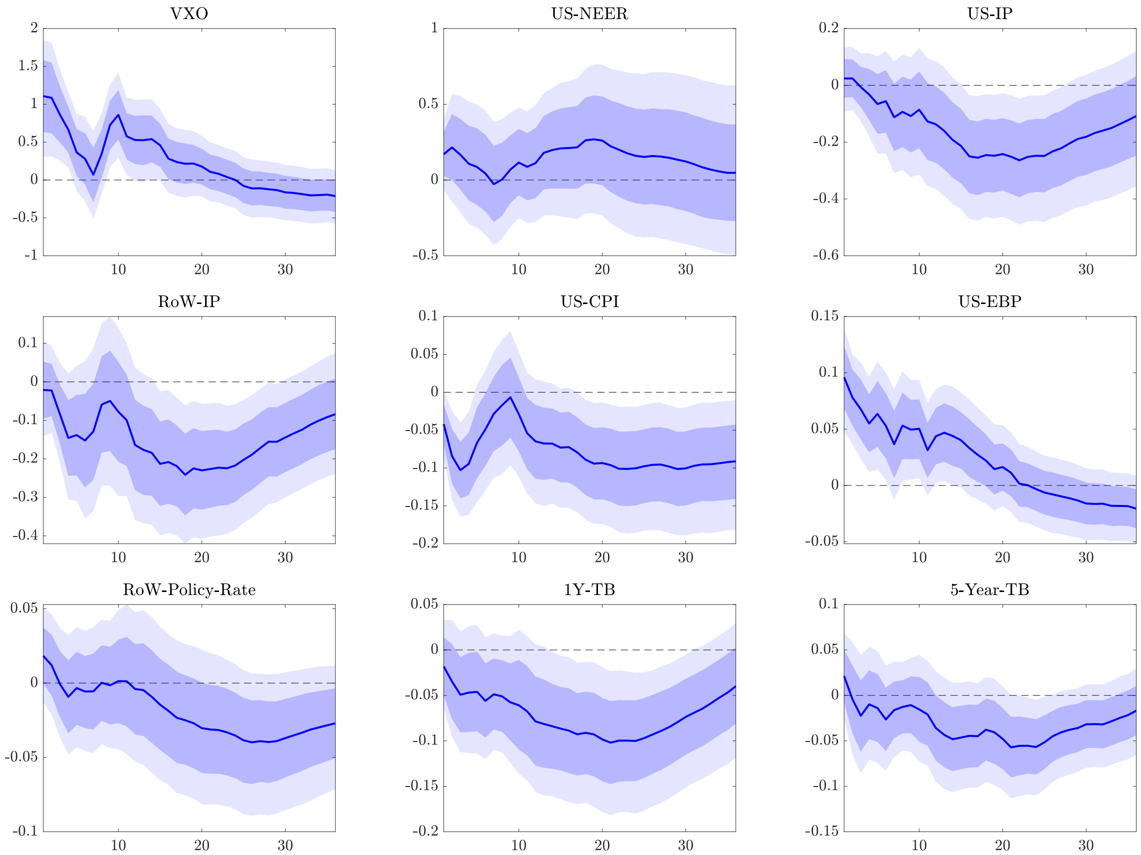
Note: The figure presents the impulse responses to a one-standard deviation US forward guidance shock when we start the estimation from 1996 and don't replace the pre 1996 missing values of the 5 year Treasury Bill futures with zeros. Horizontal axis measures time in months, vertical axis deviation from pre-shock level. Blue solid line represents point-wise posterior mean and shaded areas 68%/90% equal-tailed, point-wise credible sets.

Figure C.18: Responses to a contractionary conventional US monetary policy shock when using the proxies of Jarocinski (2021)



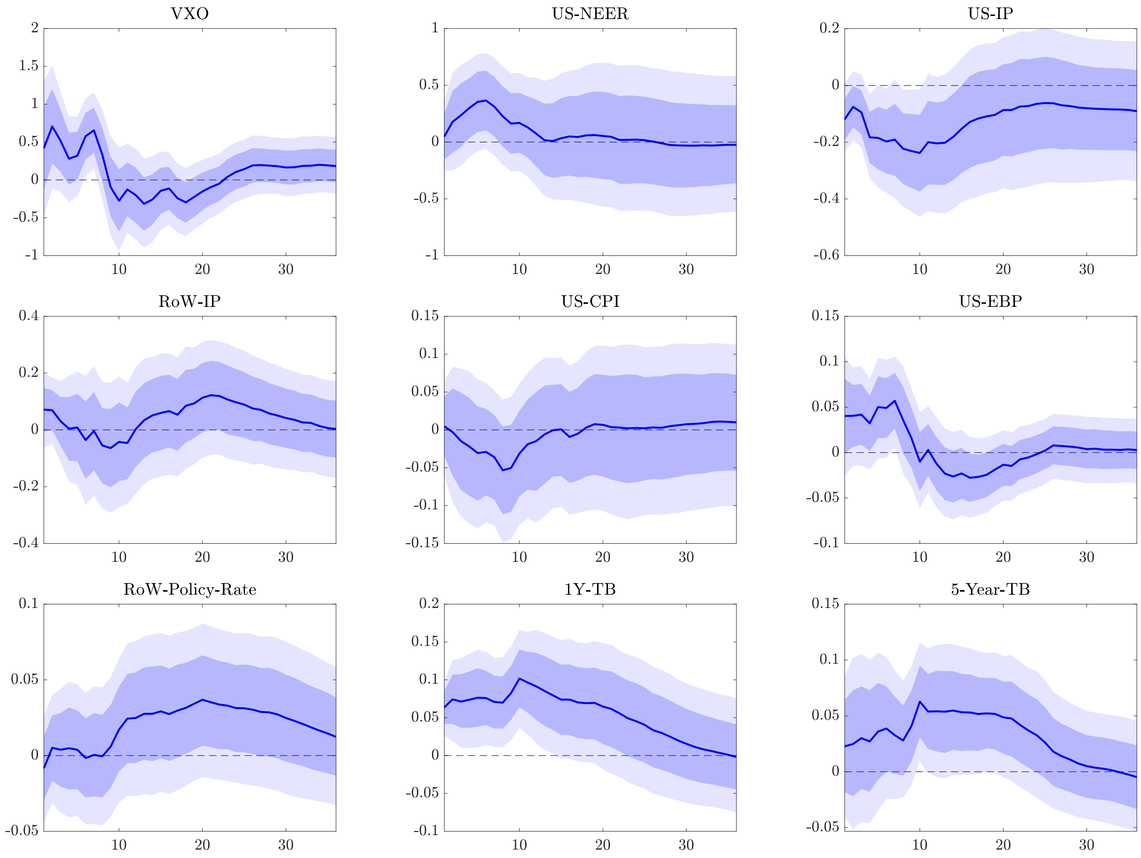
Note: The figure presents the impulse responses to a one-standard deviation US monetary policy shock when using the monetary policy proxies provided in Jarocinski (2021) instead of the raw surprises. Horizontal axis measures time in months, vertical axis deviation from pre-shock level. Blue solid line represents point-wise posterior mean and shaded areas 68%/90% equal-tailed, point-wise credible sets.

Figure C.19: Responses to a contractionary US forward guidance shock when using the proxies of Jarocinski (2021)



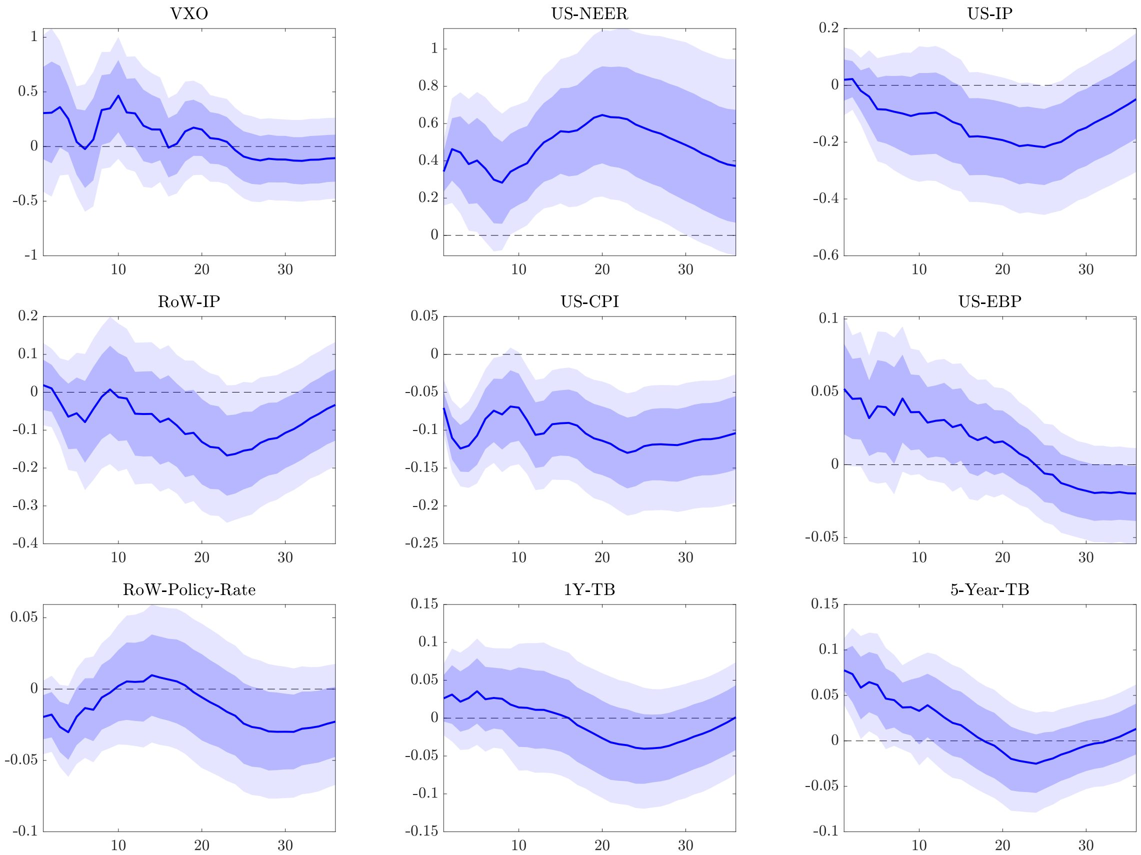
Note: The figure presents the impulse responses to a one-standard deviation US FG shock when using the monetary policy proxies provided in Jarocinski (2021) instead of the raw surprises. Horizontal axis measures time in months, vertical axis deviation from pre-shock level. Blue solid line represents point-wise posterior mean and shaded areas 68%/90% equal-tailed, point-wise credible sets.

Figure C.20: Responses to a contractionary conventional US monetary policy shock when using the proxies of Lewis (2024)



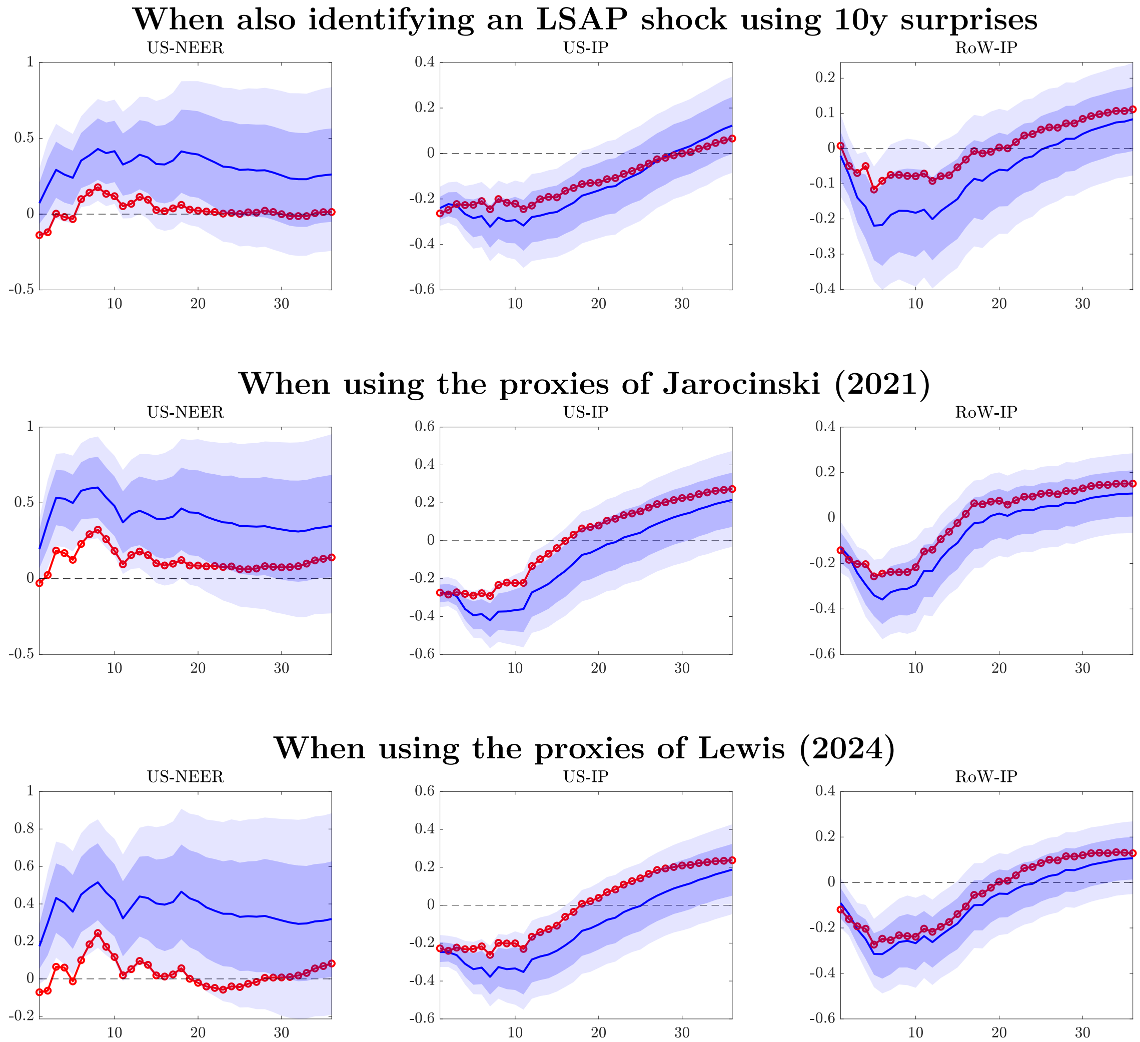
Note: The figure presents the impulse responses to a one-standard deviation US monetary policy shock when using the monetary policy proxies provided in Lewis (2024) instead of the raw surprises. Horizontal axis measures time in months, vertical axis deviation from pre-shock level. Blue solid line represents point-wise posterior mean and shaded areas 68%/90% equal-tailed, point-wise credible sets.

Figure C.21: Responses to a contractionary US forward guidance shock when using the proxies of Lewis (2024)



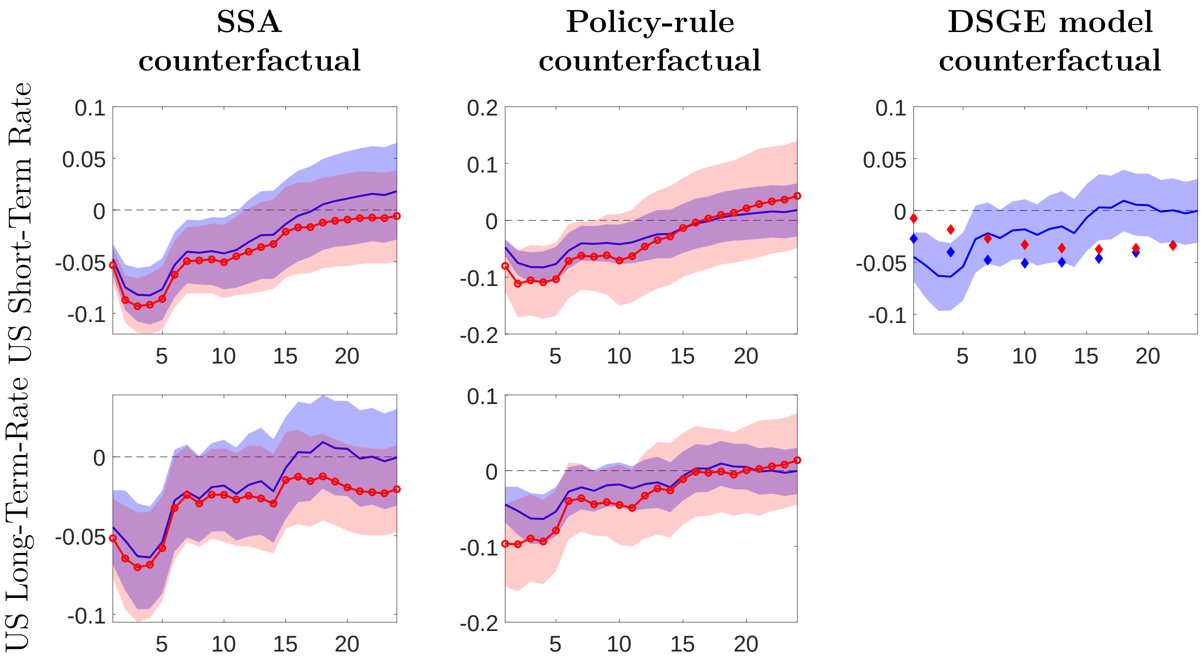
Note: The figure presents the impulse responses to a one-standard deviation US FG shock when using the monetary policy proxies provided in Lewis (2024) instead of the raw surprises. Horizontal axis measures time in months, vertical axis deviation from pre-shock level. Blue solid line represents point-wise posterior mean and shaded areas 68%/90% equal-tailed, point-wise credible sets.

Figure C.22: Robustness for the Policy Rule counterfactual



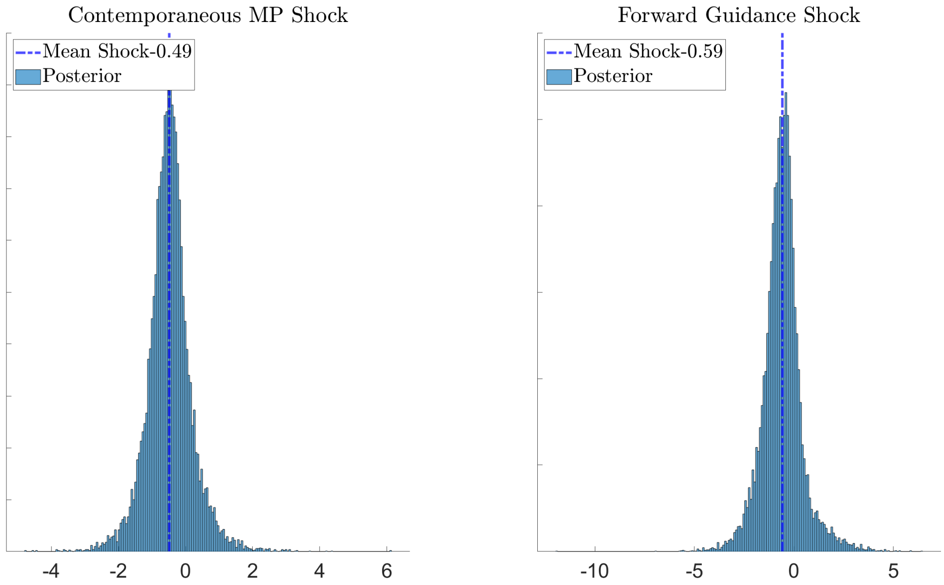
Note: The figures plots the estimated IRFs to the global risk (blue) against the pointwise mean of the IRFs under the counterfactual policy rule, where the FED commits to stabilizing the US-$. Note that, in order to stabilize the dollar perfectly over the entire impulse response horizon using the approach of McKay and Wolf (2023), one would need to identify 36 different policy shocks. As we only identify 2 shocks, we compute the least squares solution to the problem as suggested in McKay and Wolf (2023)

Figure C.23: Baseline and counterfactual responses US policy rates



Note: See notes to Figure 4 in the main text. In the BPSVAR (trinity model) we treat the 1-year (3-month) policy rate as the short-term policy indicator consistent with the empirical (theoretical) literature. Furthermore, in the BPSVAR we treat the 5-year treasury bill rate as the "long-term-rate". As in the trinity model, we do not explicitly model long-term rates, the corresponding panel in the second row is intentionally left blank.

Figure C.24: Fed policy shocks used to enforce the VAR-based policy-rule counterfactual



Note: The figure plots the posterior distribution of the policy shocks needed to enforce the counterfactual policy rule in the empirical counterfactual, where the Fed stabilizes the dollar. Shocks are measured in terms of standard deviation units and a negative sign corresponds to an expansionary shock.

C.3 Additional tables

Table C.1: Data description

Variable	Description	Source	Coverage
US 1-year TB rate	1-year Treasury Bill yield at constant maturity	US Treasury/Haver	1990m1 - 2019m12
US IP	Industrial production excl. construction	FRB/Haver	1990m1 - 2019m12
US CPI	US consumer price index	BLS/Haver	1990m1 - 2019m12
US EBP		Favara et al. (2016)	
US dollar NEER	Nominal broad trade-weighted Dollar index	FRB/Haver	1990m1-2019m12
VXO	CBOE market volatility index VXO	Wall Street Journal/Haver	1990m1 - 2019m12
RoW IP	Industrial production, see Martínez-García et al. (2015)	Dallas Fed Global Economic Indicators/Haver	1990m1 - 2019m12
RoW CPI	Consumer price index	Dallas Fed Global Economic Indicators/Haver (Martínez-García et al. 2015)	1990m1 - 2019m12
RoW policy rate	Short-term official/policy rate, see Martínez-García et al. (2015)	Dallas Fed Global Economic Indicators/Haver	1990m1 - 2019m12
Yen, euro, Swiss franc, British pound NEER	Nominal broad effective exchange rate	J.P. Morgan/Haver	1990m1-2019m12
US real exports	Exports of goods and services (chnd. 2012$)	BEA/Haver	1990q1-2019q2, interpolated to monthly frequency
US real imports	Imports of goods and services (chnd. 2012$)	BEA/Haver	1990q1-2019q2, interpolated to monthly frequency
Non-US USD cross-border bank credit	Banks' external liabilities in USD of banks owned by the world less externalliabilities in USD of banks owned by US nationals	BIS Locational Banking Statistics, Table A7/Haver	1990q1-2019q2, interpolated to monthly frequency
Non-US non-USD cross-border bank credit	Banks' external liabilities in non-USD of banks owned by the world lessexternal liabilities in non-USD of banks owned by US nationals	BIS Locational Banking Statistics, Table A7/Haver	1990q1-2019q2, interpolated to monthly frequency
EMBI spread	EMBI Brady bonds sovereign spread	JP Morgan Emerging Markets Bond Indexes /Haver	1990m1-2019m12
International debt securities	Debt securities issued outside of the resident's home market	BIS International Debt Issuance Statistics/Haver	1990q1-2019q4, interpolated to monthly frequency
AE and EME IP	Industrial production, see Martínez-García et al. (2015)	Dallas Fed Global Economic Indicators/Haver	1990m1 - 2019m12
AE and EME CPI	Consumer price index, see Martínez-García et al. (2015)	Dallas Fed Global Economic Indicators/Haver	1990m1 - 2019m12
AE and EME policy rate	Short-term official/policy rate, see Martínez-García et al. (2015)	Dallas Fed Global Economic Indicators/Haver	1990m1 - 2019m12
US dollar AE NEER	Nominal broad trade-weighted Dollar index against AEs	FRB/Haver	1990m1-2019m12
US dollar EME NEER	Nominal broad trade-weighted Dollar index against EMEs	FRB/Haver	1990m1-2019m12
US Treasury premium	Defined as the deviation from covered interest parity between US and G10 government bond yields	Du et al. (2018)	1991m4-2019m12
Commercial banks' Treasury and agency securities	Used for calculation of liquidity ratio	FRB/Haver	1990m1-2019m12
Total reserve balances with Federal Reserve banks	Used for calculation of liquidity ratio	FRB/Haver	1990m1-2019m12
Total demand deposits	Used for calculation of liquidity ratio	FRB/Haver	1990m1-2019m12
Financial commercial paper outstanding	Used for calculation of liquidity ratio	FRB/Haver	2001m1-2019m12
S&P 500	S&P 500 Composite	S&P/Haver	1990m1 - 2019m12
MSCI World excl. US	MSCI world excluding US	MSCI/Bloomberg	1990m1 - 2019m12
Macroeconomic uncertainty		Jurado et al. (2015)	1990m1 - 2019m12

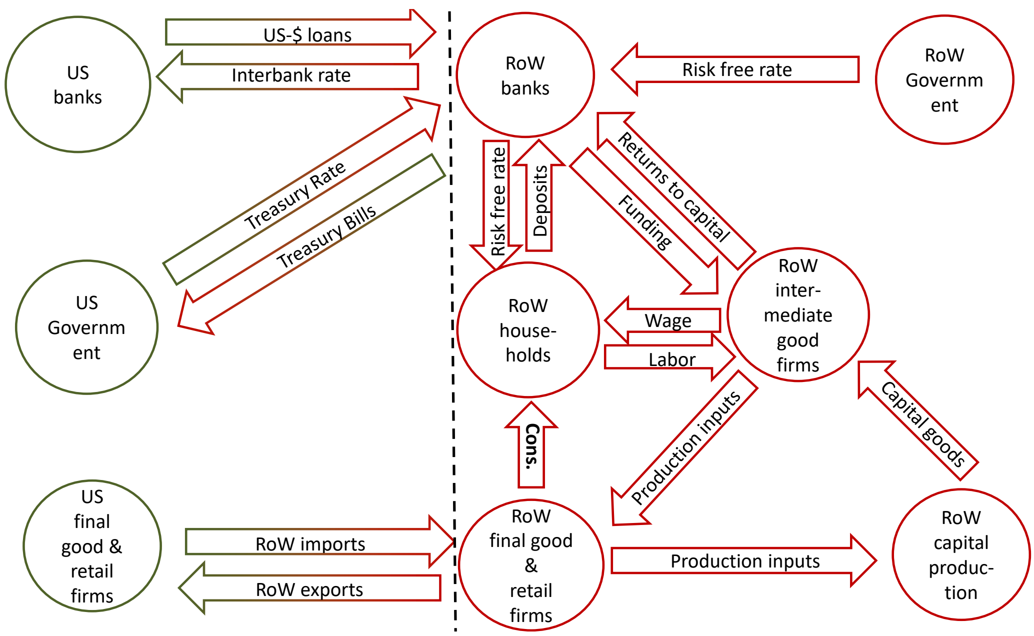
Notes: BLS stands for Bureau of Labour Statistics, FRB for Federal Reserve Board, BEA for Bureau of Economic Analysis, and BIS for Bank for International Settlements.

C.4 Additional model details

C.4.1 Model structure

The model of Georgiadis et al. (2023) consists of two economies, the US denoted by U, and a RoW block denoted by R, which are linked through trade, cross-border bank lending and investment in US Treasuries. The model features standard real and nominal frictions such as sticky prices and wages, habit formation in consumption, investment adjustment costs and variable capital utilization. At the heart of the model are US and RoW banks that engage in leveraged domestic and cross-border lending and borrowing. We assume the structure of frictions is (up to parametrization) symmetric for the US and the RoW; the key exceptions are financial frictions and global trade. In particular on the financial side, we assume US banks intermediate domestic dollar funds to the RoW and that US Treasuries are the global safe asset. Regarding international trade we assume that (i) for trade between the US and the RoW is largely prices are largely sticky in US$ and (ii) a fraction of intra RoW trade is also sticky in US$. The latter comes about because the RoW block is supposed to be an aggregate of countries and as document by Gopinath et al. (2020b), even if the US is not directly involved in the trade, countries tend to invoice a lot of their trade in US$. Figure C.25 gives a schematic overview of the model structure. As frictions are largely symmetric for the two blocks, we lay out the equations for the RoW block unless indicated otherwise. Generally, the exposition closely follows the model description in Georgiadis et al. (2023).

Figure C.25: Schematic overview of the model



C.4.2 Households and unions

In each period a household consumes a non-traded final good subject to habit formation in consumption. Furthermore each households is a monopolistic supplier of a differentiated labor service $L_{R,t}(h)$ and sells this to a perfectly competitive union that transforms it into an aggregate labor supply using a constant elasticity of substitution (CES) technology. Households satisfy demand for labor given the wage rate $W_{R,t}$, with wage setting being subject to frictions à la Calvo. The period-by-period utility function is given by

$$U(C_{R,t}, L_{R,t}) = \frac{1}{1-\sigma^c}(C_{R,t} - h_R C_{R,t-1})^{1-\sigma^c} - \frac{\kappa_{R,w}}{1+\varphi} L_{R,t}^{1+\varphi}. \tag{C.2}$$

with $\sigma^c, \varphi, h_R, \kappa_{R,w}$ as the intertemporal elasticity of substitution, the inverse Frisch elasticity of labor supply, the habit formation parameter and an exogenous labor scale parameter respectively. Households maximize utility subject to the following budget constraint

$$\frac{B^n_{R,t}}{P^C_{R,t}} + C_{R,t} = \frac{B^n_{R,t-1} R_{R,t-1}}{P^C_{R,t}} + \frac{W_{R,t}(h) L_{R,t}(h) + IS_{R,t}(h)}{P^C_{R,t}} + \frac{\Pi^C_{R,t}}{P^C_{R,t}} + \frac{\Pi^R_{R,t}}{P^C_{R,t}},$$

where we chose the final consumption and investment good price $P^C_{R,t}$ as the numeraire. $R_{R,t-1}$ is the predetermined domestic risk-free rate paid on nominal deposits with domestic banks $B^n_{R,t}$. $IS_{R,t}$ furthermore denotes an income stream from domestic state-contingent securities ensuring that all households will choose the same consumption and savings plans, despite temporarily receiving different wages due to the assumption of Calvo-type wage setting. Lastly $\Pi^C_{R,t}$ and $\Pi^R_{R,t}$ represent nominal profits from domestic (RoW) capital producing and retail firms respectively. The first-order condition of the household with respect to the choice of consumption is given by

$$\Lambda_{R,t} = (C_{R,t} - h_R C_{R,t-1})^{-\sigma_c} - \beta_R h_R \mathbb{E}_t[(C_{R,t+1} - h_R C_{R,t})^{-\sigma_c}] \tag{C.3}$$

with $\Lambda_{R,t}$ as the marginal utility of consumption. The intertemporal optimality conditions for the individual holdings of deposits with the local bank reads as

$$\Lambda_{R,t} = \mathbb{R}_t \Big[\beta_R \Lambda_{R,t+1} \frac{R_{R,t}}{1+\pi^C_{R,t+1}}\Big]. \tag{C.4}$$

where $\pi^C_{R,t+1}$ corresponds to the net inflation rate of the final consumption good. The working part of the household also sells its differentiated labor services $L_{R,t}(h)$ to a competitive union, which combines the differentiated labor services into a composite labor good using CES technology. Lastly the union leases the combined labor service to the intermediate good firms at the aggregate nominal wage rate $W_{R,t}$. The worker optimally chooses its wage given labor demand by the union taking into account that wage setting is subject to frictions à la Calvo, meaning that in each period they face a constant probability $(1-\theta_{w,R})$

of being able to adjust their nominal wage. As such the aggregate real wage index evolves as

$$w_{R,t}^{1-\psi_w} = (1-\theta_{w,R})\tilde{w}_{R,t}^{1-\psi_w} + \theta_{w.R}(1+\pi_{R,t}^{C})^{\psi_w-1} w_{R,t-1}^{1-\psi_w} \tag{C.5}$$

with $\tilde{w}_{R,t}$ as the optimal reset wage and $w_{R,t}$ as the economy wide real wage.

C.4.3 RoW financial intermediaries

RoW banks

We assume RoW banks raise funds through domestic deposits and cross-border dollar loans from US banks and use them to finance claims on domestic capital and holdings of US Treasuries. Specifically, consider RoW bank j and let $K_{R,j,t}$ be its claims on domestic capital in period t, $Q_{R,t}$ the price of such a claim relative to the price of the RoW final consumption good $P_{R,t}^{C}$, $GB_{R,j,t}$ holdings of US Treasuries, $B_{R,j,t}$ deposits from households, $CBDL_{R,j,t}$ funding through cross-border dollar loans, and $N_{R,j,t}$ net worth. The bank's balance sheet identity in real terms is

$$Q_{R,t}K_{R,j,t} + RER_t GB_{R,j,t} = B_{R,j,t} + RER_t CBDL_{R,j,t} + N_{R,j,t}, \tag{C.6}$$

where $RER_t = \mathcal{E}_t P_{U,t}^{C}/P_{R,t}^{C}$ represents the real exchange rate in terms of relative consumer-price levels and $\mathcal{E}_t$ the nominal exchange rate defined as the price of a dollar in units of RoW currency; an increase in $\mathcal{E}_t$ thus represents an appreciation of the dollar.

On the asset side of the RoW bank's balance sheet in Equation (C.6), claims on domestic capital $K_{R,j,t}$ earn the rate $R_{R,t}^{K}$, and—when converted to RoW currency—holdings of US Treasuries $GB_{R,j,t}$ earn the rate $D\mathcal{E}_t R_{U,t-1}^{GB}$, $D\mathcal{E}_t \equiv \mathcal{E}_t/\mathcal{E}_{t-1}$. On the liability side, deposits of domestic households $B_{R,j,t}$ cost the rate $R_{R,t-1}$—which we assume equals the RoW risk-free, central bank rate—and cross-border dollar loans from US banks $CBDL_{R,j,t}$ the rate $D\mathcal{E}_t R_{U,t-1}^{CBDL}$. The law of motion for the RoW bank's net worth is

$$\begin{aligned} N_{R,j,t} = \frac{1}{1+\pi_{R,t}^{C}}\Big\{ R_{R,t-1}N_{R,j,t-1} + \Big[&(1-\alpha_{R,j,t-1}^{GB})R_{R,t}^{K} + \alpha_{R,j,t-1}^{GB} D\mathcal{E}_t R_{U,t-1}^{GB} \\ &- (1-\ell_{R,j,t-1}^{CBDL})R_{R,t-1} - \ell_{R,j,t-1}^{CBDL} D\mathcal{E}_t R_{U,t-1}^{CBDL}\Big] AS_{R,j,t-1}\Big\}, \end{aligned} \tag{C.7}$$

where $AS_{R,j,t} \equiv Q_{R,t}K_{R,j,t} + RER_t GB_{R,j,t}$ denotes the bank's total assets, and $\alpha_{R,j,t}^{GB} \equiv RER_t GB_{R,j,t}/AS_{R,j,t}$ the share of total assets accounted for by US Treasuries, and $\ell_{R,j,t}^{CBDL} \equiv RER_t CBDL_{R,j,t}/AS_{R,j,t}$ the share of total assets funded by cross-border dollar loans.

Equation (C.7) shows that a RoW bank's net worth generally fluctuates with the dollar exchange rate. In particular, even when returns on US Treasuries equal the costs of cross-border dollar loans ($R_{U,t-1}^{GB} = R_{U,t-1}^{CBDL}$), if the share of assets funded by cross-border dollar loans exceeds the asset share of Treasuries ($\ell_{R,j,t}^{CBDL} - \alpha_{R,j,t}^{GB} > 0$) the bank's net worth drops when the dollar appreciates ($D\mathcal{E}_t > 0$).

The objective of a RoW bank is to maximize the discounted value of current and expected future equity streams. The bank's value function is

$$V_{R,j,t} = \max\ \mathbb{E}_t \sum_{s=0}^{\infty} (1-\theta_B)\Theta_{R,t,t+s} N_{E,j,t+1+s}, \tag{C.8}$$

where $\Theta_{R,t,t+s}$ is the household's real stochastic discount factor.

In order to put a ceiling on the amount of leverage a RoW bank can take on we assume it faces a balance-sheet constraint in the spirit of Gertler and Karadi (2011). We motivate this balance-sheet constraint as an eligibility requirement the bank needs to satisfy in order for creditors to provide funding. In particular, for the bank to attract creditors and be able to leverage, the sum of its discounted current and expected future equity streams has to be larger than a risk-weighted sum of its current assets

$$V_{R,j,t} \geq \delta_{R,j,t}(Q_{R,j,t}K_{R,j,t} + \Gamma_R^{GB} RER_t GB_{R,j,t}). \tag{C.9}$$

We assume creditors apply two types of risk weights in the balance-sheet constraint in Equation (C.9). First, the *asset-specific* risk weight Γ_R^{GB} represents the perceived riskiness of Treasuries relative to claims on domestic capital (for a similar interpretation see Karadi and Nakov 2021; Coenen et al. 2018). In particular, we assume that US Treasuries are perceived to be less risky than claims on domestic capital ($\Gamma_R^{GB} < 1$).

Second, the *balance-sheet-specific* risk weight $\delta_{R,j,t}$ represents the perceived riskiness of the bank's relative *asset and liability* composition. The balance-sheet constraint in Equation (C.9) thus shows how creditors weigh the perceived riskiness of the size and structure of the bank's asset and liability portfolio on the right-hand side against its discounted current and expected future level of equity on the left-hand side.[96] In particular, we assume the balance-sheet-specific risk weight varies with the asset and liability shares according to

$$\delta_{R,j,t}\left(\ell_{R,j,t}^{CBDL}, \alpha_{R,j,t}^{GB}\right) = \bar{\delta}_R \left[1 + \frac{\kappa_{R,\alpha,\ell}}{2}\left(\alpha_{R,j,t}^{GB} - \ell_{R,j,t}^{CBDL}\right)^2 - \epsilon_{R,\alpha}\alpha_{R,j,t}^{GB}\right] + \epsilon_t^{\delta_R}, \tag{C.10}$$

where $\epsilon_t^{\delta_R}$ is an exogenous shock which we interpret as a shock to the willingness of creditors to provide funding for a given level of net worth. In other words we assume that this shock raises the risk aversion of creditors. Because we are interested in a *global* risk aversion shock, we assume that for each country i, the shock $\epsilon^{\delta_{i,B}}$ has a factor structure with a domestic component $\eta_t^{\delta_i}$ and a global component $\eta_t^{\delta_G}$ and evolves as

$$\epsilon_t^{\delta_{i,B}} = \rho_\delta \epsilon_{t-1}^{\delta_{i,B}} + \eta_t^{\delta_i} + \eta_t^{\delta_G}. \tag{C.11}$$

[96]The balance-sheet constraint in Equation (C.9) is algebraically very similar to that postulated in Gertler and Karadi (2011), who motivate it referring to the idea that the banker can 'abscond' with a fraction of assets.

The specification of the balance-sheet-specific risk weight in Equation (C.10) is key for the transmission mechanisms in the model. First, cross-border dollar loan funding increases the balance-sheet-specific risk weight as long as it is not met by corresponding dollar assets in terms of holdings of US Treasuries ($\kappa_{R,\alpha,\ell} > 0$). We make this assumption because unhedged cross-border dollar borrowing exposes the RoW bank's net worth to fluctuations in the exchange rate and dollar funding shortages.[97] Second, apart from hedging funding through cross-border dollar loans, holding US Treasuries reduces the balance-sheet-specific risk weight ($\epsilon_{R,\alpha} > 0$). We make this assumption because Treasuries are viewed as the safe asset whose market price is relatively stable so that it can be sold with limited losses or even gains on its face value in times of stress in order to provide liquidity buffer in any type of funding shortage (Bianchi et al. 2021). In other words, Equation (C.10) incorporates a general and a dollar-specific incentive for holding Treasuries as liquidity-buffer.[98]

It can be shown that the value function of a bank, just like the law of motion its equity, is linear in its components. In particular after guessing the value function can be written as

$$V_{R,j,t} = \Big[(1-\alpha^{GB}_{R,j,t})v_{R,t} + \alpha^{GB}_{R,j,t}v^{GB}_{R,t} - \ell_{R,j,t}u_{R,t})\Big]AS_{R,j,t} + n_{R,t}N_{R,j,t} \tag{C.12}$$

its possible to verify procedure the solution to the bankers problem can be characterized by the following set of equations.

$$v_{R,t} = \mathbf{E}_t\Big(\Omega_{R,t,t+1}(R_{R,k,t+1} - R_{R,t})\Big) \tag{C.13}$$

$$v^{GB}_{R,t} = \mathbf{E}_t\Big(\Omega_{R,t,t+1}(\mathcal{E}_{t+1}/\mathcal{E}_t R^{GB}_{R,t} - R_{R,t})\Big) \tag{C.14}$$

$$n_{R,t} = \mathbf{E}_t\Big(\Omega_{E,t,t+1}(R_{R,t})\Big) \geq 1 \tag{C.15}$$

$$u_{R,t} = \mathbf{E}_t\Big(\Omega_{R,t,t+1}\mathcal{E}_{t+1}/\mathcal{E}_t R^{CBDL}_{U,t} - R_{E,t}\Big) \tag{C.16}$$

$$\Omega_{R,t,t+1} = \mathbf{E}_t\Big(\frac{\beta_R\Lambda_{R,t+1}}{\Lambda_{R,t}(1+\pi^c_{R,t+1})}\Big[(1-\theta_B) \tag{C.17}$$

$$+\theta_B([v_{R,t+1}(1-\alpha^{GB}_{R,j,t+1}) + v^{GB}_{R,t+1}\alpha^{GB}_{R,j,t+1} - u_{R,t+1}\ell^{CBDL}_{R,j,t+1}]\phi_{R,j,t+1} + n_{R,t+1})\Big]$$

Equations C.13, C.14, C.15, C.16, represent the discounted excess returns from borrowing domestically and lending domestically, the discounted excess returns from borrowing domestically and investing into US government bonds, the discounted excess costs of borrowing in US-$ instead of acquiring domestic deposits and the discounted marginal value

[97] Under the 'absconding' interpretation of the balance-sheet constraint of Gertler and Karadi (2011) this assumption entails that the amount of assets the bank can run away with increases with the unhedged share of funding through cross-border dollar loans. This assumption may be motivated by the observation that bankruptcy laws are biased towards domestic lenders (Akinci and Queralto 2019).

[98] Note that strictly speaking Equation (C.10) states that also a positive net dollar exposure ($\alpha^{GB}_{R,j,t} - \ell^{CBDL}_{R,j,t} > 0$) increases the balance-sheet-specific risk weight. Thus, Equation (C.10) can also be read as stating that the bank has an incentive to take on cross-border dollar loans to hedge holdings of Treasuries. However, as we discuss in the calibration below, in the steady state the bank has a *negative* net dollar exposure.

of an additional unit of equity. Equation C.17 is the bankers "augmented" real stochastic discount factor, which accounts for marginal value of funds internal to the financial intermediary and the fact that the bank may have to close with a probability of $1-\theta_B$. Lastly $\phi_{R,j,t} = AS_{R,j,t}/N_{R,j,t}$ corresponds to the optimal leverage ratio of the RoW bank described below.

With a closed form solution for $V_{R,j,t}$ at hand its straightforward to derive the first order conditions taking into account the balance sheet constraint in Equation (C.9).Regarding the choice of the optimal composition of asset side its possible to show that this the following first order condition.

$$\mathbf{E}_t \left[\Omega_{R,j,t,t+1} \left(D\mathcal{E}_{t+1} R^{GB}_{U,t} - R_{R,t}\right)\right] + CY_{R,j,t} = RP^{GB}_{R,j,t}. \tag{C.18}$$

The first term on the left-hand side coincides with the UIP condition in a standard model without financial frictions and safe asset demand. In particular, in a standard setup, in order to rule out arbitrage profits for RoW banks the exchange-rate-adjusted return of Treasuries—whose dollar-return equals the US risk-free rate $R^{GB}_{U,t} = R_{U,t}$ by assumption—has to equal the cost of funding through domestic deposits in terms of the risk-free rate $R_{R,t}$. Equation (C.18) shows that our model gives rise to two UIP deviations $CY_{R,j,t}$ and $RP^{GB}_{R,j,t}$.

The first UIP deviation is given by

$$RP^{GB}_{R,j,t} = \Gamma^{GB}_R \mathbf{E}_t \Big[\Omega_{R,j,t,t+1} \Big(R^K_{R,t+1} - R_{R,t}\Big)\Big], \tag{C.19}$$

and arises because optimal portfolio choice requires that in equilibrium the overall, exchange-rate-adjusted excess return of US Treasuries on the left-hand side in Equation (C.18) has to equal the risk-weight-adjusted excess return of the alternative investment in domestic capital on the right-hand side in Equation (C.18).

The second UIP deviation is given by

$$CY_{R,j,t} = -\frac{\partial \delta_{R,j,t}/\partial \alpha^{GB}_{R,j,t}}{\delta_{R,j,t}} \left[(1-\alpha^{GB}_{R,j,t}) + \Gamma^{GB}_R \alpha^{GB}_{R,j,t}\right] \mathbf{E}_t \Big[\Omega_{R,j,t,t+1} \Big(R^K_{R,t+1} - R_{R,t}\Big)\Big], \tag{C.20}$$

and arises because in our setup the *overall* return of US Treasuries for a RoW bank on the left-hand side is made up of the direct component $R^{GB}_{U,t}$ and an additional, *indirect* component: Holding Treasuries loosens a RoW bank's balance-sheet constraint in Equations (C.9) and (C.10), thereby allows it to leverage more, exploit more investment opportunities and generate additional profits. In other words, because of their dominant safe asset property, holding Treasuries may be optimal for a RoW bank even if their direct, expected, exchange-rate-adjusted return is lower than the risk-weight-adjusted return of domestic capital $RP^{GB}_{R,j,t}$. We interpret this indirect return $CY_{R,j,t}$ as a convenience yield.

Equation (C.20) shows that the magnitude of the convenience yield is pinned down by the degree to which holding Treasuries reduces a RoW bank's balance-sheet-specific risk weight, how the freed leverage translates into additional claims on domestic capital, and the corresponding excess return. For example, when domestic credit spreads are high, holding additional Treasuries and thereby relaxing a RoW bank's balance-sheet constraint is particularly profitable, and hence the convenience yield is high. Note that Equation (C.18) instills a structural interpretation to the convenience yield in the UIP condition in the no-arbitrage finance framework in Krishnamurthy and Lustig (2019). Apart from the risk premium $RP^{GB}_{R,j,t}$, Equation (C.18) also coincides with the UIP condition in the structural model of Jiang et al. (2023). However, in their setup the convenience yield is introduced *ad hoc* as a UIP deviation that is assumed to decline in the global stock of safe assets. In contrast, in our model the convenience yield and its relation to global financing conditions emerge endogenously from the optimal portfolio choice of RoW banks.

As a UIP condition Equation (C.18) pins down the evolution of the dollar exchange rate. First, for a given RoW domestic deposit rate ($R_{R,t}$), in standard UIP logic an increase in the US risk-free rate and hence by assumption the return on Treasuries ($R^{GB}_{U,t}$) requires an expected depreciation of the dollar ($D\mathcal{E}_{t+1}$ declines), which is in part achieved by a contemporaneous appreciation. Second, for a given RoW domestic deposit rate ($R_{R,t}$) and US risk-free rate ($R^{GB}_{U,t}$), an increase in the convenience yield ($CY_{R,j,t}$) has to be accompanied by an expected depreciation of the dollar ($D\mathcal{E}_{t+1}$ declines), which is again in part achieved by a contemporaneous appreciation.

Regarding the optimal choice of the liability composition, it can be shown that the total returns on cross border dollar loans $R^{CBDL}_{U,t}$ have to equal the costs of domestic funding up to an endogenous wedge.

$$\mathbf{E}_t\left(\Omega_{R,j,t,t+1}R_{R,t}\right) = \mathbf{E}_t\left(\Omega_{R,j,t,t+1}D\mathcal{E}_{t+1}R^{CBDL}_{U,t}\right) + RP^{CBDL}_{R,j,t}, \tag{C.21}$$

with

$$\begin{aligned} RP^{CBDL}_{R,j,t} = \frac{\partial\delta_{R,j,t}/\partial\ell^{CBDL}_{R,j,t}}{\delta_{R,j,t}}\mathbf{E}_t\Omega_{R,j,t,t+1}\Big[(1-\alpha^{GB}_{R,j,t})(R^K_{R,t+1}-R_{R,t}) \\ +\alpha^{GB}_{R,j,t}\left(D\mathcal{E}_{t+1}R^{GB}_{U,t}+CY_{R,j,t}-R_{R,t}\right)\Big]. \end{aligned} \tag{C.22}$$

Cross-border dollar borrowing has an additional, *indirect* cost, as it tightens the RoW bank's balance-sheet constraint in Equations (C.9) and (C.10), thereby limits its leverage and thus reduces profits. This risk premium implies that in order for the RoW bank to borrow cross-border dollar funds the *direct* cost has to be lower than for domestic deposits. Thus, consistent with the data, in our model cross-border dollar borrowing is—or at least appears to be—cheap compared to domestic funding (Caramichael et al. 2021; Gutierrez et al. 2023). Analogous to the UIP condition in Equation (C.18), also Equation (C.21) provides intuition for the evolution of the dollar exchange rate. For example, when global financing conditions tighten so that domestic credit spreads rise, the risk premium on cross-border dollar loans increases. Equation (C.21) shows that for a given deposit rate and cross-border dollar credit rate this rise in the risk premium has to be accompanied by an expected depreciation of the dollar. This is partly accomplished by a contemporaneous appreciation. This mechanism is similar to the "two-way feedback between balance sheets and exchange rates" in Akinci and Queralto (2019, p.3).

The remaining equations of the RoW banking block are fairly standard. In particular, we impose market clearing for domestic capital, US treasuries and specify the start-up funds for a newly entering bank n as a fraction of last period's portfolio, $N_{R,n,t} = \omega_R AS_{R,t-1}$. In equilibrium all banks choose the same portfolio structure as they face the same returns and costs. The law of motion for aggregate net worth of the RoW banking sector is given by

$$N_{R,t} = \frac{\theta_B}{1+\pi^C_{R,t}} \Bigg\{ R_{R,t-1} N_{R,t-1} + \Big[(1-\alpha^{GB}_{R,t-1}) R^K_{R,t} + \alpha^{GB}_{R,t-1} D\mathcal{E}_t R^{GB}_{U,t-1} \tag{C.23}$$

$$- \left(1 - \ell^{CBDL}_{R,t-1}\right) R_{R,t-1} - \ell^{CBDL}_{R,t-1} D\mathcal{E}_t R^{CBDL}_{U,t-1} \Big] AS_{R,t-1} \Bigg\} + \omega_R AS_{R,t-1}$$

When the model is parameterized so that the balance-sheet constraint in Equation (C.9) binds in a neighbourhood of the steady-state, the maximum equilibrium leverage ratio is given by

$$\phi_{R,j,t} \equiv \frac{AS_{R,j,t}}{N_{R,j,t}} = \frac{n_{R,j,t}}{\mathcal{R}_{R,j,t} - \mathcal{P}_{R,j,t}}, \tag{C.24}$$

where

$$\mathcal{R}_{R,j,t} \equiv \delta_{R,j,t} \left[(1-\alpha^{GB}_{R,j,t}) + \Gamma^{GB}_R \alpha^{GB}_{R,j,t} \right], \tag{C.25}$$

$$\mathcal{P}_{R,j,t} \equiv \mathbf{E}_t \Omega_{R,j,t,t+1} \Big[(1-\alpha^{GB}_{R,j,t}) R^K_{R,t+1} + \alpha^{GB}_{R,j,t} D\mathcal{E}_{t+1} R^{GB}_{U,t}$$

$$- (1-\ell^{CBDL}_{R,j,t}) R_{R,t} - \ell^{CBDL}_{R,j,t} D\mathcal{E}_{t+1} R^{CBDL}_{U,t} \Big], \tag{C.26}$$

are the RoW bank's asset-share-weighted bank and asset-specific risk weight and its expected profitability, respectively; the terms $\Omega_{R,j,t,t+1}$ and $n_{R,j,t}$ denote the bank's stochastic discount factor and the expected discounted returns to equity respectively. Equation (C.24) shows that the RoW bank's maximum leverage is pinned down by its portfolio's expected profitability and perceived riskiness in terms of risk weights. In particular, the RoW bank can attain a higher leverage ratio, thereby exploit more investment opportunities and generate more profits if (i) the perceived riskiness in terms of $\mathcal{R}_{R,j,t}$ is low, (ii) its expected profitability in terms of $\mathcal{P}_{R,j,t}$ is high, and/or (iii) expected discounted returns to equity in terms of $n_{R,j,t}$ are large.

C.4.4 US financial intermediaries

US banks differ from RoW banks in four ways. First, a US bank acts as cross-border lender rather than borrower, and so dollar loans appear on the asset side of its balance sheet

$$Q_{U,t}K_{U,j,t} + CBDL_{U,j,t} = B_{U,j,t} + N_{U,j,t}, \tag{C.27}$$

where $K_{U,j,t}$, $CBDL_{U,j,t}$, $B_{U,j,t}$ and $N_{U,j,t}$ are the total amount of claims on domestic capital, cross-border dollar loans, domestic deposits and net worth, respectively, deflated by the price of the US consumption good.

Second, for simplicity and in order to focus on the RoW, we assume US banks do not hold Treasuries. In contrast to RoW banks a US bank's net worth

$$\begin{aligned} N_{U,j,t} = & \frac{1}{1+\pi^C_{U,t}}\Big[(R^K_{U,t} - R_{U,t-1})Q_{U,t-1}K_{U,j,t-1} \\ & \qquad + (R^{CBDL}_{U,t-1} - R_{U,t-1})CBDL_{U,j,t-1} + R_{U,t-1}N_{U,j,t-1}\Big], \end{aligned} \tag{C.28}$$

is not affected by exchange rate valuation effects as its liabilities and assets are all denominated in dollar.

Third, for a US bank we assume the balance-sheet constraint

$$V_{U,j,t} \geq \delta_{U,j,t}(Q_{U,t}K_{U,j,t} + \Gamma^{CBDL}_{U,t}CBDL_{U,j,t}), \tag{C.29}$$

with the asset-specific risk weight creditors attach to cross-border dollar loans

$$\Gamma^{CBDL}_{U,t} = \bar{\Gamma}^{CBDL}_{U} + \Phi_{U,\phi}\phi_{R,j,t}, \tag{C.30}$$

and where $\phi_{R,j,t}$ is the leverage ratio of RoW banks from Equation (C.24). Specifically, in Equation (C.30) we assume cross-border dollar lending is perceived to be more risky by a US bank's creditors when RoW banks are more leveraged. The motivation for this specification is that while RoW banks lend to the US government (the least risky borrower by assumption) and US firms (which pledge the entire return to capital), US banks also lend to leveraged and thus risky RoW banks, whose leverage (and thereby riskiness) endogenously fluctuates with the state of the economy.

Fourth, in contrast to RoW banks, a US bank does not engage in foreign-currency borrowing so that there is no asset-liability currency mismatch creditors may be concerned about. Therefore, we assume the balance-sheet-specific risk weight $\delta_{U,j,t}$ for a US bank does not vary endogenously and is given by

$$\delta_{U,j,t} = \overline{\delta}_U + \epsilon^{\delta_U}_t, \tag{C.31}$$

where $\epsilon^{\delta_U}_t$ is an exogenous risk aversion shock discussed previously.

We assume for simplicity that the return on US Treasuries equals the risk-free, monetary policy rate: $R^{GB}_{U,t} = R_{U,t}$. [99].

As in the RoW case, the objective of the US banker is to maximize the discounted value of current and future equity streams subject to the balance sheet constraint. The bank's value function is

$$V_{U,j,t} = \max\ \mathbb{E}_t \sum_{s=0}^{\infty} (1-\theta_B)\Theta_{U,t,t+s} N_{U,j,t+1+s}, \tag{C.32}$$

where $\Theta_{U,t,t+s}$ is the household's real stochastic discount factor.

Defining $\alpha^{CBDL}_{U,j,t} = \frac{CBDL_{U,j,t}}{AS_{U,j,t}}$ as the asset ratio of cross border dollar loans to total assets of the banks assuming that the value function $V_{U,j,t}$ is linear in the components of the LOM for net worth its possible to show that

$$V_{U,j,t} = \Big[(1-\alpha^{CBDL}_{U,j,t})v_{U,t} + \alpha^{CBDL}_{U,j,t} v^{CBDL}_{U,t})\Big] AS_{U,j,t} + n_{U,t}N_{U,j,t} \tag{C.33}$$

$$v_{U,t} = \mathbb{E}_t\Big(\Omega_{U,t,t+1}(R^K_{U,t+1} - R_{U,t})\Big) \tag{C.34}$$

$$v^{CBDL}_{U,t} = \mathbb{E}_t\Big(\Omega_{U,t,t+1}(R^{CBDL}_{U,t} - R_{U,t})\Big) \tag{C.35}$$

$$n_{U,t} = \mathbb{E}_t\Big(\Omega_{U,t,t+1}(R_{U,t})\Big) \tag{C.36}$$

$$\begin{aligned} \Omega_{U,t,t+1} = \mathbb{E}_t\Big(& \frac{\Theta_{U,t,t+1}}{(1+\pi^c_{U,t+1})}\Big[(1-\theta_B) \\ & + \theta_B\Big([(1-\alpha^{CBDL}_{U,j,t+1})v_{U,t+1} + \alpha^{CBDL}_{U,j,t+1} v^{CBDL}_{E,t+1}]\phi_{U,j,t+1} + n_{U,t+1}\Big)\Big]\Big). \end{aligned} \tag{C.37}$$

With $V_{U,j,t}, v_{U,t}, v^{CBDL}_{U,t}, n_{U,t}$ and $\Omega_{F,t,t+1}$ as the slightly different versions of their RoW counterparts touched up on the previous section.

As in the RoW case the optimal portfolio choice of US banks choice requires

$$\Gamma^{CBDL}_{U,t}\mathbb{E}_t\Big[\Omega_{U,j,t,t+1}\Big(R^K_{U,t+1} - R_{U,t}\Big)\Big] = \mathbb{E}_t\Big[\Omega_{U,j,t,t+1}\Big(R^{CBDL}_{U,t} - R_{U,t}\Big)\Big] - RP^{CBDL}_{U,j,t}, \tag{C.38}$$

stating that the expected risk-weight-adjusted excess returns on domestic capital on the left-hand side and cross-border dollar loans on the right-hand side have to equalize.

Apart from the term $RP^{CBDL}_{U,j,t}$, Equation (C.38) coincides with the equilibrium condition in a standard model without financial frictions on cross-border dollar lending and borrowing. In particular, in a standard setup expected, risk-weight-adjusted returns of different assets have to equalize. In Equation (C.38) this means that the expected, risk-weight-adjusted excess returns on claims on domestic capital have to equal the expected excess returns on cross-border lending. Equation (C.38) shows that in our model the *direct* expected excess

[99] This would result endogenously if we assumed US banks can hold Treasuries, if the corresponding asset-specific risk weight in the balance-sheet constraint in Equation (C.29) was zero, and if the balance-sheet-specific risk weight in Equation (C.31) was independent of these holdings

return of cross-border dollar lending has to be higher than the risk-weight-adjusted excess return of claims on domestic capital due to a risk premium $RP_{U,j,t}^{CBDL}$.

In particular, this risk premium on cross-border lending is given by

$$RP_{U,j,t}^{CBDL} = \frac{\partial \Gamma_{U,t}^{CBDL}}{\partial \alpha_{U,j,t}^{CBDL}} \alpha_{U,j,t}^{CBDL} \mathbf{E}_t \Omega_{U,j,t,t+1} \Big[(1 - \alpha_{U,j,t}^{CBDL})(R_{U,t+1}^{K} - R_{U,t}) + \alpha_{U,j,t}^{CBDL}(R_{U,t}^{CBDL} - R_{U,t}) \Big], \tag{C.39}$$

and arises because the US bank's cross-border dollar lending raises the RoW bank's leverage, which feeds back and raises the US bank's asset-specific risk weight (see Equation (C.30)) and thereby has an additional, *negative indirect* return: It tightens the US bank's balance-sheet constraint in Equations (C.29) and (C.30), which limits its leverage and thus reduces profits.[100]

Equation (C.39) shows that the magnitude of this risk premium is pinned down by the degree to which cross-border dollar lending raises the US bank's asset-specific risk weight on cross-border dollar lending, how the ensuing reduction in the bank's leverage cuts into claims on domestic capital and cross-border dollar lending, and their corresponding excess returns. For example, when domestic credit spreads are high, the foregone profits implied by the tightening in the bank's balance-sheet constraint due to cross-border dollar lending are particularly high, and hence the risk premium on cross-border dollar lending is high.

The remaining equations of the US banking block are fairly standard. In particular, we impose market clearing for domestic capital, cross border dollar loans and specify the start-up funds for a newly entering bank n as a fraction of last period's portfolio, $N_{U,n,t} = \omega_U AS_{U,t-1}$. The law of motion for aggregate net worth of the US banking sector is given by

$$N_{U,t} = \frac{\theta_B}{1 + \pi_{U,t}^{C}} \Big\{ R_{U,t-1} N_{U,t-1} + \Big[(1 - \alpha_{U,t-1}^{CBDL})(R_{U,t}^{K} - R_{U,t-1}) + \alpha_{U,t-1}^{CBDL}(R_{U,t-1}^{GB} - R_{U,t-1}) \Big] AS_{U,t-1} \Big\} + \omega_U AS_{U,t-1} \tag{C.40}$$

When the model is parameterized so that the balance-sheet constraint in Equation (C.9) binds in a neighbourhood of the steady-state, the maximum equilibrium leverage ratio again reflects a risk-profitability trade-off

$$\phi_{U,j,t} \equiv \frac{AS_{U,j,t}}{N_{U,j,t}} = \frac{Q_{U,t} K_{U,j,t} + CBDL_{U,j,t}}{N_{U,j,t}} = \frac{n_{U,j,t}}{\mathcal{R}_{U,j,t} - \mathcal{P}_{U,j,t}}, \tag{C.41}$$

[100] Using the market clearing conditions alongside the balance sheets of the two banks it can be shown that $\frac{\partial \Gamma_{U,t}^{CBDL}}{\partial \alpha_{U,j,t}^{CBDL}} = \Phi_{U,\phi}^{F} \frac{\frac{1-s}{s} RER_t AS_{U,t}}{N_{R,t}}$

where

$$\mathcal{R}_{U,j,t} = \delta_{U,j,t}\left[(1-\alpha^{CBDL}_{U,j,t}) + \Gamma^{CBDL}_{U,t}\alpha^{CBDL}_{U,j,t}\right], \tag{C.42}$$

$$\mathcal{P}_{U,j,t} = \mathbf{E}_t\Omega_{U,j,t,t+1}\left[(1-\alpha^{CBDL}_{U,j,t})R^K_{U,t+1} + \alpha^{CBDL}_{U,j,t}R^{CBDL}_{U,t} - R_{U,t}\right], \tag{C.43}$$

Intermediate good firms

In each economy there exists a continuum of perfectly competitive intermediate goods firms that sell their output to domestic retailers. We assume that at the end of period t but before the realization of shocks the intermediate good firm acquires capital for use in next period's production. To do so, the intermediate good firm i claims equal to the number of units of capital acquired, and prices each claim at the real price of a unit of capital $Q_{R,t}$. The production function is

$$Z_{R,i,t} = \left(U_{R,i,t}K_{R,i,t-1}\right)^{\alpha} L^{(1-\alpha)}_{R,i,t}, \tag{C.44}$$

with $Z_{R,i,t}$ the amount of output produced by the individual RoW intermediate good firm in period t, $L_{R,i,t}$ the labor used in production, and $U_{R,i,t}$ the employed utilization rate of capital.

Cost minimization yields the standard equations for the optimal amount of production inputs

$$MC^r_{R,t} = \frac{w^{1-\alpha}_{R,t}\tau_{R,t}(U_{R,t})'^{\alpha}}{(1-\alpha)^{(1-\alpha)}\alpha^{\alpha}}. \tag{C.45}$$

$$\frac{w_{R,t}}{\tau_{R,t}(U_{R,t})'} = \frac{1-\alpha}{\alpha}\frac{(U_{R,t}K_{R,t-1})}{L_{R,t}}, \tag{C.46}$$

where $MC^r_{R,t}$ denote the real marginal costs of the intermediate good firms deflated by the RoW final good price $P^C_{R,t}$ and $\tau_{R,t}(U_{R,t})'$ as the derivative of the adjustment cost function, which maps a change in utilization rate into a change in the depreciation rate[101]. The optimal choice of capital gives the resulting gross nominal returns on capital, which are transferred to the bank in exchange for funding

$$R_{K,E,t} = (1+\pi^c_{R,t})\frac{\left(MC^r_{R,t}\alpha\frac{Z_{R,t}}{K_{t-1}}\right) + (Q_{R,t} - \tau_{R,t}U_{R,t})}{Q_{R,t-1}}. \tag{C.47}$$

C.4.5 Capital producers

Capital producing firms buy and refurbish depreciated capital from the intermediate goods firm at price $P^C_{R,t}$ and also produce new capital using the RoW final good, which consists of domestically produced and imported retail goods, as an input. Furthermore we assume

[101]The adjustment cost function is given by $\tau_{R,t}(U_{R,t}) = \tau_{R,ss,scale} + \zeta_{R,1}\frac{U_t^{1+\zeta_2}}{1+\zeta_2}$ with $\tau_{R,ss,scale}$ as an exogenous scale parameter in order to normalize utilization in the steady state.

that they face quadratic adjustment costs on net investment[102] and that profits, which arise outside of the steady state, are distributed lump sum to the households. The optimal choice of investment yields the familiar *Tobins Q* relation for the evolution of the relative price of capital

$$\begin{aligned} Q_{R,t} = 1 &+ \frac{\Psi}{2}\Big(\frac{In_{R,t}+Iss_R}{In_{R,t-1}+Iss_R}-1\Big)^2 + \Psi\Big(\frac{In_{R,t}+Iss_R}{In_{R,t-1}+Iss_R}-1\Big)\frac{In_{R,t}+Iss_R}{In_{R,t-1}+Iss_R} \\ &- \beta\frac{\Lambda_{E_{t+1}}}{\Lambda_{E_t}}\Psi\Big(\frac{In_{R,t+1}+Iss_R}{In_{R,t}+Iss_R}-1\Big)\Big(\frac{In_{R,t+1}+Iss_R}{In_{R,t}+Iss_R}\Big)^2 \end{aligned} \tag{C.48}$$

alongside the law of motion for capital

$$K_{R,t} = K_{R,t-1} + In_{R,t} \tag{C.49}$$

C.4.6 Goods bundling and pricing

The third key element in our model is dollar dominance in terms of DCP in bilateral trade between the US and the RoW, following the seminal work of Gopinath et al. (2020b). This means that the prices of both US and RoW exports are sticky in dollar.

In our model we go beyond DCP in bilateral trade between the US and the RoW and assume that prices of a share of domestic sales in the RoW are also sticky in dollar. In particular, Boz et al. (2022) document that a large share of trade among countries in the RoW is also priced in dollar; this is the actual meaning of a dominant—in the context of trade also often termed 'vehicle'—currency. It implies that when the dollar appreciates expenditure switching does not only affect imports from the US, but imports in general. Therefore, dollar pricing in third-country trade—in our model captured by domestic sales in the RoW—may be consequential for the effects of dollar appreciation in the context of a global risk aversion shock. To incorporate dollar pricing of a share of domestic sales in the RoW, we consider a multi-layered production structure in the spirit of Georgiadis and Schumann (2021) and depicted in Figure C.26

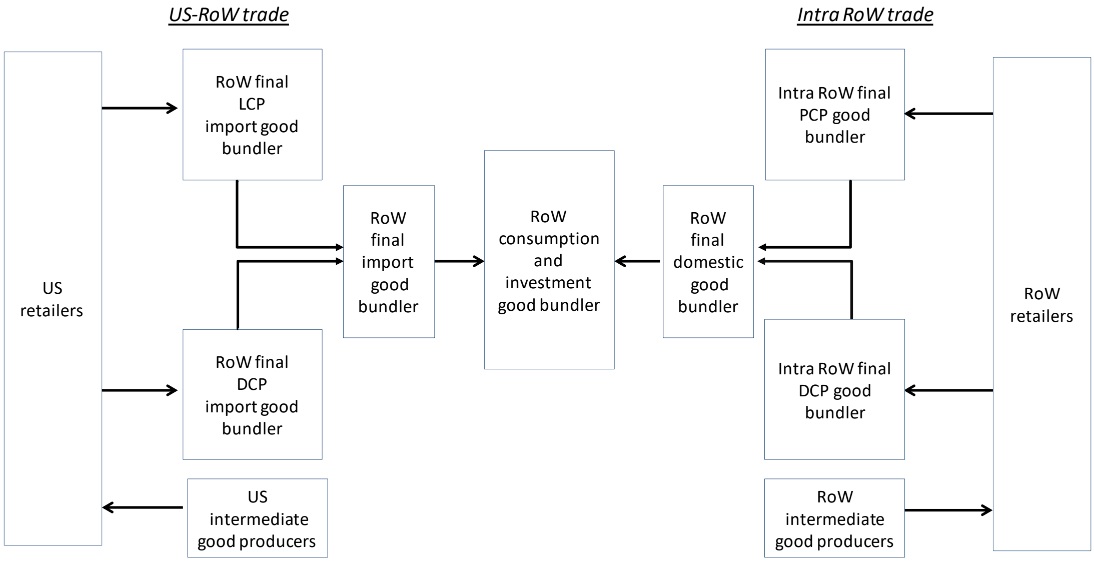
Import good bundling

As shown on the left side in Figure C.26, the RoW import good $Y^R_{U,t}$ is produced analogously to the RoW final domestic good $Y^R_{R,t}$. [103] In particular, RoW final import good producers use inputs from two branches of firms that operate under perfect competition and aggregate goods from US retail goods producers. The latter operate under monopolistic competition and set prices that are either sticky in the producer's currency (PCP goods) or in the

[102] Following Gertler and Karadi (2011) we assume that adjustment costs are only present when changing net investment in order for the optimal choice of the utilization rate to be independent from fluctuations in the relative price of capital $Q_{R,t}$

[103] Notice that the subscript indicates the country where the good is produced and the superscript the country where it is consumed.

Figure C.26: Multi-layered production structure for the RoW consumption and investment good



Note: The figure lays out the multi-layered production structure in the structural model, focusing on the RoW consumption and investment good.

importer's currency (LCP goods). Likewise, we assume that when exporting a fraction $(1-\gamma_U^R)$ of RoW and $(1-\gamma_R^U)$ of US retailers faces prices that are sticky in the currency of the importer, while the prices of the remaining firms are sticky in the producer's currency.

Table C.2 provides an overview of the core equations and first order conditions for the multistage bundling process of the final import good in the US. Equations are analogues for the RoW import good bundling process.

Final consumption and investment good

This sector operates at the top layer of this producttion structure and is populated by a continuum of firms that operate under perfect competition and combine a final domestically produced good $Y_{R,t}^R$ and a final import good $Y_{U,t}^R$ into a combined final good, employing the following CES technology

$$Y_{R,t}^C = \left[n_R^{\frac{1}{\psi_f}} Y_{R,t}^{R\,\frac{\psi_f - 1}{\psi_f}} + (1-n_R)^{\frac{1}{\psi_f}} Y_{U,t}^{R\,\frac{\psi_f - 1}{\psi_f}} \right]^{\frac{\psi_f}{\psi_f - 1}}. \tag{C.50}$$

The parameter n_R governs the share of domestically produced goods and thereby the degree of home bias in the assembling process[104]. The parameter ψ_f on the other hand corresponds

[104]The home bias parameter is adjusted in order to take into account the differences in country size as in Sutherland (2005). In particular, given a degree of general trade openness op_R and the relative country size of the RoW s, the parameter n_R takes the value $n_R = 1 - op_R(1-s)$ with a similar adjustment for the US

Table C.2: US import good bundling

Production function/Price index	Demand functions
US final import goods	
$Y^U_{R,t} = \left[\gamma_U^{R\,\frac{1}{\psi_i}} \tilde{Y}_{R,t}^{U\,\frac{\psi_i-1}{\psi_i}} + (1-\gamma_U)^{R\,\frac{1}{\psi_i}} \hat{Y}_{R,t}^{U\,\frac{\psi_i-1}{\psi_i}}\right]^{\frac{\psi_i}{\psi_i-1}}$ $P^{R^I}_{U,t} = \left[\gamma_F^E \left(\frac{\tilde{P}^F_{E,t}}{\mathcal{E}^F_{E,t}}\right)^{1-\psi_i} + (1-\gamma_F^E)\hat{P}^{F\,1-\psi_i}_{E,t}\right]^{\frac{1}{1-\psi_i}}$	$\tilde{Y}^U_{R,t} = \gamma_U^R \left(\frac{\tilde{P}^R_{U,t}}{\mathcal{E}_t P^{R^I}_{U,t}}\right)^{-\psi_i} Y^U_{R,t}$ $\hat{Y}^U_{R,t} = (1-\gamma_U^R)\left(\frac{\hat{P}^U_{R,t}}{P^{R^I}_{U,t}}\right)^{-\psi_i} Y^U_{R,t}.$
US imported PCP good	
$\tilde{Y}^U_{R,t} = \left[\left(\frac{1}{\gamma_F^E}\right)^{\frac{1}{\psi_i}}\left(\int_0^{\gamma_U^R} \tilde{Y}^U_{R,t}(i)^{\frac{\psi_i-1}{\psi_i}} di\right)\right]^{\frac{\psi_i}{\psi_i-1}}$ $\frac{\tilde{P}^U_{R,t}}{\mathcal{E}_t} = \left[\frac{1}{\gamma_U^R}\int_0^{\gamma_U^R}(\frac{\tilde{P}^U_{R,t}(i)}{\mathcal{E}_t})^{1-\psi_i}di\right]^{\frac{1}{1-\psi_i}}$	$\tilde{Y}^U_{R,t}(i) = \frac{1}{\gamma_U^R}\left(\frac{\tilde{P}^U_{R,t}(i)}{\tilde{P}^U_{U,t}}\right)^{-\psi_i}\tilde{Y}^U_{R,t}$ $= \left(\frac{\tilde{P}^U_{R,t}(i)}{\mathcal{E}_t P^{R^I}_{U,t}}\right)^{-\psi_i} Y^U_{R,t}$
US imported DCP good	
$\hat{Y}^U_{R,t} = \left[\left(\frac{1}{1-\gamma_U^R}\right)^{\frac{1}{\psi_i}}\left(\int_{\gamma_U^R}^1 \hat{Y}^U_{R,t}(i)^{\frac{\psi_i-1}{\psi_i}}di\right)\right]^{\frac{\psi_i}{\psi_i-1}}$ $\hat{P}^U_{R,t} = \left[\frac{1}{(1-\gamma_U^R)}\int_{\gamma_F^E}^1 \hat{P}^U_{R,t}(i))^{1-\psi_i}di\right]^{\frac{1}{1-\psi_i}}$	$\hat{Y}^U_{R,t}(i) = \frac{1}{1-\gamma_U^R}\left(\frac{\hat{P}^U_{R,t}(i)}{\hat{P}^U_{R,t}}\right)^{-\psi_i}\hat{Y}^U_{R,t}$ $= \left(\frac{\hat{P}^U_{R,t}(i)}{P^{R^I}_{U,t}}\right)^{-\psi_i}Y^U_{R,t}$

to the elasticity of substitution between the final domestic and import good.

Taking the prices of the domestic final good $P^R_{R,t}$ and the price of the final import good expressed in domestic currency $(\mathcal{E}_t P^R_{U,t})$[105] as well as total demand from consumers and capital producers as given, the optimal demand for goods produced domestically and abroad is governed by

$$Y^R_{R,t} = n_R\left(\frac{P^R_{R,t}}{P^C_{R,t}}\right)^{-\psi_f} Y^C_{R,t} \tag{C.51}$$

$$Y^R_{U,t} = (1-n_R)\left(\frac{\mathcal{E}_t P^R_{U,t}}{P^C_{R,t}}\right)^{-\psi_f} Y^C_{R,t}. \tag{C.52}$$

Lastly note that the three equations above imply that the price of the final consumption and investment good in the RoW $P^C_{E,t}$ is (up to first order) a weighted average of the prices of the final domestic and import good

$$P^c_{R,t} = \left[n_E P^{E^{1-\psi_f}}_{R,t} + (1-n_R)(\mathcal{E}P^R_{U,t})^{1-\psi_f}\right]^{\frac{1}{1-\psi_f}}. \tag{C.53}$$

RoW domestically produced and sold final good

We assume markets are partly segmented and firms set different prices in different markets depending on demand conditions. We assume a fraction of RoW firms $1-\gamma^R_R$ sets their prices for domestic sales in dollar, while the remaining prices are sticky in RoW currency. As in Gopinath et al. (2020b), we assume firms cannot choose their pricing currency, but

counterpart

[105]Note that because of the pricing-to-market assumption the price for US exports expressed in US-\$ $P^R_{U,t}$ will in general be different from the price charged for US goods sold in the US $P^U_{U,t}$.

Table C.3: RoW domestic sales bundling

Production function/Price index	Demand functions
RoW domestically produced final good	
$Y_{R,t}^{R} = \left[{\gamma_R^R}^{\frac{1}{\psi_i}} {\tilde{Y}_{R,t}^{R}}^{\frac{\psi_i - 1}{\psi_i}} + (1-\gamma_R)^{R^{\frac{1}{\psi_i}}} {\hat{Y}_{R,t}^{R}}^{\frac{\psi_i - 1}{\psi_i}} \right]^{\frac{\psi_i}{\psi_i - 1}}$ $P_{R,t}^{R} = \left[\gamma_R^R {\tilde{P}_{R,t}^{R}}^{1-\psi_i} + (1-\gamma_R^R) \left(\mathcal{E}_t \hat{P}_{R,t}^{R} \right)^{1-\psi_i} \right]^{\frac{1}{1-\psi_i}}$	$\tilde{Y}_{R,t}^{R} = \gamma_R^R \left(\frac{\tilde{P}_{R,t}^{R}}{P_{R,t}^{R}} \right)^{-\psi_i} Y_{R,t}^{R}$ $\hat{Y}_{R,t}^{R} = (1-\gamma_R^R) \left(\frac{\mathcal{E}_t \hat{P}_{R,t}^{R}}{P_{R,t}^{R}} \right)^{-\psi_i} Y_{R,t}^{R}$
RoW domestically sold PCP good	
$\tilde{Y}_{R,t}^{R} = \left[\left(\frac{1}{\gamma_R^R} \right)^{\frac{1}{\psi_i}} \int_0^{\gamma_E^R} \tilde{Y}_{R,t}^{R}(i)^{\frac{\psi_i - 1}{\psi_i}} di \right]^{\frac{\psi_i}{\psi_i - 1}}$ $\tilde{P}_{R,t}^{R} = \left[\frac{1}{\gamma_R^R} \int_0^{\gamma_R^R} \tilde{P}_{R,t}^{R}(i)^{1-\psi_i} di \right]^{\frac{1}{1-\psi_i}}$	$\tilde{Y}_{R,t}^{R}(i) = \frac{1}{\gamma_R^R} \left(\frac{\tilde{P}_{R,t}^{R}(i)}{\tilde{P}_{R,t}^{R}} \right)^{-\psi_i} \tilde{Y}_{R,t}^{R}$ $= \left(\frac{\tilde{P}_{R,t}^{R}(i)}{P_{R,t}^{R}} \right)^{-\psi_i} Y_{R,t}^{R}$
RoW domestically sold DCP good	
$\hat{Y}_{R,t}^{R} = \left[\left(\frac{1}{1-\gamma_R^R} \right)^{\frac{1}{\psi_i}} \left(\int_{\gamma_R^R}^{1} \hat{Y}_{R,t}^{R}(i)^{\frac{\psi_i - 1}{\psi_i}} di \right) \right]^{\frac{\psi_i}{\psi_i - 1}}$ $\mathcal{E}_t \hat{P}_{R,t}^{R} = \left[\frac{1}{(1-\gamma_R^R)} \int_{\gamma_R^R}^{1} (\mathcal{E}_t \hat{P}_{R,t}^{R}(i))^{1-\psi_i} di \right]^{\frac{1}{1-\psi_i}}$	$\hat{Y}_{R,t}^{R}(i) = \frac{1}{1-\gamma_R^R} \left(\frac{\mathcal{E}_t \hat{P}_{R,t}^{R}(i)}{\mathcal{E}_t \hat{P}_{R,t}^{R}} \right)^{-\psi_i} \hat{Y}_{R,t}^{R}$ $= \left(\frac{\mathcal{E}_t \hat{P}_{R,t}^{R}(i)}{P_{R,t}^{R}} \right)^{-\psi_i} Y_{R,t}^{R}$

are assigned to it exogenously and do not change it over time.

The firms that put together the RoW final domestic good $Y_{R,t}^{R}$ shown on the right side in Figure C.26 operate under perfect competition and combine inputs $\widetilde{Y}_{R,t}^{R}$ and $\widehat{Y}_{R,t}^{R}$ using a CES technology. The inputs are produced by two branches of firms that also operate under perfect competition and combine RoW retail goods. The firms in the first branch combine RoW retail goods $\widehat{Y}_{R,t}^{R}(i)$ priced in dollar (DCP goods) into the RoW final DCP good $\widehat{Y}_{R,t}^{R}$; analogously, the firms in the second branch combine RoW retail goods $\widetilde{Y}_{R,t}^{R}(i)$ priced in the producer's currency (PCP goods) into the RoW final PCP good $\widetilde{Y}_{R,t}^{R}$.

The next layer consists of RoW retail-goods-producing firms which buy and repackage RoW intermediate goods. These firms operate under monopolistic competition and serve the RoW as well as the US market; for simplicity Figure C.26 only shows their domestic sales. The share of RoW retail-goods-producing firms whose domestic sales prices are sticky in dollar is given by $(1-\gamma_R^R)$. Therefore, $(1-\gamma_R^R)$ also reflects the degree to which movements in the dollar exchange rate cause fluctuations in the RoW aggregate producer-price index $P_{R,t}^{R}$.

Table C.3 provides an overview of the core equations and first order conditions for the multistage bundling process.

C.4.7 Retail good pricing

There exists a continuum of firms that operate under monopolistic competition and use intermediate goods to produce a retail good that is eventually sold to the specialized branches farther up. Each retail firm sells its product in the domestic and foreign markets; as mentioned above, for simplicity we only show sales to RoW in Figure C.26. When selling in the RoW (i.e. domestic) market, a fraction γ_R^R of RoW retail-goods-producing firms sets prices in RoW currency, while the remaining $(1-\gamma_R^R)$ share of firms sets their prices in dollar. A similar setting exists in the market for US imports, with γ_U^R indicating the fraction of RoW firms that price their exports in the producer's currency. Regardless of the pricing currency, all firms use the same production technology and face the same factor costs. Because firms are subject to Calvo-style price-setting frictions and can only change their price with a probability $(1-\theta_p^R)$ each period, the mark-up of a firm whose price is sticky in dollar fluctuates with the exchange rate. As RoW firms serving domestic and US markets, respectively, set their prices optimally and as in each market they use different pricing currencies, their profit functions differ as shown in table C.4. As standard in Calvo-style price setting, firms choose their optimal reset price given demand and their pricing currency while taking into account that they might not be able to reset their price in the future. For instance the optimal price choice of a DCP firm i for its sales in the RoW market, taking into account the fact that it may not be able to reset its US-\$ denominated price $\hat{P}_{R,t}^R(i)$, can be written as

$$\underset{\hat{P}_{R,t}^R(i)}{max}\mathbb{E}_t\sum_{s=0}^{\infty}\theta_p^{R^s}\Theta_{R_{t,t+s}}\Big[\mathcal{E}\hat{P}_{R,t}^R(i)Y_{R,t}^R(i)-MC_{R,t}Y_{R,t}^R(i)\Big]. \tag{C.54}$$

It is possible to show that the optimal reset price of a firm that sets its price for the RoW market in US-\$, relative to the aggregate RoW DCP sales price index $\hat{P}_{R,t}^R$, is given by

$$\frac{\hat{P}_{R,t}^R(i)}{\hat{P}_{R,t}^R}=\hat{p}_{R,t}^R=\frac{\psi_i}{(\psi_i-1)}\frac{\hat{x}_{R,1,t}^R}{\hat{x}_{R,2,t}^R}. \tag{C.55}$$

The auxiliary recursive variables $\hat{x}_{R,1,t}^R$ and $\hat{x}_{R,2,t}^R$ read as

$$\hat{x}_{R,1,t}^R=\Lambda_{R,t}\Big(\frac{\mathcal{E}\hat{P}_{R,t}^R}{P_{R,t}^R}\Big)^{-\psi_i}Y_{R,t}^R\frac{P_{R,t}^R}{P_{R,t}^C}MC_{R,t}^{rp}+\beta\theta_p\mathbb{E}_t\hat{x}_{R,1,t+1}^R(1+\hat{\pi}_{R,t+1}^R)^{\psi_i} \tag{C.56}$$

$$\hat{x}_{R,2,t}^R=\Lambda_{E,t}\Big(\frac{\mathcal{E}\hat{P}_{R,t}^R}{P_{R,t}^E}\Big)^{-\psi_i}Y_{R,t}^E\Big(\frac{\mathcal{E}\hat{P}_{R,t}^E}{P_{R,t}^C}\Big)+\beta\theta_p^R\mathbb{E}_t\hat{x}_{R,1,t+1}^R(1+\hat{\pi}_{R,t+1}^R)^{\psi_i-1}, \tag{C.57}$$

with $MC_{R,t}^{rp}$ as marginal costs deflated in by the aggregate producer price $P_{R,t}^R$. It becomes apparent that not only does the exchange rate $\mathcal{E}$ impact the optimal DCP price setting decision as it determines the demand for DCP goods via the relative price $\frac{\mathcal{E}\hat{P}_{R,t}^R}{P_{R,t}^R}$, it also impacts the optimal reset price via the term $\frac{\mathcal{E}\hat{P}_{R,t}^R}{P_{R,t}^C}$, which translates the local currency revenues that a DCP firm makes from selling one unit of its good $\mathcal{E}\hat{P}_{R,t}^R$ into the unit of account that the firm's owners (households) care about $P_{R,t}^C$. Everything else equal, an appreciation of the US-\$ exchange rate, will cause the local currency revenues per unit of DCP good sold to rise, while the input costs, which are denominated in the RoW currency, remain roughly stable. Thus the mark-up rises above the optimal mark-up and a DCP good firm would like to lower its US-\$ price in response to an appreciation of the US-\$ over and

Table C.4: Market and pricing paradigm specific profit functions of RoW firms

Type of firm and market	Profit function
RoW market PCP firm	$\tilde{\Pi}^R_{R,t}(i) = \tilde{P}^R_{R,t}(i)\tilde{Y}^R_{R,t}(i) - MC_{R,t}\tilde{Y}^E_{R,t}(i)$
RoW market DCP firm	$\hat{\Pi}^R_{R,t}(i) = \mathcal{E}\hat{P}^R_{R,t}(i)\hat{Y}^R_{R,t}(i) - MC_{R,t}\hat{Y}^R_{R,t}(i)$
US import market PCP firm	$\tilde{\Pi}^U_{R,t}(i) = \tilde{P}^U_{R,t}(i)\tilde{Y}^U_{R,t}(i) - MC_{R,t}\tilde{Y}^U_{R,t}(i)$
US import market DCP firm	$\hat{\Pi}^U_{R,t}(i) = \mathcal{E}\hat{P}^U_{R,t}(i)\hat{Y}^U_{R,t}(i) - MC_{R,t}\hat{Y}^U_{R,t}(i)$

above what the induced fall in RoW demand for the DCP good would dictate. It is easy to verify that when aggregating across intra RoW sales of RoW DCP firms the inflation rate of the aggregate RoW sales DCP price (expressed in US-$) is given by

$$1 = (1-\theta_p)\hat{p}^{R^{1-\psi_i}}_{R,t} + \theta_p(1+\hat{\pi}^R_{R,t})^{(\psi_i-1)}, \tag{C.58}$$

where $\hat{p}^R_{R,t}$ denotes the ratio of the optimal reset price relative to the aggregate price index. Using the profit functions in table C.4 its possible to show similar equations hold for the optimal price of RoW retail firms that set their prices in the US import market in US-$ as well as, with slight adaptions, for PCP firms.

C.4.8 Market clearing and the aggregate budget constraint

Turning to the market clearing conditions, aggregate demand for the domestic consumption good $Y^C_{E,t}$ is given by the sum of individual demand from all sources that either consume the good or use it as an input in production

$$Y^C_{R,t} = C_{R,t} + I_{R,t} + \frac{\Psi}{2}\Big(\frac{In_{R,t}+Iss_R}{In_{R,t-1}+Iss_R} - 1\Big)^2(In_{R,t}+Iss_R). \tag{C.59}$$

Aggregating across all intermediate and retail goods firms and imposing market clearing yields the aggregate production function of the economy

$$Z_{R,t} = (U_{R,t}K_{R,t-1})^\alpha L^{(1-\alpha)}_{R,t} = \delta^R_{R,t}Y^R_{R,t} + \delta^F_{R,t}Y^F_{R,t}, \tag{C.60}$$

with $\delta^R_{R,t}$ and $\delta^F_{R,t}$ as price dispersion terms which are zero up to a first order approximation. $Y^R_{R,t}$ corresponds to the aggregate domestic demand for the final *domestically produced* RoW good given by

$$Y^R_{R,t} = n_R\Big(\frac{P^R_{R,t}}{P^C_{R,t}}\Big)^{-\psi_f}Y^C_{R,t}, \tag{C.61}$$

with $Y^C_{R,t}$ as the households and firms demand for the final good. Furthermore the aggregate demand for RoW goods produced for exports reads as

$$Y^F_{R,t} = \frac{1-s}{s}(1-n_F)\left(\frac{\mathcal{E}_t P^F_{R,t}}{P^C_{F,t}}\right)^{-\psi_f} Y^C_{F,t}, \tag{C.62}$$

where it it is important to note that variables are expressed in per capita terms and therefore, following Sutherland (2005), the relative population size has to be taken when aggregating across countries as indicated by the ratio $\frac{1-s}{s}$.

We assume financial markets clear, which implies $GB_{U,t} = \frac{s}{1-s} GB_{R,t}$ and $CBDL_{U,t} = \frac{s}{1-s} CBDL_{R,t}$, where s is the relative country size parameter. When aggregating across budget constraints in the RoW, we recover the national accounting identity

$$\begin{aligned} RER_t\Big[\Big(GB_{R,t} - \frac{R^{GB}_{U,t-1}}{1+\pi^C_{U,t}} GB_{R,t-1}\Big) - \Big(CBDL_{R,t} - \frac{R^{CBDL}_{U,t-1}}{1+\pi^C_{U,t}} CBDL_{R,t-1}\Big)\Big] = \\ \frac{P^R_{R,t}}{P^C_{R,t}} Y^R_{R,t} + \frac{\mathcal{E}_t P^F_{R,t}}{P^C_{R,t}} Y^U_{R,t} - Y^C_{R,t}. \end{aligned} \tag{C.63}$$

The left-hand side represents the sum of the changes in the RoW net foreign asset position and the net financial account, while the right-hand side is the trade balance (taking into account that prices charged differ across domestically produced and exported goods). Importantly, and in contrast to Akinci and Queralto (2019), Devereux et al. (2020) and many others, we explicitly model *gross* rather than only net financial flows. As a consequence, the national accounting identity does not dictate the evolution of all financial flows as in a net-flows model. In a net-flow model, where, for instance, RoW banks can only borrow in dollars but not hold dollar assets (i.e. gross liabilities equal net liabilities), the trade balance and costs of funds borrowed in the previous period determine uniquely the foreign banking sector's liability position in the next period. In contrast, in our model the national accounting identity only uniquely determines the *sum of the changes* in gross assets and liabilities has to equal the sum of the trade balance and the financial account.

C.4.9 Calibration

We generally allow parameter values to differ across the US and the RoW (see Table C.5). For parameters that govern standard model elements, to the extent possible we draw on estimates from existing literature. In particular, for US parameters we rely on Justiniano et al. (2010). For the RoW it is more difficult to find suitable estimates, as it reflects an aggregate of countries. Since the euro area accounts for roughly one quarter of the RoW in the data in terms of output, we use the estimates in Coenen et al. (2018) for many of the RoW parameters. We next discuss the calibration of the parameters that govern DCP in trade and cross-border credit.

Regarding DCP in trade we first calibrate the relative country size s such that the steady-state share of US real GDP in global output is 25%. Given the country sizes, we set the general RoW openness vis-à-vis the US (op_R) such that the steady-state share of imports from the US in the aggregate RoW bundle ($1-\eta_R$) is roughly 5.1%, in line with the data over 1990-2019. In the same vein, we set US trade openness (op_U) such that the share of imports in the US bundle ($1-\eta_U$) is roughly 14%. We set the share of RoW firms that face sticky dollar prices when exporting to the US ($1-\gamma_U^R$) to 93%, in line with invoicing shares documented in Gopinath (2015). Based on the calculations in Georgiadis and Schumann (2021) we assume that US exporters almost exclusively face sticky prices in dollar and set γ_R^U to 3%. We set the share of intra-RoW *sales* that is priced in dollar ($1-\gamma_R^R$) to 9%, which implies that 37.5% of intra-RoW *exports* are priced in dollar as indicated by the invoicing data in Boz et al. (2022).[106] We almost exclusively choose the parameters that govern the endogenous portfolio choices of RoW and US banks in order to meet some steady-state targets. For both the US and the RoW banking sectors we follow Akinci and Queralto (2019) and assume a (risk weight adjusted) steady-state leverage ratio of five. Furthermore, we impose that the steady-state domestic credit spread ($R_i^K - R_i$) equals 200 basis points, which roughly corresponds to the average of the GZ-spread of Gilchrist and Zakrajsek (2012). These two assumptions imply the country specific values for the start-up fund parameter (ω_B) and the constants in balance-sheet-specific risk weights ($\bar{\delta}$) shown in Table C.5. We assume an average bank planning horizon of 7.5 years, which lies in between the 10 year of Gertler and Karadi (2011) and the one in Akinci and Queralto (2019). This implies that we set $\theta_{U,B} = \theta_{R,B}$ of 0.9667. For the parameters governing the portfolio choice of US banks we target a risk premium that is a fifth of the US domestic credit spread (a conservative choice) as well as an annualized steady-state 'exorbitant privilege' (Gourinchas and Rey 2007) of 1%, which pins down $\bar{\Gamma}_U^{CBDL}$ and $\Phi_{U,\phi}$. For RoW banks we jointly determine the parameters $\epsilon_{R,\alpha}$, $\bar{\delta}_R$ and $\kappa_{R,\alpha,\ell}$ in order to hit three steady state targets: A leverage ratio of five and a portfolio such that RoW banks invest 15% of their total liabilities in US Treasuries and finance 25% of their total assets using cross-border dollar loans. The latter roughly corresponds to the average liability structure of non-US, internationally active banks in the BIS Locational Banking Statistics.[107]

Finally, we impose that the US and RoW steady-state risk-free rates are 2% and 3.5%,

[106]We first calculate the fraction of intra-RoW trade (global exports without US imports and exports) over global non-US GDP and then take the yearly average from 1990-2019 ($\approx 24\%$). Next, we use the average share of global exports invoiced in dollar as calculated in Boz et al. (2022) and subtract the fraction of US trade in global trade to arrive at 37.5%. Multiplying the two numbers we arrive at about 9%.

[107]Combined with the assumption that banks are the only entities engaging in global financial markets our model calibration implies that the RoW has a negative net dollar exposure and is a net debtor to the US ($\alpha_R^{TREAS} - \ell_R^{CBDL} < 0$). While this is in line with the negative net dollar exposures of the RoW banking sector documented in Shin (2012), the entire RoW economy has a positive net dollar exposure vis-à-vis and is a net creditor to the US. This lies at the heart of the 'exorbitant duty' (Gourinchas et al. 2012; Gourinchas and Rey 2022). In Georgiadis et al. (2023) we consider a simple extension in which we introduce an additional RoW entity whose asset holdings render the aggregate RoW economy a net creditor with a negative net dollar exposure. We show that when this entity is unconstrained—thus to be thought of as a central bank holding foreign exchange reserves, pension or sovereign wealth funds—the exorbitant duty is an exchange rate valuation effect without real implications.

respectively. These values roughly correspond to the averages in the data and pin down the discount factors β_U and β_R. These assumptions imply that the steady-state trade deficit to GDP ratio of the US is 1.8%, which is close to the average in the data. The US finances this trade deficit by a positive net financial income, which results from the US earning higher returns from cross-border dollar lending to the RoW than it pays for Treasuries held by the RoW. Therefore, the US maintains a higher steady-state per capita consumption than the RoW as a direct consequence of the exorbitant privilege.

Table C.5: Parameter values used in the simulations I

Param.	**Val.**	**Description**	**Source**
		Households	
h_R	0.620	Habit persistence in consumption RoW	CKSW(2018)[a]
h_U	0.790	Habit persistence in consumption US	JPT(2010)
σ_c	1.002	Intertemporal elasticity of substitution	$\approx$ log utility
φ	2.000	Inverse Frisch elasticity of labor	CKSW(2018)
β_U	0.995	Discount factor US	2% ann. US rate
β_R	0.9913	Discount factor ROW	3.5% ann. RoW rate
		RoW financial intermediaries	
ω_B^U	0.00036	Start up funds RoW	endogenous in SS
θ_B^U	0.9667	Survival probability of Banks RoW	AQ(2019)+GK(2011)
$\epsilon_{R,\alpha}$	0.5479	IC parameter for US GB	endogenous in SS
Γ_R^{GB}	0	Risk weight for US GB	endogenous in SS
$\kappa_{R,\alpha,\ell}$	2.7397	IC parameter unhedged US$ debt	endogenous in SS
$\bar{\delta}_{B,U}$	0.6790	Constant in incentive constraint (IC)	endogenous in SS
		US financial intermediaries	
ω_B^U	0.00026	Start-up funds parameter US	endogenous in SS
θ_B^U	0.966	Survival probability of Banks US	AQ(2019)+GK(2011)
$\bar{\delta}_{B,U}$	1.0468	Constant in incentive constraint (IC)	endogenous in SS
$\bar{\Gamma}_U^{CBDL}$	0.3	SS Risk weight of global interbank loans	endogenous in SS
$\Phi_{\Gamma,U}$	0.1012	Semi elasticity of Γ_U^{CBDL} wrt $\phi_{R,t}$	endogenous in SS
ρ_δ	0.95	Common persistence of global risk shock	VAR dynamics
		Wage decision	
ψ_w	6.000	Elasticity of substitution labor services	20% wage mark up
θ_w^R	0.780	Calvo parameter wages RoW	CKSW(2018)
θ_w^U	0.840	Calvo parameter wages US	JPT(2010)
		International trade	
ψ_f	1.120	Trade price elasticity	CKSW(2018)
op_R	0.200	General trade openness RoW	$\eta_R \approx 0.95$
op_U	0.185	General trade openness US	$\eta_U \approx 0.86$
n	0.750	Share of RoW in global economy	$1 - \frac{GDP_{US}}{GDP_{RoW}}$

[a] GK(2011), JPT(2010), CKSW(2018), GZ(2012), JKL(2021), AQ(2019), G(2015), represent abbreviations for Gertler and Karadi (2011), Justiniano et al. (2010), Coenen et al. (2018), Gilchrist and Zakrajsek (2012), Jiang et al. (2021) Akinci and Queralto (2019) and Gopinath (2015) respectively.

Table C.6: Parameter values used in the simulations II

Param.	**Val.**	**Description**	**Source**
		Intermediate goods production	
α	0.333	Share of capital in production	AQ(2019)
ζ_2	5.800	Elasticity of depreciation wrt. to utilization	JPT(2010) [a]
$\tau_{R,ss}$	0.020	Normalization parameter depreciation RoW	endogenous in SS
ζ_1^R	0.035	Normalization of utilization parameter RoW	endogenous in SS
ζ_1^U	0.035	Normalization of utilization parameter US	endogenous in SS
$\tau_{U,ss}$	0.020	Normalization parameter depreciation US	endogenous in SS
		Retail good pricing	
ψ_i	6.000	Elasticity of substitution retail goods	20% mark up
θ_P^R	0.820	Calvo parameter retail firms RoW	CKSW(2018)
θ_P^U	0.840	Calvo parameter retail firms US	JPT(2010)
$\widehat{\gamma_R^R} = 1 - \gamma_R^R$	0.09	Share of RoW domestic sales DCP firms	37.5% intra RoW exp.
$\gamma_{U^R} \widehat{= 1} - \gamma_U^R$	0.97	Share of RoW export to US DCP firms	$\approx$ G(2015) invoicing
$\widetilde{\gamma_R^U} = 1 - \gamma_R^U$	0.05	Share of US export LCP firms	$\approx$ G(2015) invoicing
		Capital goods production	
Ψ_R	5.770	Investment adjustment costs RoW	CKSW(2018)
Ψ_U	2.950	Investment adjustment costs US	JPT(2010)
		Monetary Policy	
$\rho_{U,r}$	0.930	RoW interest rate smoothing	CKSW(2018)
$\phi_{U,\pi}$	2.740	RoW Taylor Rule coefficient inflation	CKSW(2018)
$\phi_{U,z}$	0.030	RoW Taylor Rule coefficient output	CKSW(2018)
$\rho_{R,r}$	0.810	US interest rate smoothing	JPT (2010)
$\phi_{R,\pi}$	1.970	US Taylor Rule coefficient inflation	JPT(2010)
$\phi_{R,z}$	0.050	US Taylor Rule coefficient output	JPT(2010)

[a] GK(2011), JPT(2010), CKSW(2018), GZ(2012), JKL(2021), AQ(2019), G(2015), represent abbreviations for Gertler and Karadi (2011), Justiniano et al. (2010), Coenen et al. (2018), Gilchrist and Zakrajsek (2012), Jiang et al. (2021) Akinci and Queralto (2019) and Gopinath (2015) respectively.

Table C.7: Steady State Targets

Param.	**Val.**	**Description**	**Source**
Steady State targets			
$L_{R,ss}$	0.333	SS labor target RoW	GK(2011) [a]
U_{ss}	1.000	SS utilization rate target RoW and US	JPT(2010)
τ_{ss}	0.025	SS depreciation rate target RoW and US	JPT(2010)
$S_{R,ss}$	0.005	SS credit spread target RoW (quarterly)	≈ CKSW(2018)
$S_{U,ss}$	0.005	SS credit spread target US (quarterly)	≈ avg. GZ spread
$\phi_{R,ss}$	5.00	SS (risk weighted) leverage target, RoW	CKSW(2018)
$\phi^{F}_{U,ss}$	5.00	SS (risk weighted) local leverage target, US	GK(2011)
$\ell^{CBDL}_{R,j,t}$	0.25	SS dollar debt portfolio share RoW	≈ LBS avg.
$\alpha^{GB}_{R,j,t}$	0.15	SS US treasuries portfolio share RoW	≈ LBS avg.
$R^{CBDL}_{U,ss} - R^{GB}_{U,ss}$	0.0025	SS Exorbitant priviledge	1% annualized
$CY_{R,ss}$	0.015/4	SS convenience yield	≈ JKL(2012)
$RP^{CBDL}_{U,ss}$	0.001	SS interbank risk premium	1/5 of credit spread

[a] GK(2011), JPT(2010), CKSW(2018), GZ(2012), JKL(2021), AQ(2019), G(2015), represent abbreviations for Gertler and Karadi (2011), Justiniano et al. (2010), Coenen et al. (2018), Gilchrist and Zakrajsek (2012), Jiang et al. (2021) Akinci and Queralto (2019) and Gopinath (2015) respectively.

C.5 List of all model equations

This section contains all the relevant model equations of the Trinity model of Georgiadis et al. (2023) as they appear in the corresponding code.[108]

C.5.1 Households

Marginal Utility RoW

$$\Lambda_{Rt} = exp\left(\varepsilon^{\beta}_{Rt}\right)\,\left(C_{Rt} - h_R\,C_{Rt-1}\right)^{(-\sigma_c)} - \beta_R\,h_R\,\left(exp\left(\varepsilon^{\beta}_{Rt+1}\right)\,C_{Rt+1} - C_{Rt}\,h_R\right)^{(-\sigma_c)} \tag{C.64}$$

Euler equation RoW

$$\Lambda_{Rt} = \beta_R\,\left(1 + R_{Rt}\right)\,\frac{\Lambda_{Rt+1}}{1 + \pi^{C}_{Rt+1}} \tag{C.65}$$

Demand shock RoW

$$\varepsilon^{\beta}_{Rt} = \rho^{\beta}\,\varepsilon^{\beta}_{Rt-1} + \frac{\eta^{\beta}_{Rt}}{100} \tag{C.66}$$

Marginal Utility US

$$\Lambda_{Ut} = exp\left(\varepsilon^{\beta}_{Ut}\right)\,\left(C_{Ut} - h_U\,C_{Ut-1}\right)^{(-\sigma_c)} - \beta_U\,h_U\,\left(exp\left(\varepsilon^{\beta}_{Ut+1}\right)\,\left(C_{Ut+1} - C_{Ut}\,h_U\right)\right)^{(-\sigma_c)} \tag{C.67}$$

Euler equation US

$$\Lambda_{Ut} = \beta_U\,\left(1 + R_{Ut}\right)\,\frac{\Lambda_{Ut+1}}{1 + \pi^{C}_{Ut+1}} \tag{C.68}$$

Demand Shock US

$$\varepsilon^{\beta}_{Ut} = \rho^{\beta}\,\varepsilon^{\beta}_{Ut-1} + \frac{\eta^{\beta}_{Ut}}{100} \tag{C.69}$$

UIP deviation

$$\widehat{UIP}_t = \left(1 + R_{Ut}\right)\,\left(1 + D\mathcal{E}_{t+1}\right) - \left(1 + R_{Rt}\right) \tag{C.70}$$

C.5.2 RoW financial intermediaries

Discounted excess return to investing in domestic capital RoW

$$v_{Rt} = \Omega_{Rt+1}\,\left(R_{K,R_{t+1}} - \left(1 + R_{Rt}\right)\right) \tag{C.71}$$

Discounted return to equity RoW

$$n_{Rt} = \left(1 + R_{Rt}\right)\,\Omega_{Rt+1} \tag{C.72}$$

[108] The corresponding *DYNARE* file is available upon request

Aggregate Net worth RoW financial sector

$$N_{Rt} = N_{R,e_t} + N_{R,n_t} \tag{C.73}$$

RoW credit spread

$$S_{Rt} = R_{K,R_{t+1}} - (1 + R_{Rt}) \tag{C.74}$$

RoW capital price expressed in dollars

$$Q_{R,US\$_t} = \frac{Q_{Rt}}{RER_t} \tag{C.75}$$

Aggregate Assets RoW (taking into account that $\phi_{R,t}$ is the *risk adjusted* leverage ratio in the code)

$$AS_{Rt} = \frac{N_{Rt}\,\phi_{Rt}}{(1-\alpha_R^{GB})_t + \Gamma_R^{GB}\,\alpha_R^{GB}{}_t} \tag{C.76}$$

Net Worth of new banks RoW

$$N_{R,n_t} = \omega^R\ (AS_{R,t-1}) \tag{C.77}$$

Discounted excess costs of borrowing in Dollars

$$u_{Rt} = \Omega_{Rt+1}\ \left((1 + D\mathcal{E}_{t+1})\ R_{U,t}^{CBDL} - (1 + R_{Rt})\right) \tag{C.78}$$

RoW banks stochastic discount factor

$$\Omega_{Rt} = \beta_R\,\frac{\Lambda_{Rt}}{\Lambda_{Rt-1}}\,\frac{1}{1+\pi_{Rt}^C}\ \left(1 - \theta_B^R + \theta_B^R\ \left(n_{Rt} + \left(v_{Rt}\,(1-\alpha_R^{GB})_t + \alpha_R^{GB}{}_t\,v_R^{GB}{}_t - u_{Rt}\,\ell_{R,t}^{CBDL}\right)\ \frac{\phi_{Rt}}{(1-\alpha_R^{GB})_t + \Gamma_R^{GB}\,\alpha_R^{GB}{}_t}\right)\right) \tag{C.79}$$

FOC optimal liability choice RoW

$$-u_{Rt} = \frac{\delta'_{R,\ell_t}}{\delta_{R,B_t}}\ \left(v_{Rt}\,(1-\alpha_R^{GB})_t + \alpha_R^{GB}{}_t\ \left(v_R^{GB}{}_t + CV_{Rt}\right)\right) \tag{C.80}$$

Risk weight adjusted optimal leverage ratio RoW

$$\phi_{Rt} = \frac{n_{Rt}\ \left((1-\alpha_R^{GB})_t + \Gamma_R^{GB}\,\alpha_R^{GB}{}_t\right)}{u_{Rt}\,\ell_{R,t}^{CBDL} + (1-\alpha_R^{GB})_t\,\delta_{R,B_t} + \alpha_R^{GB}{}_t\,\Gamma_R^{GB}\,\delta_{R,B_t} - v_{Rt}\,(1-\alpha_R^{GB})_t - \alpha_R^{GB}{}_t\,v_R^{GB}{}_t} \tag{C.81}$$

Time varrying balance sheet specific risk weight RoW

$$\delta_{R,B_t} = \overline{\delta}_R\ \left(1 - \alpha_R^{GB}{}_t\,\epsilon_{R,\alpha} + \frac{\kappa_{R,\alpha,\ell_t}}{2}\ \left(\alpha_R^{GB}{}_t - \ell_{R,t}^{CBDL}\right)^2\ exp\left(\epsilon^{\delta^R}{}_t\right)\right) \tag{C.82}$$

Risk aversion shock RoW

$$\epsilon^{\delta^R}{}_t = \rho^\delta\,\epsilon^{\delta^R}{}_{t-1} + \sigma_\eta^{\delta_R}\,\eta_{Ut}^\delta + \sigma_\eta^{\delta_G}\,\eta_{Gt}^\delta \tag{C.83}$$

LOM aggregate equity of existing banks RoW banking sector

$$N_{R,e_t} = \frac{1}{1+\pi^C_{Rt}} \, \theta^R_B \, \Big(\big((R_{K,R_t} - (1+R_{Rt-1})\big) \, (1-\alpha^{GB}_R)_{t-1} + \big((1+D\mathcal{E}_t) \, R^{GB}_{R\;t-1} - (1+R_{Rt-1})\big) \, \alpha^{GB}_{R\;t-1} - \big((1+D\mathcal{E}_t) \, R^{CBDL}_{U,t-1} - (1+R_{Rt-1})\big) \, \ell^{CBDL}_{R,t-1} \, AS_{Rt-1} + (1+R_{Rt-1}) \, N_{Rt-1} \tag{C.84}$$

Definition of CBDL portfolio share

$$\ell^{CBDL}_{R,t} = \frac{RER_t \, CBDL_{R,t}}{AS_{Rt}} \tag{C.85}$$

Aggregate assets RoW banking sector

$$AS_{Rt} = Q_{Rt} \, K_{Rt} + RER_t \, GB_{Rt} \tag{C.86}$$

Definition of US treasury portfolio share RoW

$$\alpha^{GB}_{R\;t} = \frac{GB_{R,val_t}}{AS_{Rt}} \tag{C.87}$$

Definition of domestic investment portfolio share RoW (redundant)

$$(1-\alpha^{GB}_R)_t = \frac{Q_{Rt} \, K_{Rt}}{AS_{Rt}} \tag{C.88}$$

Total value of US treasuries held by RoW banks

$$GB_{R,val_t} = RER_t \, GB_{Rt} \tag{C.89}$$

Return on treasuries (in US-$)

$$R^{GB}_{R\;t} = 1 + R_{Ut} \tag{C.90}$$

Discounted excess returns (in RoW currency) from investing in US treasuries

$$v^{GB}_{R\;t} = \Omega_{Rt+1} \, \big((1+D\mathcal{E}_{t+1}) \, R^{GB}_{R\;t} - (1+R_{Rt})\big) \tag{C.91}$$

Derivative of time varying balance sheet specific risk weight wrt. CBDL share

$$\delta'_{R,\ell_t} = \bar{\delta}_R \, \big(\epsilon_{R,\ell} \, \big(\ell^{CBDL}_{R,t} - \bar{\ell}_R\big) + \kappa_{R,\alpha,\ell_t} \, \big(\ell^{CBDL}_{R,t} - \alpha^{GB}_{R\;t}\big)\big) \tag{C.92}$$

Derivative of time varying balance sheet specific risk weight wrt. treasury share

$$\delta'_{R,\alpha_t} = \bar{\delta}_R \, \big(\kappa_{R,\alpha,\ell_t} \, \big(\alpha^{GB}_{R\;t} - \ell^{CBDL}_{R,t}\big) - \epsilon_{R,\alpha}\big) \tag{C.93}$$

Convenience yield from investing in treasuries RoW banks

$$CV_{Rt} = v_{Rt} \, \big(-\big((1-\alpha^{GB}_R)_t + \Gamma^{GB}_R \, \alpha^{GB}_{R\;t}\big)\big) \, \frac{\delta'_{R,\alpha_t}}{\delta_{R,B_t}} \tag{C.94}$$

FOC asset choice

$$v^{GB}_{R\;t} = v_{Rt} \, \Gamma^{GB}_R - CV_{Rt} \tag{C.95}$$

C.5.3 US financial intermediaries

Discounted returns to investing domestically US

$$v_{U\,t} = \Omega_{U\,t+1}\,\Big(R_{K,U\,t+1} - (1 + R_{U\,t})\Big) \tag{C.96}$$

Discounted returns to equity US

$$n_{U\,t} = (1 + R_{U\,t})\;\Omega_{U\,t+1} \tag{C.97}$$

US balance sheet specific risk weight (constant up to shock)

$$\delta^{U}{}_{t} = \overline{\delta_U}\;exp\left(\epsilon^{\delta}_{U\,t}\right) \tag{C.98}$$

US risk aversion shock

$$\epsilon^{\delta}_{U\,t} = \sigma_{\eta}^{\delta_G}\;\eta^{\delta}_{G\,t} + \rho^{\delta}\;\epsilon^{\delta}_{U\,t-1} + \sigma_{\eta}^{\delta_U}\;\eta^{\delta}_{R\,t} \tag{C.99}$$

Aggregate equity US financial sector

$$N_{U\,t} = N_{U,e_t} + N_{U,n_t} \tag{C.100}$$

Credit spread US

$$S_{U\,t} = R_{K,U\,t+1} - (1 + R_{U\,t}) \tag{C.101}$$

Aggregate equity US financial sector

$$N_{U,n_t} = \omega^{U}\,AS_{U,t-1} \tag{C.102}$$

Time varying asset specific risk weight of cross border lending

$$\Gamma^{CBDL}_{U_t} = \Gamma^{CBDL}_{R,ss}\;exp\left(\epsilon_{\Gamma\,t}\right) + \Phi^{\Gamma}_{U}\;\left(\phi_{R\,t} - (\bar{\phi_R})\right) \tag{C.103}$$

Shock to asset specific risk weight of cross border dollar lending

$$\epsilon_{\Gamma\,t} = \rho_{\Gamma}\;\epsilon_{\Gamma\,t-1} + \sigma^{\Gamma}_{\eta}\;\eta^{\Gamma}_{U\,t} \tag{C.104}$$

Ratio of total CBDL to dometic lending (This is equivalent to $\alpha^{CBDL}_{U,t}/(1-\alpha^{CBDL}_{U,t})$)

$$\xi^{CBDL}_{U_t} = \frac{AS_{R\,t}\;\ell^{CBDL}_{R,t}\;\frac{s}{1-s}}{RER_t\;Q_{U\,t}\;K_{U\,t}} \tag{C.105}$$

Ratio of total CBDL to dometic lending excluding valuation effects

$$\xi^{CBDL}_{U,real,t} = \frac{AS_{R\,t}\;\ell^{CBDL}_{R,t}\;\frac{s}{1-s}}{RER_t\;K_{U\,t}} \tag{C.106}$$

Stochastic discount factor US Banks

$$\Omega_{U\,t} = \beta_U\;\frac{\Lambda_{U\,t}}{\Lambda_{U\,t-1}}\;\frac{1}{1+\pi^{C}_{U\,t}}\;\left(1 - \theta^{U}_{B} + \theta^{U}_{B}\;\left(n_{U\,t} + \frac{v_{U\,t} + \xi^{CBDL}_{U,t}\;v^{CBDL}_{U,t}}{1 + \Gamma^{CBDL}_{U,t}\;\xi^{CBDL}_{U,t}}\;\phi_{U\,t}\right)\right) \tag{C.107}$$

Discounted excess returns from cross border lending

$$v_{U,t}^{CBDL} = \Omega_{U\,t+1}\, \left(R_{U,t}^{CBDL} - (1 + R_{U\,t})\right) \tag{C.108}$$

CBDL risk premium in Dollar

$$RP_{U,t}^{CBDL} = \Phi_U^{\Gamma}\, \left(v_{U\,t} + \xi_{U,t}^{CBDL}\, v_{U,t}^{CBDL}\right)\, \frac{K_{U\,t}\, Q_{U\,t}\, RER_t\, \frac{(1-s)}{s}\, \xi_{U,t}^{CBDL}}{N_{R\,t}} \tag{C.109}$$

FOC optimal asset choice US

$$v_{U,t}^{CBDL} = RP_{E,b_t}^{F} + v_{U\,t}\, \Gamma_{U,t}^{CBDL} \tag{C.110}$$

Existing banks equity US

$$\begin{aligned} N_{U,e_t} =& \frac{1}{1+\pi_{U\,t}^{C}}\, \theta_B^U \Big(K_{U\,t-1}\, \left(R_{K,U_t} - (1 + R_{U\,t-1}) + \left(R_{U,t-1}^{CBDL} - (1 + R_{U\,t-1})\right)\, \xi_{U,t-1}^{CBDL}\right)\, Q_{U\,t-1} \\ &+ (1 + R_{U\,t-1})\, N_{U\,t-1}\Big) \end{aligned} \tag{C.111}$$

Definition of aggregate assets US banks

$$AS_{U\,t} = K_{U\,t}\, Q_{U\,t}\, \left(1 + \xi_{U,t}^{CBDL}\right) \tag{C.112}$$

Definition of of portfolio share of domestic investment (redundant)

$$(1 - \alpha_{U,t}^{CBDL}) = \frac{Q_{U\,t}\, K_{U\,t}}{AS_{U\,t}} \tag{C.113}$$

Definition of of portfolio share of CBDL investment US

$$\alpha_{U,t}^{CBDL} = \frac{CBDL_{R\,t}\, \frac{s}{1-s}}{AS_{U\,t}} \tag{C.114}$$

Risk weight adjusted optimal leverage ratio US

$$\phi_{U\,t} = \frac{n_{U\,t}\, \left((1 - \alpha_{U,t}^{CBDL}) + \Gamma_{U,t}^{CBDL} \alpha_{U,t}^{CBDL}\right)}{{\delta^U}_t\, \Big((1 - \alpha_{U,t}^{CBDL}) + \Gamma_{U,t}^{CBDL}\, \alpha_{U,t}^{CBDL}\Big) - v_{U,t}\,(1 - \alpha_{U,t}^{CBDL}) - v_{U,t}^{CBDL}\, \alpha_{U,t}^{CBDL}} \tag{C.115}$$

Aggregate Assets US (taking into account that $\phi_{U,t}$ is the *risk adjusted* leverage ratio in the code)

$$AS_{U\,t} = \frac{N_{U\,t}\, \phi_{U\,t}}{(1 - \alpha_{U,t}^{CBDL}) + \Gamma_{U,t}^{CBDL}\, \alpha_{U,t}^{CBDL}} \tag{C.116}$$

Cross border lending spread (in US-$)

$$S_{U,t}^{CBDL} = R_{U,t}^{CBDL} - (1 + R_{U\,t}) \tag{C.117}$$

C.5.4 Wage setting

Numerator Calvo style wages RoW

$$X_{1,R_t}^{w} = \kappa_w^R\, exp\left(\epsilon_{R\,t}^{W}\right)\, {w_{R\,t}}^{\psi_w\,(1+\varphi)}\, {L_{R\,t}}^{1+\varphi} + \beta_R\, \theta_w^R\, \left(1 + \pi_{R\,t+1}^{C}\right)^{\psi_w\,(1+\varphi)}\, X_{1,R_{t+1}}^{w} \tag{C.118}$$

Denominator Calvo style wages RoW

$$X^w_{2,R_t} = L_{R\,t}\,\Lambda_{R\,t}\,w_{R\,t}^{\psi_w} + \beta_R\,\theta^R_w\,\left(1+\pi^C_{R\,t+1}\right)^{\psi_w-1}\,X^w_{2,R_{t+1}} \tag{C.119}$$

Optimal real reset wage RoW

$$\tilde{w}_{R\,t}^{1+\psi_w\,\varphi} = \frac{X^w_{1,R_t}\,\frac{\psi_w}{\psi_w-1}}{X^w_{2,R_t}} \tag{C.120}$$

Evolution real wage RoW

$$w_{R\,t}^{1-\psi_w} = \left(1-\theta^R_w\right)\,\tilde{w}_{R\,t}^{1-\psi_w} + \theta^R_w\,\left(1+\pi^C_{R\,t}\right)^{\psi_w-1}\,w_{R\,t-1}^{1-\psi_w} \tag{C.121}$$

Labor supply shock RoW (redundant)

$$\epsilon^W_{R\ t} = \rho_w\,\epsilon^W_{R\ t-1} + \frac{\eta^W_{R\ t}}{100} \tag{C.122}$$

Numerator Calvo style wages US

$$X^w_{1,U_t} = \kappa^U_w\,exp\left(\epsilon^W_{U\ t}\right)\,w_{U\,t}^{\psi_w\,(1+\varphi)}\,L_{U\,t}^{1+\varphi} + \beta_U\,\theta^U_w\,\left(1+\pi^C_{U\,t+1}\right)^{\psi_w\,(1+\varphi)}\,X^w_{1,U_{t+1}} \tag{C.123}$$

Denominator Calvo style wages US

$$X^w_{2,U_t} = L_{U\,t}\,\Lambda_{U\,t}\,w_{U\,t}^{\psi_w} + \beta_U\,\theta^U_w\,\left(1+\pi^C_{U\,t+1}\right)^{\psi_w-1}\,X^w_{2,U_{t+1}} \tag{C.124}$$

Optimal real reset wage US

$$\tilde{w}_{U\,t}^{1+\psi_w\,\varphi} = \frac{\frac{\psi_w}{\psi_w-1}\,X^w_{1,U_t}}{X^w_{2,U_t}} \tag{C.125}$$

Evolution of real wage US

$$w_{U\,t}^{1-\psi_w} = \left(1-\theta^U_w\right)\,\tilde{w}_{U\,t}^{1-\psi_w} + \theta^U_w\,\left(1+\pi^C_{U\,t}\right)^{\psi_w-1}\,w_{U\,t-1}^{1-\psi_w} \tag{C.126}$$

Labour Supply Shock US (redundant)

$$\epsilon^W_{U\ t} = \rho_w\,\epsilon^W_{U\ t-1} + \frac{\eta^W_{U\ t}}{100} \tag{C.127}$$

C.5.5 Final Good Bundler

RoW demand for domestically produced goods

$$Y^R_{R\ t} = \eta_{R,t}\,exp\left(\epsilon^\eta_{R\,t}\right)\,IP_{R\,t}^{(-\psi_f)}\,Y^C_{R\ t} \tag{C.128}$$

RoW demand for import good from the US

$$Y^R_{U\ t} = Y^C_{R\ t}\,\frac{n}{1-n}\,\left(1-\eta_{R,t}\,exp\left(\epsilon^\eta_{R\,t}\right)\right)\,\left(IP_{R\,t}\,IT^U_{R\ t}\right)^{(-\psi_f)} \tag{C.129}$$

RoW home bias shock (redundant)

$$\epsilon^{\eta}_{Rt} = \rho_{\eta}\, \epsilon^{\eta}_{Rt-1} + \frac{\eta^{\eta}_{Rt}}{100} \tag{C.130}$$

US demand for domestically produced goods

$$Y^{U}_{U\ t} = \eta_{F,t}\, exp\left(\epsilon^{\eta}_{U\,t}\right)\ IP_{U\,t}^{(-\psi_f)}\, Y^{C}_{U\ t} \tag{C.131}$$

US demand for for import good from RoW

$$Y^{U}_{R\ t} = Y^{C}_{U\ t}\, \frac{1-n}{n}\ \left(1-\eta_{F,t}\, exp\left(\epsilon^{\eta}_{U\,t}\right)\right)\ \left(IP_{U\,t}\, IT^{R}_{U\ t}\right)^{(-\psi_f)} \tag{C.132}$$

Definition of US imports (in US per capita units)

$$Imp_{U\,t} = Y^{C}_{U\ t}\ \left(1-eta_{F,t}\, exp\left(\epsilon^{\eta}_{U\,t}\right)\right)\ \left(IP_{U\,t}\, IT^{R}_{U\ t}\right)^{(-\psi_f)} \tag{C.133}$$

US home bias shock (redundant)

$$\epsilon^{\eta}_{U\,t} = \rho_{\eta}\, \epsilon^{\eta}_{U\,t-1} + \frac{\eta^{\eta}_{U\,t}}{100} \tag{C.134}$$

Definition of US export import ratio

$$\frac{Exp}{imp}_{U\,t} = \frac{{Y_{U,t}}^{R}}{Imp_{U\,t}} \tag{C.135}$$

C.5.6 Intermediate Goods producers

Depreciation Function RoW

$$\tau_{Rt} = \tau_{R,ss,scale} + \frac{\zeta^{R}_{1} U^{1+\zeta_2}_{R_t}}{1+\zeta_2} \tag{C.136}$$

Derivative Depreciation Function RoW

$$\tau'_{Rt} = \zeta^{R}_{1} U_{Rt}^{\zeta_2} \tag{C.137}$$

Optimal RoW capital services to labor ratio (implicitly defining optimal utilization)

$$\frac{w_{Rt}}{\tau'_{Rt}} = \frac{\frac{1-\alpha}{\alpha}\, K_{Rt-1}\, U_{Rt}}{L_{Rt}} \tag{C.138}$$

Real marginal costs in CPI terms RoW

$$MC^{r}_{Rt} = \frac{{w_{Rt}}^{1-\alpha}\, {\tau'_{Rt}}^{\alpha}}{(1-\alpha)^{1-\alpha}\ \alpha^{\alpha}} \tag{C.139}$$

Real marginal costs in PPI terms RoW

$$MC^{rp}_{R\ t} = \frac{MC^{r}_{Rt}}{IP_{Rt}} \tag{C.140}$$

RoW gross returns to capital

$$R_{K,R_t} = \left(1 + \pi^C_{Rt}\right) \frac{Q_{Rt} + \frac{\alpha\, MC^r_{Rt}\, Z_{Rt}}{K_{Rt-1}} - \tau_{Rt}}{Q_{Rt-1}} \tag{C.141}$$

Depreciation Function US

$$\tau_{Ut} = \tau_{U,ss,scale} + \frac{\zeta^U_1\, U^{1+\zeta_2}_{U_t}}{1+\zeta_2} \tag{C.142}$$

Derivative Depreciation Function US

$$\tau'_{Ut} = \zeta^U_1\, U_{Ut}^{\zeta_2} \tag{C.143}$$

Optimal US capital services to labor ratio (implicitly defining optimal utilization)

$$\frac{w_{Ut}}{\tau'_{Ut}} = \frac{\frac{1-\alpha}{\alpha}\, K_{Ut-1}\, U_{Ut}}{L_{Ut}} \tag{C.144}$$

Real marginal costs in US CPI

$$MC^r_{Ut} = \frac{w_{Ut}^{1-\alpha}\, {\tau'_{Ut}}^{\alpha}}{(1-\alpha)^{1-\alpha}\, \alpha^\alpha} \tag{C.145}$$

Real marginal costs in US PPI terms

$$MC^{rp}_{U\,t} = \frac{MC^r_{Ut}}{IP_{Ut}} \tag{C.146}$$

US gross returns to capital

$$R_{K,U_t} = \left(1 + \pi^C_{Ut}\right) \frac{Q_{Ut} + \frac{\alpha\, MC^r_{Ut}\, Z_{Ut}}{K_{Ut-1}} - \tau_{Ut}}{Q_{Ut-1}} \tag{C.147}$$

C.5.7 RoW Capital Goods Producers

RoW Tobins Q/RoW Price of Capital

$$\begin{aligned} Q_{Rt} =& 1 + \frac{\Psi_R}{2} \left(\frac{In_{Rt} + (\bar{I_R})}{(\bar{I_R}) + In_{Rt-1}} - 1\right)^2 + \frac{In_{Rt} + (\bar{I_R})}{(\bar{I_R}) + In_{Rt-1}}\, \Psi_R \left(\frac{In_{Rt} + (\bar{I_R})}{(\bar{I_R}) + In_{Rt-1}} - 1\right) \\ & - \Psi_R\, \frac{\beta_R\, \Lambda_{Rt+1}}{\Lambda_{Rt}} \left(\frac{(\bar{I_R}) + In_{Rt+1}}{In_{Rt} + (\bar{I_R})} - 1\right) \left(\frac{(\bar{I_R}) + In_{Rt+1}}{In_{Rt} + (\bar{I_R})}\right)^2 \end{aligned} \tag{C.148}$$

RoW LOM for capital

$$K_{Rt} = K_{Rt-1} + In_{Rt} \tag{C.149}$$

Definition of net investment

$$In_{Rt} = I_{Rt} - K_{Rt-1}\, \tau_{Rt} \tag{C.150}$$

US Tobins Q/US Price of Capital

$$\begin{aligned} Q_{U\,t} =& 1 + \frac{\Psi_U}{2}\, \left(\frac{In_{U\,t} + (\bar{I_U})}{(\bar{I_U}) + In_{U\,t-1}} - 1\right)^2 + \frac{In_{U\,t} + (\bar{I_U})}{(\bar{I_U}) + In_{U\,t-1}}\, \Psi_U\, \left(\frac{In_{U\,t} + (\bar{I_U})}{(\bar{I_U}) + In_{U\,t-1}} - 1\right) \\ & - \Psi_U\, \frac{\beta_U\, \Lambda_{U\,t+1}}{\Lambda_{U\,t}}\, \left(\frac{(\bar{I_U}) + In_{U\,t+1}}{In_{U\,t} + (\bar{I_U})} - 1\right) \left(\frac{(\bar{I_U}) + In_{U\,t+1}}{In_{U\,t} + (\bar{I_U})}\right)^2 \end{aligned} \tag{C.151}$$

US LOM for capital

$$K_{U\,t} = K_{U\,t-1} + In_{U\,t} \tag{C.152}$$

US definition of net investment

$$In_{U\,t} = I_{U\,t} - K_{U\,t-1}\, \tau_{U\,t} \tag{C.153}$$

C.5.8 Intra RoW retail good pricing

Numerator Calvo pricing PCP intra RoW sales

$$\tilde{X}^R_{R,1_t} = Y^R_{R\,t}\, MC^{rp}_{R\,t}\, IP_{Rt}\, \Lambda_{Rt}\, \widetilde{CP}^{R(-\psi_i)}_{Rt} + \beta_R\, \theta^R_P\, \left(1 + \tilde{\pi}^R_{Rt+1}\right)^{\psi_i}\, \tilde{X}^R_{R,1_{t+1}} \tag{C.154}$$

Denominator Calvo pricing PCP intra RoW sales

$$\tilde{X}^R_{R,2_t} = Y^R_{R\,t}\, IP_{Rt}\, \Lambda_{Rt}\, \widetilde{CP}^{R^{1-\psi_i}}_{Rt} + \beta_R\, \theta^R_P\, \left(1 + \tilde{\pi}^R_{Rt+1}\right)^{\psi_i - 1}\, \tilde{X}^R_{R,2_{t+1}} \tag{C.155}$$

Optimal reset price Calvo pricing PCP intra RoW sales

$$\tilde{p}^R_{Rt} = \frac{\tilde{X}^R_{R,1_t}\, \frac{\psi_i}{\psi_i - 1}}{\tilde{X}^R_{R,2_t}} \tag{C.156}$$

RoW domestic sales PCP retailers inflation

$$1 = \left(1 - \theta^R_P\right)\, \tilde{p}^{R^{1-\psi_i}}_{Rt} + \theta^R_P\, \left(1 + \tilde{\pi}^R_{Rt}\right)^{\psi_i - 1} \tag{C.157}$$

Numerator Calvo pricing DCP intra RoW sales

$$\hat{X}^R_{R,1_t} = Y^R_{R\,t}\, MC^{rp}_{R\,t}\, IP_{Rt}\, \Lambda_{Rt}\, \widehat{CP}^{R(-\psi_i)}_{Rt} + \beta_R\, \theta^R_P\, \left(1 + \hat{\pi}^R_{Rt+1}\right)^{\psi_i}\, \hat{X}^R_{R,1_{t+1}} \tag{C.158}$$

Denominator Calvo pricing DCP intra RoW sales

$$\hat{X}^R_{R,2_t} = Y^R_{R\,t}\, IP_{Rt}\, \Lambda_{Rt}\, \widehat{CP}^{R^{1-\psi_i}}_{Rt} + \beta_R\, \theta^R_P\, \left(1 + \hat{\pi}^R_{Rt+1}\right)^{\psi_i - 1}\, \hat{X}^R_{R,2_{t+1}} \tag{C.159}$$

Optimal reset price Calvo pricing DCP intra RoW sales

$$\hat{p}^R_{Rt} = \frac{\frac{\psi_i}{\psi_i - 1}\, \hat{X}^R_{R,1_t}}{\hat{X}^R_{R,2_t}} \tag{C.160}$$

RoW domestic sales DCP retailers inflation

$$1 = \left(1 - \theta^R_P\right)\, \hat{p}^{R^{1-\psi_i}}_{Rt} + \theta^R_P\, \left(1 + \hat{\pi}^R_{Rt}\right)^{\psi_i - 1} \tag{C.161}$$

C.5.9 Intra US retail good pricing

Numerator Calvo pricing intra US sales

$$X^U_{U,1_t} = Y^U_{U\,t}\, MC^{rp}_{U\,t}\, \Lambda_{U\,t}\, IP_{U\,t} + \beta_U\, \theta^U_P\, \left(1 + \pi^U_{U\,t+1}\right)^{\psi_i}\, X^U_{U,1_{t+1}} \tag{C.162}$$

Denominator Calvo pricing intra US sales

$$X^U_{U,2_t} = Y^U_{U\,t}\, \Lambda_{U\,t}\, IP_{U\,t} + \beta_U\, \theta^U_P\, \left(1 + \pi^U_{U\,t+1}\right)^{\psi_i - 1}\, X^U_{U,2_{t+1}} \tag{C.163}$$

Optimal reset price Calvo pricing intra US sales

$$\overline{p}^U_{U\,t} = \frac{\frac{\psi_i}{\psi_i - 1}\, X^U_{U,1_t}}{X^U_{U,2_t}} \tag{C.164}$$

US domestic retail good price inflation

$$1 = \left(1 - \theta^U_P\right)\, \overline{p}^{U\,1-\psi_i}_{U\,t} + \theta^U_P\, \left(1 + \pi^U_{U\,t}\right)^{\psi_i - 1} \tag{C.165}$$

C.5.10 Export Pricing

Numerator Calvo Pricing RoW PCP exports to US

$$\tilde{X}^U_{R,1_t} = Y^U_{R\,t}\, IP_{R\,t}\, MC^{rp}_{R\,t}\, \Lambda_{R\,t}\, \widetilde{CP}^{U\,(-\psi_i)}_{R\,t} + \beta_R\, \theta^R_P\, \left(1 + \tilde{\pi}^U_{R\,t+1}\right)^{\psi_i}\, \tilde{X}^U_{R,1_{t+1}} \tag{C.166}$$

Denominator Calvo Pricing RoW PCP exports to US

$$\tilde{X}^U_{R,2_t} = Y^U_{R\,t}\, IP_{R\,t}\, \Lambda_{R\,t}\, \widetilde{CP}^{U\,(-\psi_i)}_{R\,t}\, \widetilde{EM}^U_{R\,t} + \beta_R\, \theta^R_P\, \left(1 + \tilde{\pi}^U_{R\,t+1}\right)^{\psi_i - 1}\, \tilde{X}^U_{R,2_{t+1}} \tag{C.167}$$

Optimal reset price Calvo Pricing RoW PCP exports to US

$$\tilde{p}^U_{R\,t} = \frac{\frac{\psi_i}{\psi_i - 1}\, \tilde{X}^U_{R,1_t}}{\tilde{X}^U_{R,2_t}} \tag{C.168}$$

PCP price inflation RoW exports to US

$$1 = \left(1 - \theta^R_P\right)\, \tilde{p}^{U\,1-\psi_i}_{R\,t} + \theta^R_P\, \left(1 + \tilde{\pi}^U_{R\,t}\right)^{\psi_i - 1} \tag{C.169}$$

Numerator Calvo Pricing RoW DCP exports to US

$$\hat{X}^U_{R,1_t} = Y^U_{R\,t}\, IP_{R\,t}\, MC^{rp}_{R\,t}\, \Lambda_{R\,t}\, \widehat{CP}^{U\,(-\psi_i)}_{R\,t} + \beta_R\, \theta^R_P\, \left(1 + \hat{\pi}^U_{R\,t+1}\right)^{\psi_i}\, \hat{X}^U_{R,1_{t+1}} \tag{C.170}$$

Denominator Calvo Pricing RoW DCP exports to US

$$\hat{X}^U_{R,2_t} = Y^U_{R\,t}\, IP_{R\,t}\, \Lambda_{R\,t}\, \widehat{CP}^{U\,(-\psi_i)}_{R\,t}\, \widehat{EM}^U_{R\,t} + \beta_R\, \theta^R_P\, \left(1 + \hat{\pi}^U_{R\,t+1}\right)^{\psi_i - 1}\, \hat{X}^U_{R,2_{t+1}} \tag{C.171}$$

Optimal reset price Calvo Pricing RoW DCP exports to US

$$\hat{p}^U_{R\,t} = \frac{\frac{\psi_i}{\psi_i - 1}\, \hat{X}^U_{R,1_t}}{\hat{X}^U_{R,2_t}} \tag{C.172}$$

DCP price inflation RoW exports to US

$$1 = \left(1 - \theta_P^R\right) \hat{p}_{Rt}^{U\,1-\psi_i} + \theta_P^R \left(1 + \hat{\pi}_{R\,t}^{U}\right)^{\psi_i - 1} \tag{C.173}$$

Numerator Calvo Pricing US DCP exports to RoW

$$\tilde{X}_{U,1_t}^R = Y_{U\,t}^R \, IP_{U\,t} \, MC_{U\,t}^{rp} \, \Lambda_{U\,t} \, \widetilde{CP}_{U\,t}^{R(-\psi_i)} + \beta_U \, \theta_P^U \left(1 + \tilde{\pi}_{U\,t+1}^R\right)^{\psi_i} \tilde{X}_{U,1_{t+1}}^R \tag{C.174}$$

Denominator Calvo Pricing US DCP exports to RoW

$$\tilde{X}_{U,2_t}^R = Y_{U\,t}^R \, IP_{U\,t} \, \Lambda_{U\,t} \, \widetilde{CP}_{U\,t}^{R(-\psi_i)} \, \widetilde{EM}_{U\,t}^R + \beta_U \, \theta_P^U \left(1 + \tilde{\pi}_{U\,t+1}^R\right)^{\psi_i - 1} \tilde{X}_{U,2_{t+1}}^R \tag{C.175}$$

Optimal reset price Calvo Pricing US DCP exports to RoW

$$\tilde{p}_{U\,t}^R = \frac{\frac{\psi_i}{\psi_i - 1} \tilde{X}_{U,1_t}^R}{\tilde{X}_{U,2_t}^R} \tag{C.176}$$

DCP price inflation US exports to RoW

$$1 = \left(1 - \theta_P^U\right) \tilde{p}_{U\,t}^{R\,1-\psi_i} + \theta_P^U \left(1 + \tilde{\pi}_{U\,t}^R\right)^{\psi_i - 1} \tag{C.177}$$

Numerator Calvo Pricing US LCP exports to RoW

$$\underline{X}_{U,1_t}^R = Y_{U\,t}^R \, IP_{U\,t} \, MC_{U\,t}^{rp} \, \Lambda_{U\,t} \, \underline{CP}_{U\,t}^{R(-\psi_i)} + \beta_U \, \theta_P^U \left(1 + \underline{\pi}_{U\,t+1}^R\right)^{\psi_i} \underline{X}_{U,1_{t+1}}^R \tag{C.178}$$

Denominator Calvo Pricing US LCP exports to RoW

$$\underline{X}_{U,2_t}^R = Y_{U\,t}^R \, IP_{U\,t} \, \Lambda_{U\,t} \, \underline{CP}_{U\,t}^{R(-\psi_i)} \, \underline{EM}_{U\,t}^R + \beta_U \, \theta_P^U \left(1 + \underline{\pi}_{U\,t+1}^R\right)^{\psi_i - 1} \underline{X}_{U,2_{t+1}}^R \tag{C.179}$$

Optimal reset price Calvo Pricing US LCP exports to RoW

$$\underline{p}_{U\,t}^R = \frac{\frac{\psi_i}{\psi_i - 1} \underline{X}_{U,1_t}^R}{\underline{X}_{U,2_t}^R} \tag{C.180}$$

LCP price inflation US exports to RoW

$$1 = \left(1 - \theta_P^U\right) \underline{p}_{U\,t}^{R\,1-\psi_i} + \theta_P^U \left(1 + \underline{\pi}_{U\,t}^R\right)^{\psi_i - 1} \tag{C.181}$$

C.5.11 Monetary Policy

RoW Taylor rule

$$\frac{1 + R_{Rt}}{1 + R_{E,SS}} = \left(\frac{1 + R_{Rt-1}}{1 + R_{E,SS}}\right)^{\rho_{R,r}} \left(\left(\frac{1 + \pi_{Rt}^C}{1 + (\bar{\pi_R^C})}\right)^{\phi_{R,\pi}} \left(\frac{Z_{Rt}}{(\bar{Z_R})}\right)^{\phi_{R,z}}\right)^{1-\rho_{R,r}} exp\left(\varepsilon_{Rt}^R\right) \tag{C.182}$$

US Taylor rule

$$\frac{1 + R_{U\,t}}{1 + RSS_{F,ss}} = \left(\frac{1 + R_{U\,t-1}}{1 + R_{F,ss}}\right)^{\rho_{U,r}} \left(\left(\frac{1 + \pi_{U\,t}^C}{1 + (\bar{\pi_U^C})}\right)^{\phi_{U,\pi}} \left(\frac{Z_{U\,t}}{(\bar{Z_U})}\right)^{\phi_{U,z}}\right)^{1-\rho_{U,r}} exp\left(\varepsilon_{U\,t}^R\right) \tag{C.183}$$

RoW MP shock

$$\varepsilon^{R}_{Rt} = \rho^{r}_{\epsilon}\, \varepsilon^{R}_{Rt-1} + \sigma^{r}_{R,\epsilon}\, \eta^{r}_{Rt} \tag{C.184}$$

US MP shock

$$\varepsilon^{R}_{Ut} = \rho^{r}_{\epsilon}\, \varepsilon^{R}_{U\,t-1} + \frac{\sigma^{r}_{U,\epsilon}}{100}\, \eta^{r}_{U\,t} \tag{C.185}$$

C.5.12 Relative Prices

Relative price of RoW domestic DCP sales and RoW domestic PCP sales

$$\hat{IT}^{R}_{Rt} = \hat{IT}^{R}_{Rt-1} \frac{\left(1+D\mathcal{E}_t\right)\left(1+\hat{\pi}^{R}_{Rt}\right)}{1+\tilde{\pi}^{R}_{Rt}} \tag{C.186}$$

Relative price of RoW domestic PCP sales to Aggregate RoW PPI

$$\widetilde{CP}^{R}_{Rt} = \left(\gamma^{R,PCP}_{R} + \left(1-\gamma^{R,PCP}_{R}\right) \hat{IT}^{R}_{Rt}{}^{1-\psi_i}\right)^{\frac{1}{\psi_i - 1}} \tag{C.187}$$

Relative price of RoW domestic DCP sales to Aggregate RoW PPI

$$\widehat{CP}^{R}_{Rt} = \widetilde{CP}^{R}_{Rt}\, \hat{IT}^{R}_{Rt} \tag{C.188}$$

Aggregate RoW PPI inflation as a function of domestic PCP and DCP prices

$$1+\pi^{R}_{Rt} = \left(1+\tilde{\pi}^{R}_{Rt}\right) \frac{\widetilde{CP}^{R}_{Rt-1}}{\widetilde{CP}^{R}_{Rt}} \tag{C.189}$$

Export margins for DCP exports from RoW to US in RoW currency (price of DCP exports over domestic sales price)

$$\widehat{EM}^{U}_{Rt} = \widehat{EM}^{U}_{Rt-1} \frac{\left(1+D\mathcal{E}_t\right)\left(1+\hat{\pi}^{U}_{Rt}\right)}{1+\pi^{R}_{Rt}} \tag{C.190}$$

Export margins for PCP exports from RoW to US in RoW currency (price of DCP exports over domestic sales price)

$$\widetilde{EM}^{U}_{Rt} = \widetilde{EM}^{U}_{Rt-1} \frac{1+\tilde{\pi}^{U}_{Rt}}{1+\pi^{R}_{Rt}} \tag{C.191}$$

Aggregate margins for exports from RoW to US in RoW currency (agg. export price over domestic sales PPI)

$$EM^{U}_{R\,t} = \left(\gamma^{R,PCP}_{U,t}\, \widetilde{EM}^{U}_{Rt}{}^{1-\psi_i} + \left(1-\gamma^{R,PCP}_{U,t}\right) \widehat{EM}^{U}_{Rt}{}^{1-\psi_i}\right)^{\frac{1}{1-\psi_i}} \tag{C.192}$$

Import price inflation of US imports from the RoW in US-D

$$1+\pi^{R^I}_{U\ t} = \frac{\left(1+\pi^{R}_{Rt}\right) \frac{EM^{U}_{R\,t}}{EM^{U}_{R\,t-1}}}{1+D\mathcal{E}_t} \tag{C.193}$$

Export margins for PCP exports from the US to RoW in US-D (price of PCP exports over domestic sales price)

$$\widetilde{EM}^{R}_{U\,t} = \widetilde{EM}^{R}_{U\,t-1}\,\frac{1+\tilde{\pi}^{R}_{U\,t}}{1+\pi^{U}_{U\,t}} \tag{C.194}$$

Export margins for LCP exports from the US to RoW in US-D (price of LCP exports over domestic sales price)

$$\underline{EM}^{R}_{U\,t} = \frac{\left(1+\underline{\pi}^{R}_{U\,t}\right)\,\frac{\underline{EM}^{R}_{U\,t-1}}{1+D\mathcal{E}_t}}{1+\pi^{U}_{U\,t}} \tag{C.195}$$

Aggregate margins for exports from US to RoW in US-D currency (agg. export price over domestic sales PPI)

$$EM^{R}_{U\,t} = \left(\gamma^{F,PCP}_{E,t}\,\widetilde{EM}^{R\,1-\psi_i}_{U\,t} + \left(1-\gamma^{F,PCP}_{E,t}\right)\,\underline{EM}^{R\,1-\psi_i}_{U\,t}\right)^{\frac{1}{1-\psi_i}} \tag{C.196}$$

Import price inflation of RoW imports from the US in RoW currency

$$1+\pi^{U^I}_{R\ \ t} = \left(1+D\mathcal{E}_t\right)\,\left(1+\pi^{U}_{U\,t}\right)\,\frac{EM^{R}_{U\,t}}{EM^{R}_{U\,t-1}} \tag{C.197}$$

Interior terms of trade RoW (US exports prices (in RoW currency) relative to RoW PPI)

$$IT^{U}_{R\,t} = IT^{U}_{R\,t-1}\,\frac{1+\pi^{U^I}_{R\ \ t}}{1+\pi^{R}_{R\,t}} \tag{C.198}$$

Interior Producer Price RoW (PPI over CPI)

$$IP_{R\,t} = \left(\eta_{R,t} + \left(1-\eta_{R,t}\right)\,IT^{U\,1-\psi_f}_{R\,t}\right)^{\frac{1}{\psi_f-1}} \tag{C.199}$$

RoW CPI inflation

$$1+\pi^{C}_{R\,t} = \left(1+\pi^{R}_{R\,t}\right)\,\frac{IP_{R\,t-1}}{IP_{R\,t}} \tag{C.200}$$

Interior terms of trade US (RoW exports prices (in US-D currency) relative to US PPI)

$$IT^{R}_{U\,t} = \frac{EM^{U}_{R\,t}\,EM^{R}_{U\,t}}{IT^{U}_{R\,t}} \tag{C.201}$$

Interior Producer Price US (PPI over CPI)

$$IP_{U\,t} = \left(\eta_{U,t} + \left(1-\eta_{U,t}\right)\,IT^{R\,1-\psi_f}_{U\,t}\right)^{\frac{1}{\psi_f-1}} \tag{C.202}$$

US consumer price inflation

$$1+\pi^{C}_{U\,t} = \left(1+\pi^{U}_{U\,t}\right)\,\frac{IP_{U\,t-1}}{IP_{U\,t}} \tag{C.203}$$

Definition of the Real exchange rate (in terms of CPI baskets)

$$RER_t = \frac{IP_{R\,t}\,EM^{U}_{R\,t}}{IP_{U\,t}\,IT^{R}_{U\,t}} \tag{C.204}$$

PCP export price over agg. US import price

$$\widetilde{CP}^{U}_{Rt} = \frac{IP_{Rt}\,\frac{\widetilde{EM}^{U}_{Rt}}{IT^{R}_{U\,t}}}{IP_{U\,t}}\,\frac{1}{RER_t} \tag{C.205}$$

DCP export price over agg. US import price

$$\widehat{CP}^{U}_{Rt} = \frac{1}{RER_t}\,\frac{IP_{Rt}\,\frac{\widehat{EM}^{U}_{Rt}}{IT^{R}_{U\,t}}}{IP_{U\,t}} \tag{C.206}$$

DCP export price over agg. RoW import price

$$\widetilde{CP}^{R}_{U\,t} = RER_t\,\frac{IP_{U\,t}\,\frac{\widetilde{EM}^{R}_{U\,t}}{IT^{U}_{R\,t}}}{IP_{Rt}} \tag{C.207}$$

LCP export price over agg. RoW import price

$$\underline{CP}^{R}_{U\,t} = RER_t\,\frac{IP_{U\,t}\,\frac{\underline{EM}^{R}_{U\,t}}{IT^{U}_{R\,t}}}{IP_{Rt}} \tag{C.208}$$

C.5.13 Market Clearing

Agg. demand for RoW final composite good

$$Y^{C}_{R\,t} = C_{Rt} + I_{Rt} + \left(In_{Rt} + (\bar{I_R})\right)\,\frac{\Psi_R}{2}\,\left(\frac{In_{Rt} + (\bar{I_R})}{(\bar{I_R}) + In_{Rt-1}} - 1\right)^2 \tag{C.209}$$

Agg. demand for US final composite good

$$Y^{C}_{U\,t} = C_{U\,t} + I_{U\,t} + \left(In_{U\,t} + (\bar{I_U})\right)\,\frac{\Psi_U}{2}\,\left(\frac{In_{U\,t} + (\bar{I_U})}{(\bar{I_U}) + In_{U\,t-1}} - 1\right)^2 \tag{C.210}$$

RoW aggregate production function

$$Z_{Rt} = \left(K_{Rt-1}\,U_{Rt}\right)^{\alpha}\,L_{Rt}^{1-\alpha} \tag{C.211}$$

US aggregate production

$$Z_{U\,t} = \left(K_{U\,t-1}\,U_{U\,t}\right)^{\alpha}\,L_{U\,t}^{1-\alpha} \tag{C.212}$$

RoW market clearing

$$Z_{Rt} = Y^{R}_{R\,t}\,\delta^{R}_{Rt} + Y^{U}_{R\,t}\,\delta^{U}_{Rt} \tag{C.213}$$

US market clearing

$$Z_{U\,t} = Y^{U}_{U\,t}\,\delta^{U}_{U\,t} + Y^{R}_{U\,t}\,\delta^{U}_{Rt} \tag{C.214}$$

C.5.14 Price dispersion terms

$$\tilde{\delta}^R_{Rt} = \left(1-\theta^R_P\right)\tilde{p}^{R(-\psi_i)}_{Rt} + \theta^R_P\left(1+\tilde{\pi}^R_{Rt}\right)^{\psi_i}\tilde{\delta}^R_{Rt-1} \tag{C.215}$$

$$\hat{\delta}^R_{Rt} = \left(1-\theta^R_P\right)\hat{p}^{R(-\psi_i)}_{Rt} + \theta^R_P\left(1+\hat{\pi}^R_{Rt}\right)^{\psi_i}\hat{\delta}^R_{Rt-1} \tag{C.216}$$

$$\delta^R_{Rt} = \tilde{\delta}^R_{Rt}\widetilde{CP}^{R(-\psi_i)}_{Rt}\gamma^{R,PCP}_{R,t} + \hat{\delta}^R_{Rt}\widehat{CP}^{R(-\psi_i)}_{Rt}\left(1-\gamma^{R,PCP}_{R,t}\right) \tag{C.217}$$

$$\tilde{\delta}^U_{Rt} = \left(1-\theta^R_P\right)\tilde{p}^{U(-\psi_i)}_{Rt} + \theta^R_P\left(1+\tilde{\pi}^U_{Rt}\right)^{\psi_i}\tilde{\delta}^U_{Rt-1} \tag{C.218}$$

$$\hat{\delta}^U_{Rt} = \left(1-\theta^R_P\right)\hat{p}^{U(-\psi_i)}_{Rt} + \theta^R_P\left(1+\hat{\pi}^U_{Rt}\right)^{\psi_i}\hat{\delta}^U_{Rt-1} \tag{C.219}$$

$$\delta^U_{Rt} = \tilde{\delta}^U_{Rt}\widetilde{CP}^{U(-\psi_i)}_{Rt}\gamma^{R,PCP}_{U,t} + \hat{\delta}^U_{Rt}\widehat{CP}^{U(-\psi_i)}_{Rt}\left(1-\gamma^{R,PCP}_{U,t}\right) \tag{C.220}$$

$$\delta^U_{U\,t} = \left(1-\theta^U_P\right)\overline{p}^{U(-\psi_i)}_{U\,t} + \theta^U_P\left(1+\pi^U_{U\,t}\right)^{\psi_i}\delta^U_{U\,t-1} \tag{C.221}$$

$$\tilde{\delta}^R_{U\,t} = \left(1-\theta^U_P\right)\tilde{p}^{R(-\psi_i)}_{U\,t} + \theta^U_P\left(1+\tilde{\pi}^R_{U\,t}\right)^{\psi_i}\tilde{\delta}^R_{U\,t-1} \tag{C.222}$$

$$\underline{\delta}^R_{U\,t} = \left(1-\theta^U_P\right)\underline{p}^{R(-\psi_i)}_{U\,t} + \theta^U_P\left(1+\underline{\pi}^R_{U\,t}\right)^{\psi_i}\underline{\delta}^R_{U\,t-1} \tag{C.223}$$

$$\delta^R_{U\,t} = \tilde{\delta}^R_{U\,t}\widetilde{CP}^{R(-\psi_i)}_{U\,t}\gamma^{U,PCP}_{R,t} + \underline{\delta}^R_{U\,t}\underline{CP}^{R(-\psi_i)}_{U\,t}\left(1-\gamma^{U,PCP}_{R,t}\right) \tag{C.224}$$

C.5.15 Balance of Payments

<u>RoW Current account in RoW currency</u>

$$CA^F_{R,nom_t} = Y^R_{R\,t}\,IP_{Rt} + Y^U_{R\,t}\,IT^R_{U\,t}\,RER_t\,IP_{U\,t} - Y^C_{R\,t} \tag{C.225}$$

<u>Balance of Payments</u>

$$RER_t\left(GB_{Rt} - \frac{R^{GB}_{R\ t-1}}{1+\pi^C_{U\,t}}\left(GB_{Rt-1}\right)\right) - RER_t\left(CBDL_{Rt} - CBDL_{Rt-1}\frac{R^{CBDL}_{U,t-1}}{1+\pi^C_{U\,t}}\right) = CA^F_{R,nom_t} \tag{C.226}$$

<u>Trade Balance RoW</u>

$$TB_{Rt} = Y^U_{R\,t} - \frac{(1-n)\,Y^R_{U\,t}}{n} \tag{C.227}$$

Change in the NFA (including valuation effects) relative to RoW GDP

$$\Delta NFA_{Rt} = \frac{RER_t\,(GB_{Rt-1}) - (GB_{Rt-1})\,RER_{t-1} - RER_t\,CBDL_{Rt-1} + CBDL_{Rt-1}\,RER_{t-1}}{(\bar{Z_R})} \tag{C.228}$$

C.5.16 Model local variables

Share of PCP goods in US Import Basket

$$\gamma_U^{R,PCP} = 1 - \hat{\gamma}_U^R$$

Share of PCP goods in RoW Import basket

$$\gamma_R^{U,PCP} = 1 - \tilde{\gamma}_R^U$$

Share of PCP goods in RoW Local Basket

$$\gamma_R^{R,PCP} = 1 - \hat{\gamma}_R^R$$

RoW steady state net interest rate

$$R_{R,SS} = \frac{1}{\beta_R} - 1$$

US steady state net interest rate

$$R_{R,SS} = \frac{1}{\beta_U} - 1$$

Size adjusted import share RoW

$$\eta_R = (1 - op_R)\,(1 - s)$$

Size adjusted import share US

$$\eta_{US} = (1 - op_U)s$$

C.6 What if the FED stabilized the US$ in the trinity model?

A natural question to ask is if the empirical policy rule counterfactual in section 4.2 agrees with the implications from the version of the trinity model, where we change the US policy rule from a Taylor Rule to a rule according to which the Federal Reserve stabilizes the US$.[109] We simulate the model under the counterfactual policy rule and plot the results in figures C.27, C.28, C.29. In line with the—in particular the VAR-based policy-rule—counterfactuals in the paper the global contraction caused by a global risk aversion shock is mitigated if the Fed follows such a policy rule that stabilizes the dollar in the trinity model.

[109]We thank an anonymous referee for this suggestion.

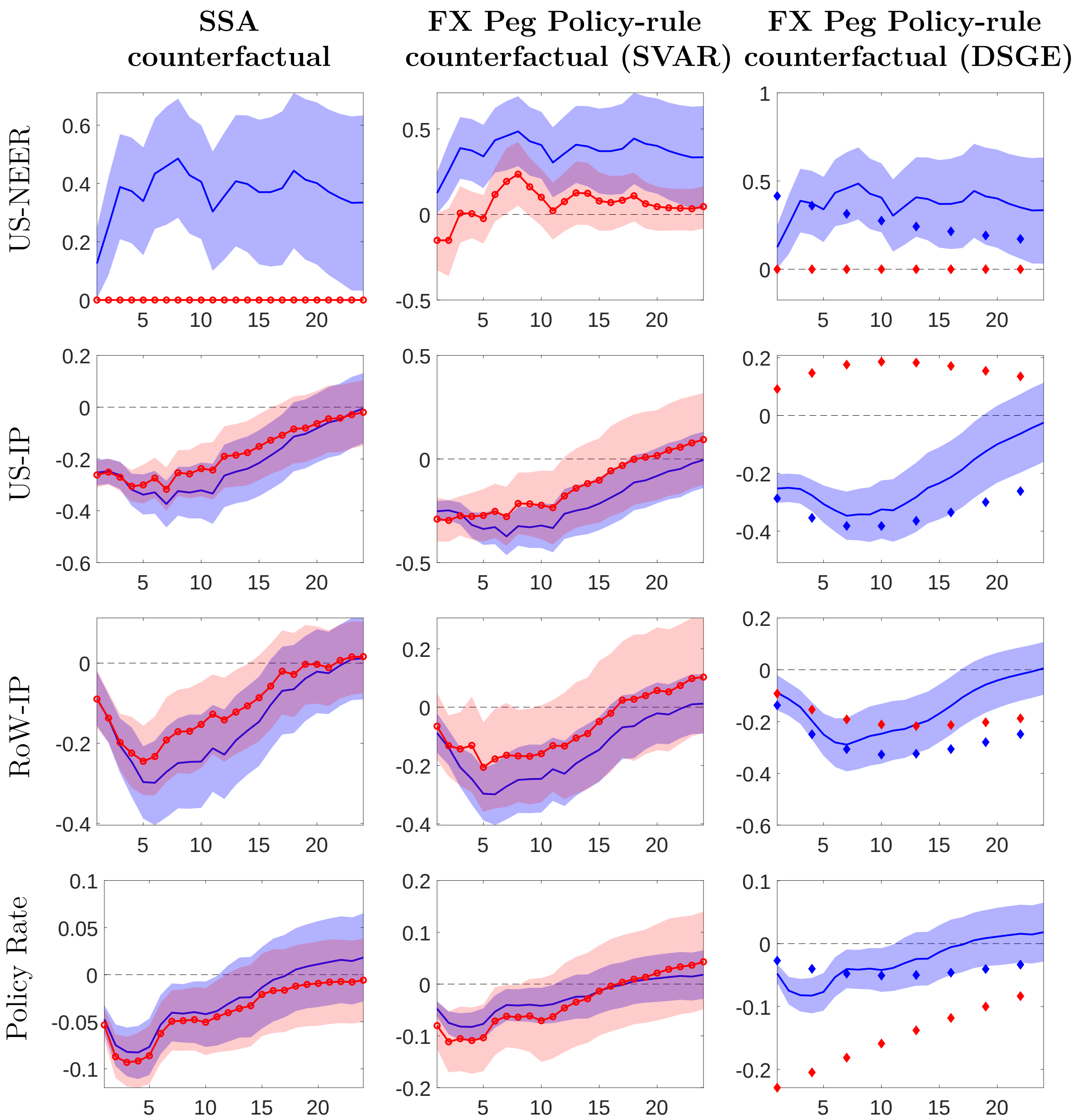
Quantitatively, in the trinity-model policy-rule counterfactual the contractionary effect on US output is mitigated more than in the VAR-based policy-rule counterfactual. In fact, when the Fed follows a policy rule that stabilizes the dollar in the trinity model US output even rises following a global risk aversion shock. This is because the Fed has to loosen much more to stabilize the dollar in the trinity-model counterfactual than in the corresponding VAR-based counterfactual. The loosening in the trinity-model counterfactual is so much larger that it even overturns the contractionary effect of the global risk shock on US output.

Recall that McKay and Wolf (2023) prove the equivalence of a change in the structural policy-rule and the method based on policy (news) shocks, which we apply in the paper. Therefore, the empirical policy-rule counterfactual and the theoretical policy-rule counterfactual should coincide if the DSGE model and the BSPVAR model agree on (i) the effects of a global risk aversion shock and (ii) the effects of policy (news) shocks and (iii) the approximation error arising from the fact that we only empirically identify a subset of all policy (news) shocks to enforce the counterfactual policy-rule and therefore have to rely on the "best Lucas-critique robust approximation" of McKay and Wolf (2023) is not too large. As shown in the main body of the paper the trinity model and the BPSVAR model largely agree on the effects of global risk aversion shocks, and therefore the disagreement has to arise either from differences in the estimated effects of policy shocks in the BPSVAR and the trinity model and/or the approximation error. Although by definition we cannot compute the approximation error for the empirical counterfactual as we do not have the true data-generating structural model at hand, we can do so for the trinity model. In Figure C.30 we plot the impulse responses to a global risk aversion shock in the trinity model (solid black line) alongside the true FX peg policy-rule counterfactual (dashed blue line) and the "best Lucas-critique robust approximation" (orange lines) under the assumption that an econometrician only has access to a limited number of identified policy shocks and impulse responses from the trinity model.[110] By focusing on the distance between the true FX peg policy-rule counterfactual and the "best Lucas-critique robust approximation" in Figure C.30 it becomes apparent that the econometrician can closely recover the true policy-rule counterfactual even when she only has access to a only two or three identified policy shocks as is the case in our application. We interpret this as tentative evidence that the differences between the empirical policy-rule counterfactual and the theoretical policy-rule counterfactual from the trinity model are not a result of the approximation error but rather arise from the differences in the estimated effects of policy shocks. This intuition is confirmed in Figure C.31, where we plot the impulse responses to a contemporaneous monetary policy shock from the BPSVAR and the trinity model normalized to bring about the same appreciation of the US-$ on average over the first year. Although never explicitly targeted the theoretical impulse responses from the trinity model qualitatively agree with the impulse responses from the BSPVAR. They even quantitatively agree for many variables

[110] We added policy (news) shocks to the baseline trinity model in order to compute the approximations to the true counterfactual based on impulse responses to those shocks.

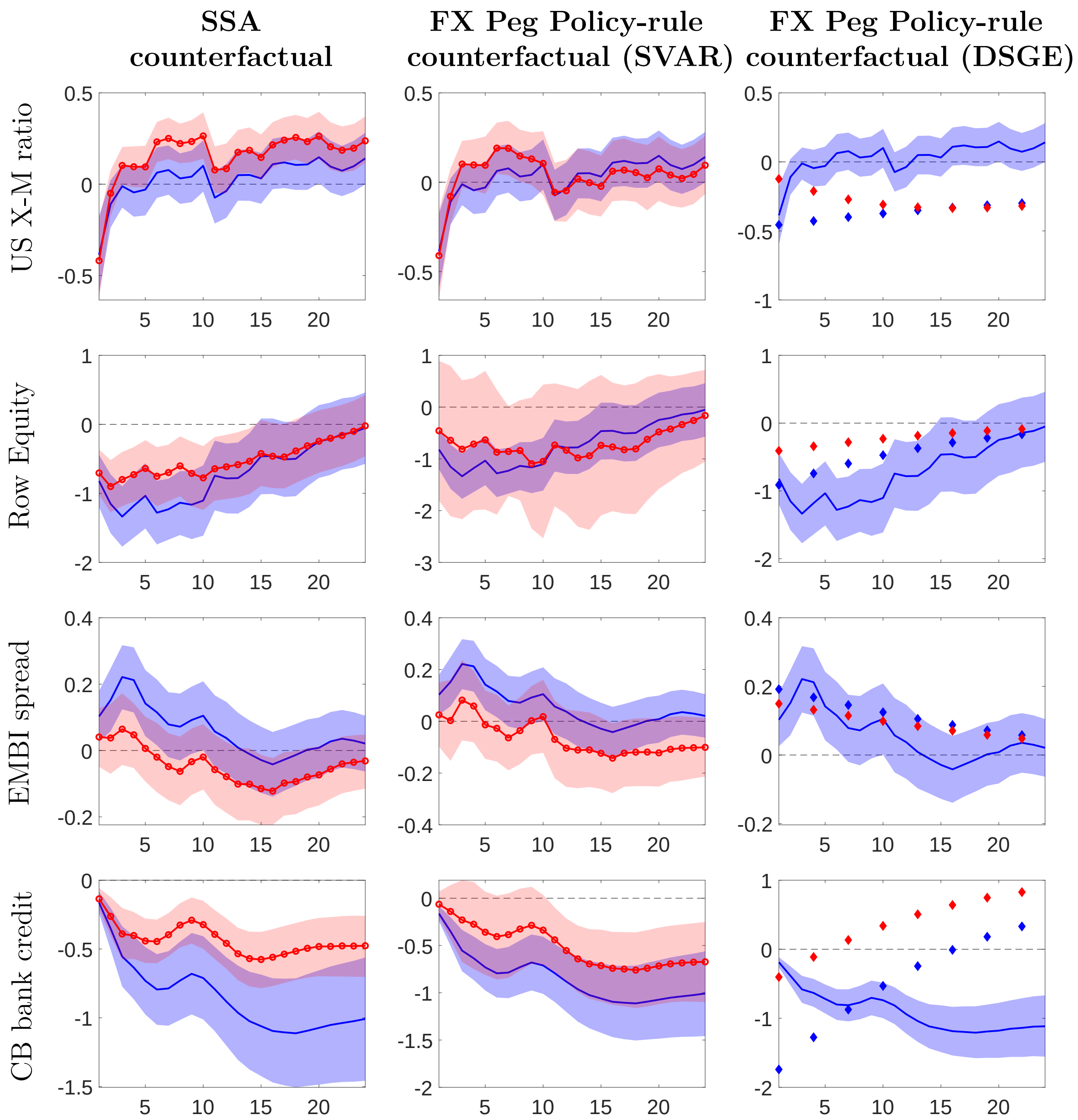
in the system but there are two crucial outliers: US output and the US policy rate. Figure C.31 shows that the trinity model underestimates the responsiveness of the exchange rate to changes in the policy rate. Leveraging the equivalence between a sequence of policy shocks and the change in the policy-rule this implies that to bring about the same level of depreciation needed to compensate for the appreciation caused by the global risk shock, the US central bank has to loosen much more (issue much stronger policy news shocks) in the model than in the data. In the theoretical model, this implies that not only is the contraction in RoW output strongly mitigated, but US output in fact rises.

Figure C.27: Baseline and counterfactual responses to a global risk shock



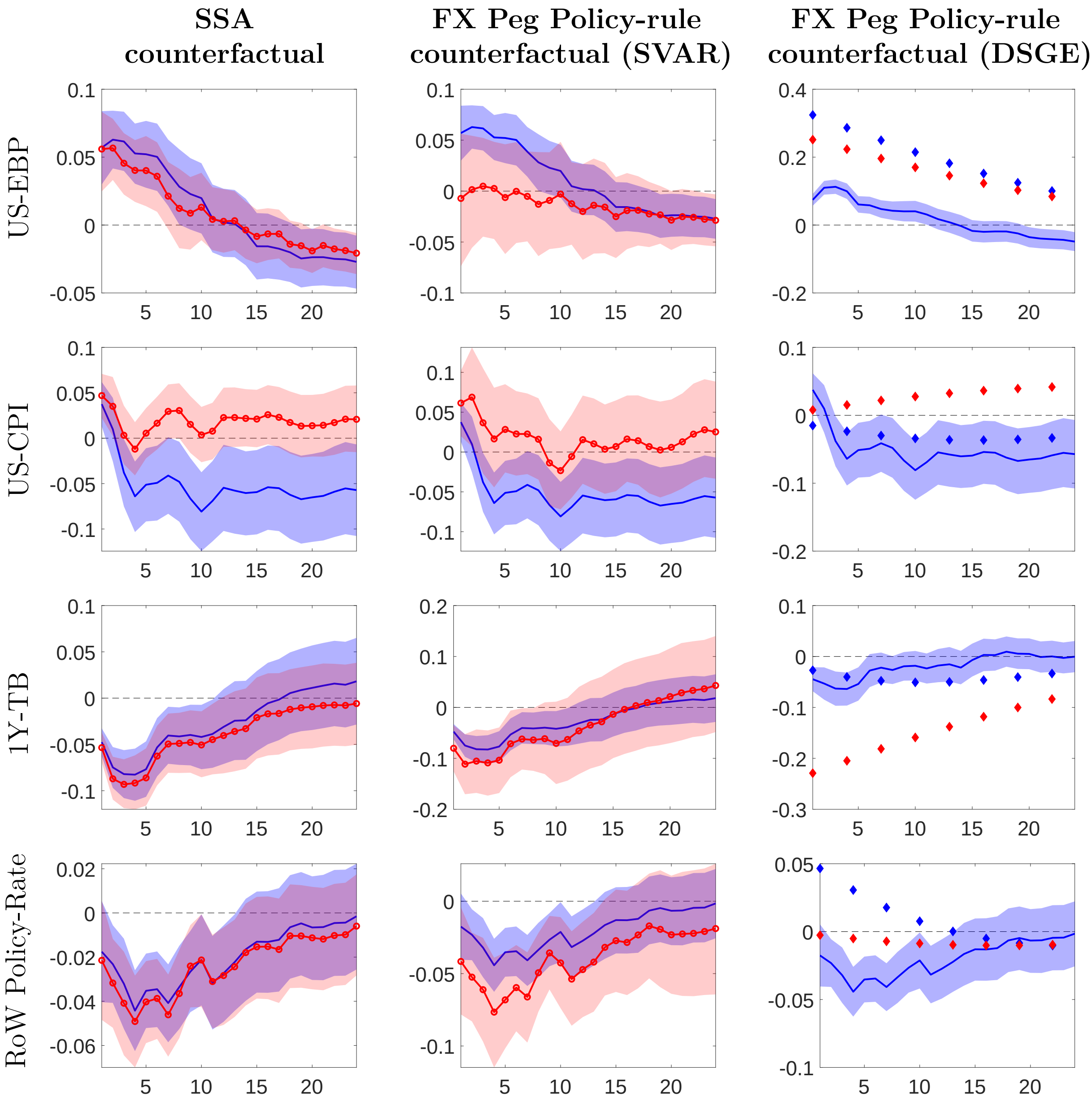
Note: The figure shows the baseline BPSVAR model (blue solid) and counterfactual (red circled) impulse responses to a global risk shock. SSA counterfactuals are shown in the first column, policy-rule counterfactuals in the second column, and the trinity-model counterfactuals in the third column. The red (blue) shaded areas represent 68% credible sets obtained from computing the counterfactual (impulse responses) for each draw from the posterior distribution. In the third column, the blue (red) diamonds depict the baseline (counterfactual) impulse responses to a global risk aversion shock in the trinity model without (with) an exchange rate peg enacted by the US central bank. We do not connect the dots depicting the counterfactual because the trinity model is calibrated to quarterly frequency while the BPSVAR model is estimated at the monthly frequency. The global risk aversion shock in the trinity model is scaled such that the average of the response of the dollar over the first year is the same as the response from the BPSVAR model. The real GDP (output) response in the trinity model is multiplied by 2.5 to make it comparable to the industrial production response from the BPSVAR model given that in the data the latter is 2.5 times more volatile than the former. In the Appendix we document that the BPSVAR model impulse response of S&P Global's US monthly GDP is indeed about 2.5 times smaller than for US industrial production (see Figure C.12), while their time profiles are rather similar.

Figure C.28: Baseline and counterfactual responses of trade and financial variables to a global risk shock



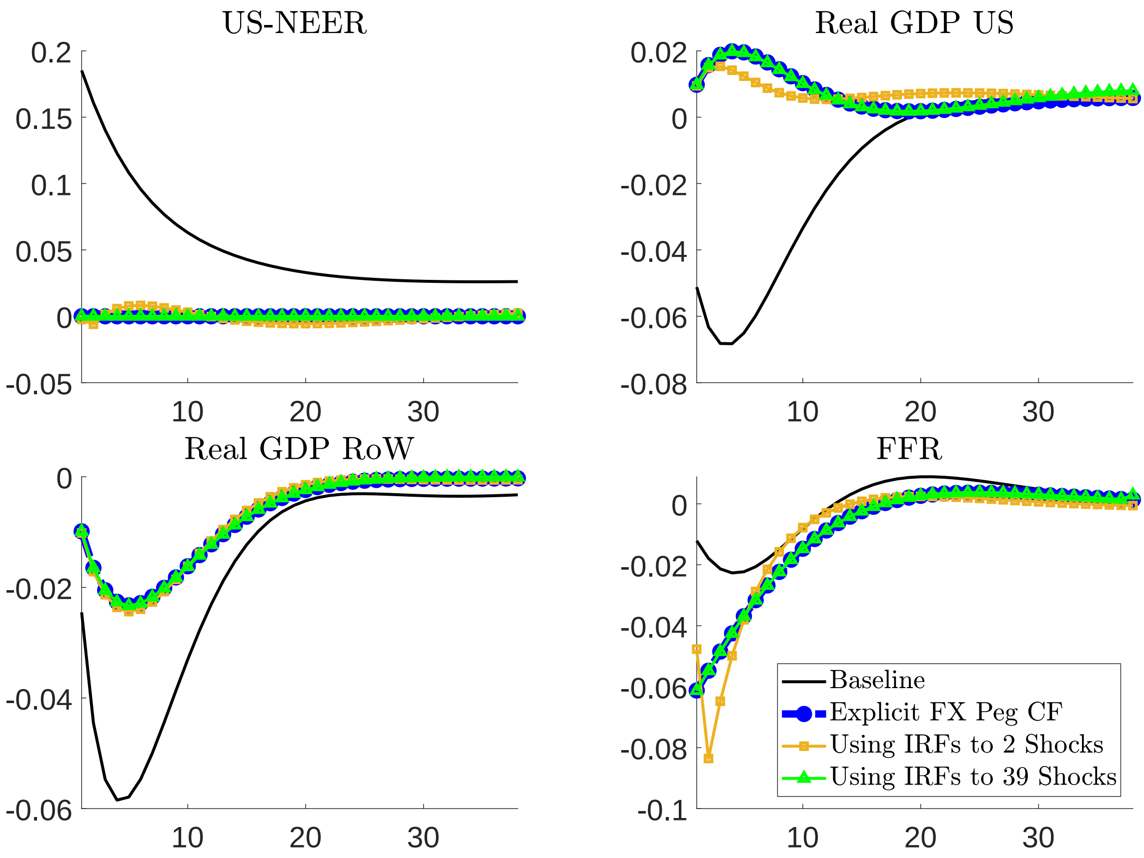
Note: See notes to Figure C.27. As the trinity model does not include an exact match for equity prices (the EMBI spread) we plot the response of the price of capital (RoW cross-border credit spread) instead. In the counterfactual structural model, the US policy rule is specified as exchange rate peg.

Figure C.29: Baseline and counterfactual responses of remaining BPSVAR model variables to a global risk shock



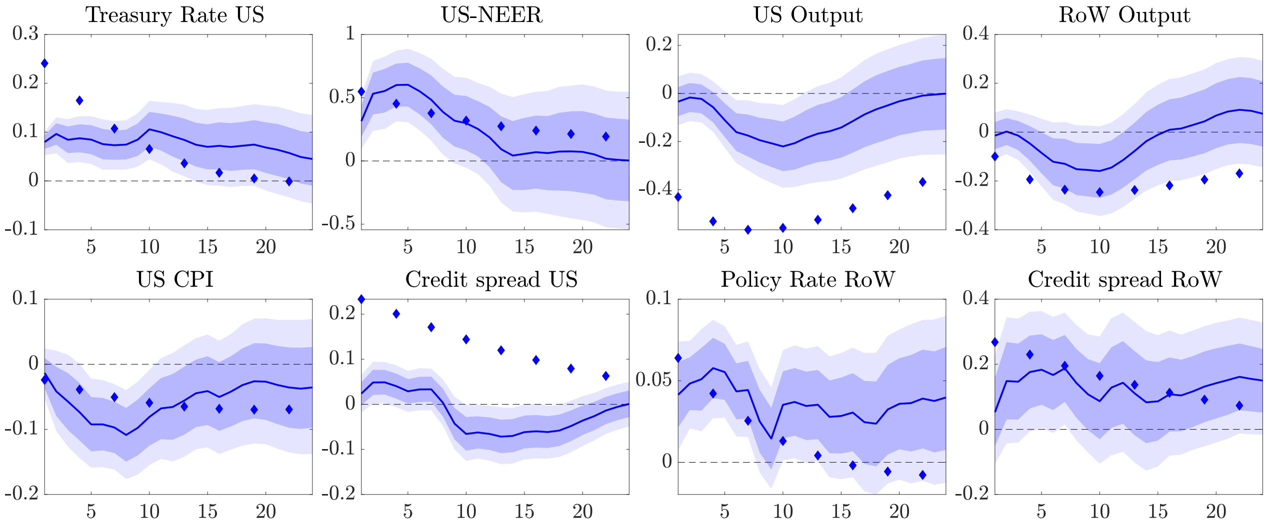
Note: See notes to 4. As the model of Georgiadis et al. (2023) does not include an exact counterpart US EBP we plot the responses of the US credit spread instead. While in the baseline specification of the BPSVAR we included the AE policy rate as our indicator for the RoW policy stance as weighted average of the policy rates of the "entire" RoW as computed in the Dallas Fed Global Economic Indicators data (Grossman et al. 2014) are extremely volatile volatility in the 1990s due to several crises involving major emerging market economies (EMEs), including Mexico, Brazil, Russia, Thailand, Indonesia, Malaysia, South Korea, Philippines, Argentina and Turkey. As in the DSGE model we want to capture the policy stance in the "entire" RoW the blue lines in the last row corresponds to results from a specification with a 'hybrid' RoW policy rate. : Due to the extreme values in the beginning of the sample we impute backwards from 2000 the levels of RoW policy rates in the AE policy rate. Given that the size of EMEs—especially of China—took off only after the late 1990s, this should introduce only mild distortions in the RoW aggregate series. Due to the extreme values in the beginning of the sample we impute backwards from 2000 the levels of RoW policy rates changes in the AE policy rates.

Figure C.30: Using US monetary policy (news) shocks to approximate the true FX Peg Policy Rule counterfactual in the Trinity Model



Note: This figure plots the impulse responses to a risk aversion shock from the trinity model (black solid line) alongside impulse responses to a risk shock under a change in the policy rule, where the US monetary authority in the trinity model stabilizes the US-$ (blue circled line). The green triangled line shows that this counterfactual can be perfectly recovered using a sequence of identified policy news shocks. The orange squared lines correspond to approximations of the true counterfactual assuming that the econometrician who applies the approach of McKay and Wolf (2023) only has impulse responses to two policy (news) shocks at hand as is the case in the empirical application in section 4.2 of the paper. In the trinity model the impulse responses to a global risk aversion shock under the counterfactual US$ dollar stabilizing policy rule can be closely recovered by combining the impulse responses to a risk shock with only a few impulse responses to policy shocks under the baseline policy rule.

Figure C.31: SVAR & DSGE impulse responses to a contemporaneous US monetary policy shock



Note: Horizontal axis measures time in months, vertical axis deviation from pre-shock level; size of shock is one standard deviation; blue solid line represents point-wise posterior mean and shaded areas 68%/90% equal-tailed, point-wise credible sets. We do not connect the dots depicting the counterfactual because the trinity model is calibrated to quarterly frequency while the BPSVAR model is estimated at the monthly frequency. The US monetary policy shock in the trinity model is scaled such that the average of the response of the dollar over the first year is the same as the response from the BPSVAR model. The real GDP (output) response in the trinity model is multiplied by 2.5 to make it comparable to the industrial production response from the BPSVAR model given that in the data the latter is 2.5 times more volatile than the former. In the Appendix we document that the BPSVAR model impulse response of S&P Global's US monthly GDP is indeed about 2.5 times smaller than for US industrial production (see Figure C.12), while their time profiles are rather similar.

Appendix D

Appendix for Chapter 4

D.1 Robustness of model prediction from partial DCP

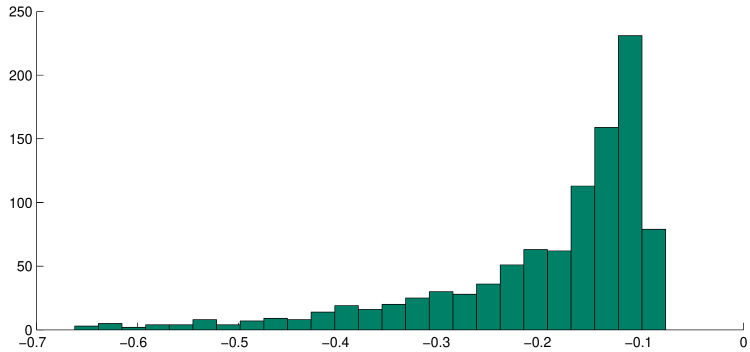
D.1.1 A Monte Carlo experiment exploring the sensitivity of the prediction from DCP to model parametrisation

One may wonder whether the prediction that the output spillovers from a positive US demand shock are negatively related to economies' non-US export-import DCP share differential only obtains under very specific parameterisations of the model. For example, the strength of expenditure switching in the face of US dollar exchange rate changes depends on the elasticity of substitution between foreign and domestically produced goods. In order to document that the prediction from partial DCP we derive is not specific to particular parameterisations of the model, we carry out a Monte Carlo experiment. In each replication, we draw values for a subset of the model parameters from a uniform distribution over a support than spans the range of values that are used in the literature. Specifically, we draw values for all economies for the inter-temporal elasticity of substitution $\sigma_c \in [0.5, 4]$, the demand elasticity for domestic and foreign final goods $\psi_f \in [0.75, 4]$, the demand elasticity for differentiated intermediates $\psi_i \in [1.1, 3]$, as well as the Calvo parameters for price stickiness $\theta_p \in [0.65, 0.85]$. Figure D.1 presents the distribution of the difference between EME output spillovers in positive and negative non-US export-import DCP pricing share scenarios across 1,000 replications of a Monte Carlo experiment. No parametrisation changes the prediction that output spillovers are larger in case of a negative export-import DCP pricing share differential qualitatively.

D.1.2 Intermediate inputs trade

Georgiadis et al. (2019) set up a structural model and show that exchange rate pass-through to important prices is reduced even in the case of DCP if an economy's trading partners are integrated in global value chains. Against this background, the question arises whether accounting for intermediate inputs trade in the model we consider in this paper would

Figure D.1: Distribution of difference between EME output spillovers in positive and negative non-US export-import DCP pricing share parameterisations across replications of a Monte Carlo experiment



Difference in EME real GDP impact response for $\Delta_E > 0$ and $\Delta_E < 0$

Note: The figure presents the distribution of the difference between the model-implied output spillover from a positive US demand shock in EMEs across the cases of 100% and 0% of EME exporters exhibiting subject to DCP. In each replication of the Monte Carlo experiment we draw values for all economies for the intertemporal elasticity of substitution $\sigma_c \in [0.5, 4]$, the demand elasticity for domestic and foreign final goods $\psi_f \in [0.75, 4]$, the demand elasticity for differentiated intermediates $\psi_i \in [1.1, 3]$, as well as the Calvo parameters for price stickiness $\theta_p \in [0.65, 0.85]$.

overturn qualitatively the prediction of partial DCP regarding the asymmetry that worsens the bilateral net exports of EME vis-à-vis RoW in the face of US dollar appreciation. Specifically, if the price of the intermediate input good bundle rises this could partly be passed through into prices mitigating the differences in competitiveness stemming from the differences in the export-import DCP pricing share differentials between EME and RoW. While this effect is indeed operative in a model with trade in intermediate goods used as inputs in production instead of only for consumption, given that prices are sticky in the short run, it does not qualitatively offset the role of export-import DCP pricing share differentials for the spillovers from shocks that appreciate the US dollar that we highlight in the baseline version of the model. This is documented in Figure D.5, which shows the spillovers from a positive US demand shock in a version of the model with trade in intermediate goods used for inputs in production.[111]

D.1.3 Capital, sticky wages and financial frictions

We extend our baseline model in that each economy is inhabited by seven rather than five types of agents: Households, financial intermediaries, intermediate goods producers, final good bundlers, capital good producers and a monetary authority. Households trade international bonds, provide deposits to local intermediaries and provide labor monopolistically while being subject to nominal rigidities in wage setting. The financial intermediaries are

[111]Another aspect to consider is that it is reasonable to assume that there is also some degree of home bias in the use of intermediate inputs. An initial rise in domestic production in response to a foreign shock will thereby cause the demand for domestic intermediates to rise by more than for foreign intermediates for a given level of prices. Given the fact that differences in competitiveness arise in the short run mainly due to differences in the export-import DCP pricing share differentials, this even slightly amplifies the effect we highlight in the baseline version of the model.

modelled as in Gertler and Karadi (2011), and thus intermediate funds between firms and households while being subject to a leverage constraint.[112] Figure D.6 documents that also in this more elaborate version of the model the predictions from partial DCP obtain.

D.1.4 Hedging, non-constant demand elasticity and local distribution services

Finally, we briefly discuss the implications of additional model features that have been put forth in the literature on export pricing and exchange rate pass-through, namely non-constant elasticity of demand (which implies complementarities in price setting) and local distribution services. Specifically, we argue that none of these features changes the predictions from partial DCP regarding the asymmetry that worsens the bilateral net exports of EMEs vis-à-vis the RoW we highlight in the baseline version of our model qualitatively. In order to suggest that the presence of the presence of each additional model feature does not qualitatively change the predictions from partial DCP, we will argue that expenditure switching in the face of an appreciation of the US dollar remains stronger in the economy with the larger export-import DCP pricing share differential. Hence, we focus on the scenario with a positive (negative) export-import DCP pricing share differential in EMEs (the RoW), i.e. when EMEs price 100% of their exports to the RoW in US dollar while the RoW prices only 50% of its exports to EMEs in US dollar.

Hedging

In the data, anecdotal evidence suggests that some importers hedge against exchange rate risk. Conceptually, hedging contracts can be designed in various ways, and the specifics have different implications regarding the effects of US dollar exchange rate variation for net exports under DCP in non-US economies. For example, hedging contracts can be designed between different households in the same economy; between households across different economies, both across the dominant-currency issuing economy and one economy in the rest of the world and across different economies in the rest of the world. In general, hedging imports against US dollar exchange rate risk does not change the qualitative prediction of partial DCP. First, because hedging is costly, it will in general not be optimal to hedge exchange rate risk completely. Second, while hedging removes the income effect implied by changes in the US dollar exchange rate, it does not remove the substitution effect. As a result, even hedged households in the rest of the world will switch away from imported goods towards domestically produced goods in response to US dollar appreciation in case of DCP.[113] Moreover, if hedging was such that it did annihilate expenditure switching

[112] All model equations are explicitly derived in an appendix available from the authors.

[113] As firms are owned by households, hedging nominal export revenues against US dollar exchange rate variation is akin to households insuring against income effects, at least for given export quantities, and hence does not annihilate substitution effects.

completely, then export-import DCP share differentials would just be inconsequential for the output spillovers from US dollar appreciation (unless of course hedging would increase the output spillovers, which is however the exact opposite of what hedging contracts are designed to achieve). Our empirical analysis can therefore be understood as additionally testing the hypothesis that not all of imports subject to US dollar DCP are hedged and/or that hedging is not complete.

Non-constant demand elasticity

Under a non-constant demand elasticity, the elasticity of demand faced by a producer depends on the deviation of its price from that of its competitors. In particular, Gust et al. (2010) study an environment with PCP and show that with non-constant demand elasticity exporters' prices in the importers' currency are less responsive to exchange rate changes, which exporters achieve by adjusting their mark-ups correspondingly. The goal of the paper by Gust et al. (2010) is to show that non-constant elasticity of demand is sufficient to generate incomplete exchange rate pass-through, independently of exogenously imposed LCP and sticky prices.

To facilitate understanding how non-constant elasticity of demand would affect the implications of partial DCP in our model, consider the case of entirely fixed prices first. Under non-constant elasticity of demand and fixed prices, the appreciation of the US dollar which raises the local-currency price of imports in the RoW (and in EMEs) elicits a *stronger* expenditure switching away from imports towards domestically produced goods than under constant elasticity of demand. Therefore, and because all EME exports to the RoW are priced in US dollar while only 50% of RoW exports to EME are priced in US dollar, the loss of competitiveness is stronger for EME exports to the RoW than vice versa. As a result, in the short term with fixed prices non-constant elasticity of demand amplifies the asymmetry that worsens the bilateral net exports of EME vis-à-vis RoW we highlight in the baseline version of our model.[114]

In the medium term when exporters start to adjust their prices, EME exporters will tend to lower the US dollar price of their exports to RoW in order to limit deviation of their RoW currency price from that of RoW competitors, hence dampening—but not muting completely—the amplification of expenditure switching due to non-constant elasticity of demand in RoW. Therefore, even in the medium term non-constant elasticity of demand amplifies quantitatively—even though to a somewhat smaller degree—the asymmetry in expenditure switching across EME with a positive and RoW with a negative export-import DCP pricing share differential.

[114]Notice moreover that non-constant elasticity of demand as introduced in Gust et al. (2010) is irrelevant in a first-order approximation of the model. In the non-linear model, the magnitude of the effect of non-constant demand elasticity depends on the curvature of the demand function. Specifically, only if the curvature is large will a small shock trigger large changes in the elasticity of substitution.

Local distribution services

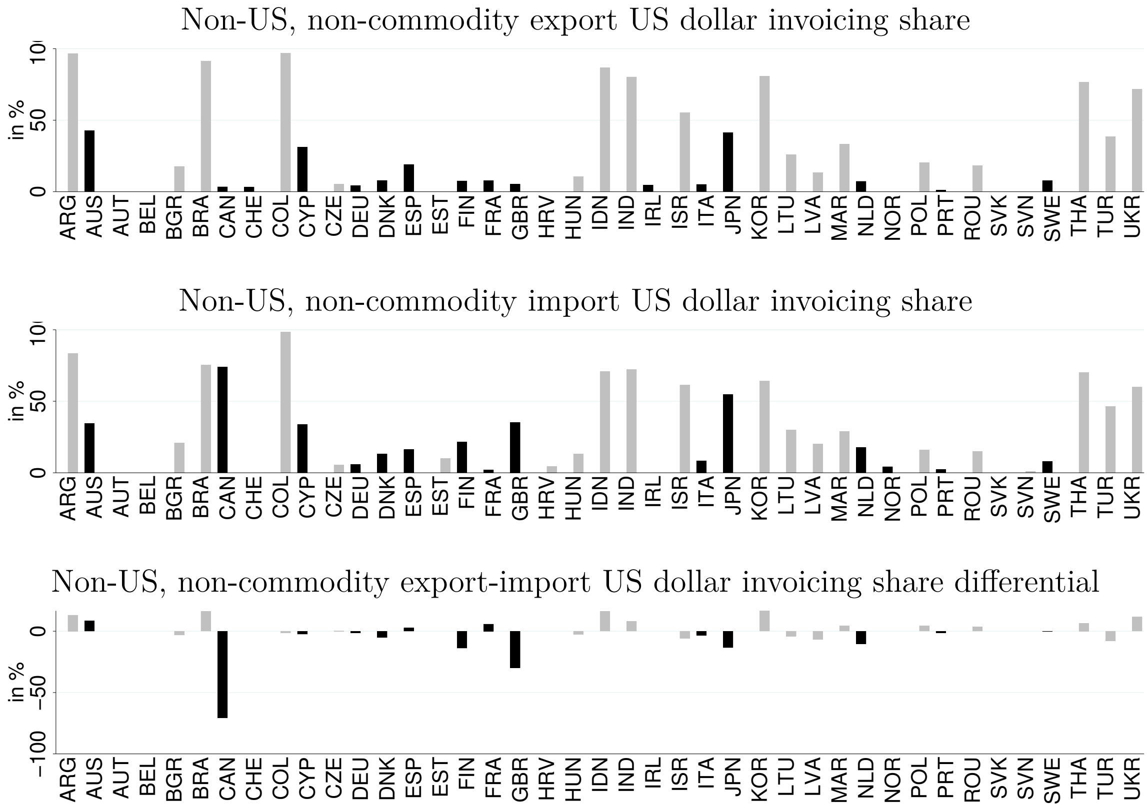
Under local distribution services local labour is required in order to distribute imported (and domestic) tradable goods to consumers. Corsetti and Dedola (2005) consider an environment with PCP and show that the presence of local distribution services gives rise to different elasticities of demand across countries and hence deviations of the law-of-one-price by limiting the exchange rate pass-through to in particular local consumer prices.

To facilitate understanding how local distribution services would affect the implications of the baseline version of our model we highlight, consider again the case of fixed prices first. While the appreciation of the US dollar still raises the RoW currency price of imports from EME, the rise is smaller than in a model without local distribution costs. This is because the price of an imported good from EME that is eventually distributed in RoW is an average of the RoW currency price of the imports from EME and local labour in RoW. Hence, the use of local labor in distributing imports from EME to RoW consumers dampens the rise in RoW consumer prices and hence expenditure switching. However, the mechanism applies symmetrically in EME as well: The use of local labor in distributing imports from the RoW to EME consumers dampens the rise in EME consumer prices and hence expenditure switching. And since the rise in the local currency price of imports resulting from the appreciation of the US dollar is stronger in RoW due to the higher DCP import share than in EME, in the short term with fixed prices the asymmetry that worsens the bilateral net exports of EME vis-à-vis RoW we highlight in the baseline version of the model continues to prevail even in the presence of local distribution services.

Finally, because even in the presence of local distribution services marginal costs of EME exporters do not change more than in the baseline version of our model, even when prices adjust in the medium term EME exporters will not lower their prices in a way that would completely annihilate expenditure switching in RoW.

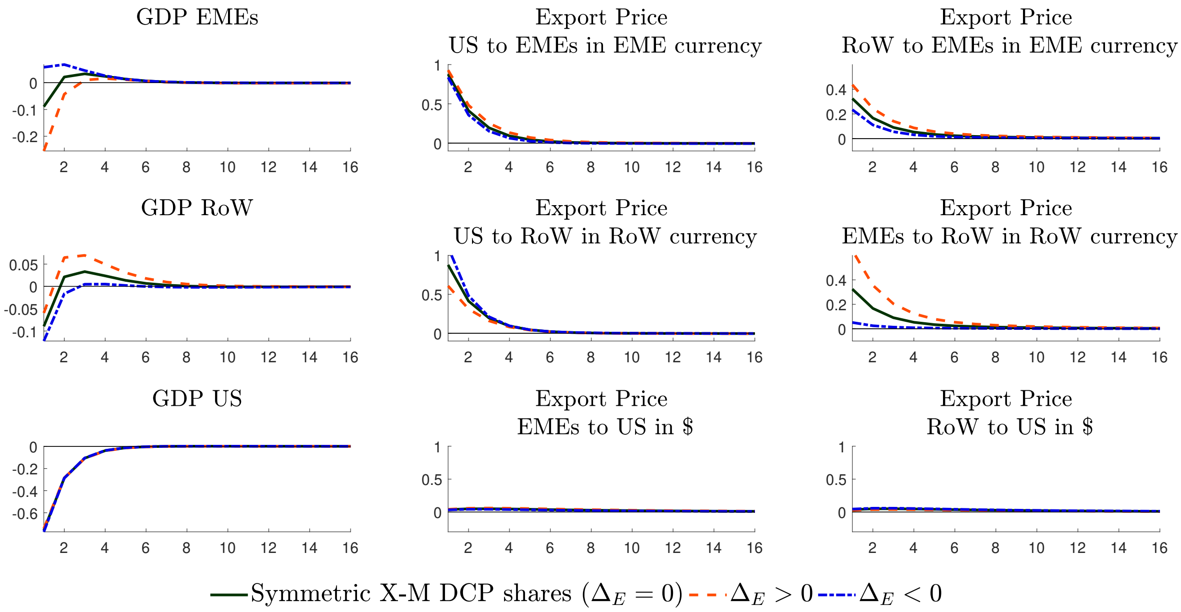
D.2 Additional Figures

Figure D.2: Non-commodity trade US dollar invoicing share differential



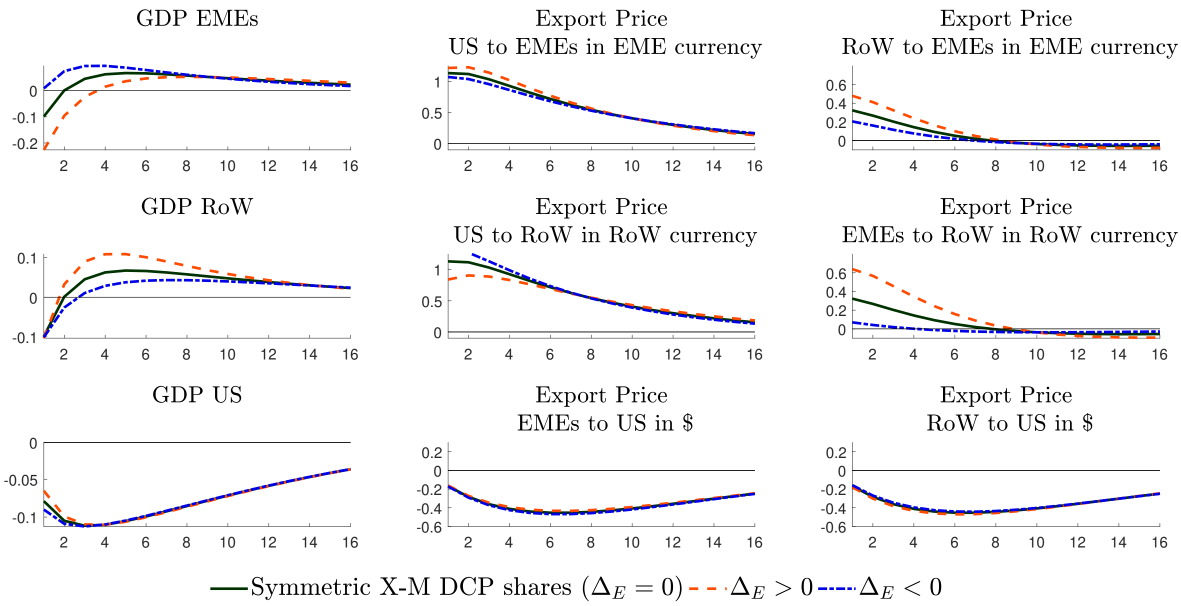
Note: The figure presents non-US, non-commodity trade, export-import US dollar invoicing share differentials. See Footnote 81 in the main text for a description of the construction of this variable.

Figure D.3: Dynamic effects of a contractionary US monetary policy shock under partial DCP



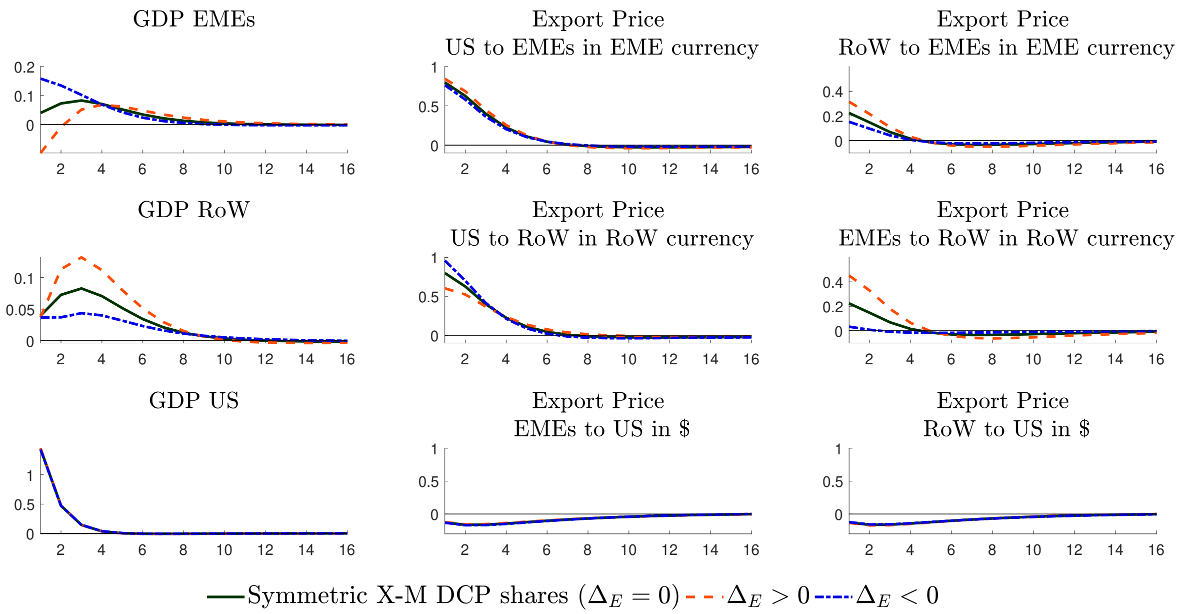
Note: The figure presents the dynamic effects of a contractionary US monetary policy shock on real GDP and bilateral export prices denoted in the currency of the importer for the case of partial DCP. See also the note to Figure 4.3.

Figure D.4: Dynamic effects of a UIP shock under partial DCP



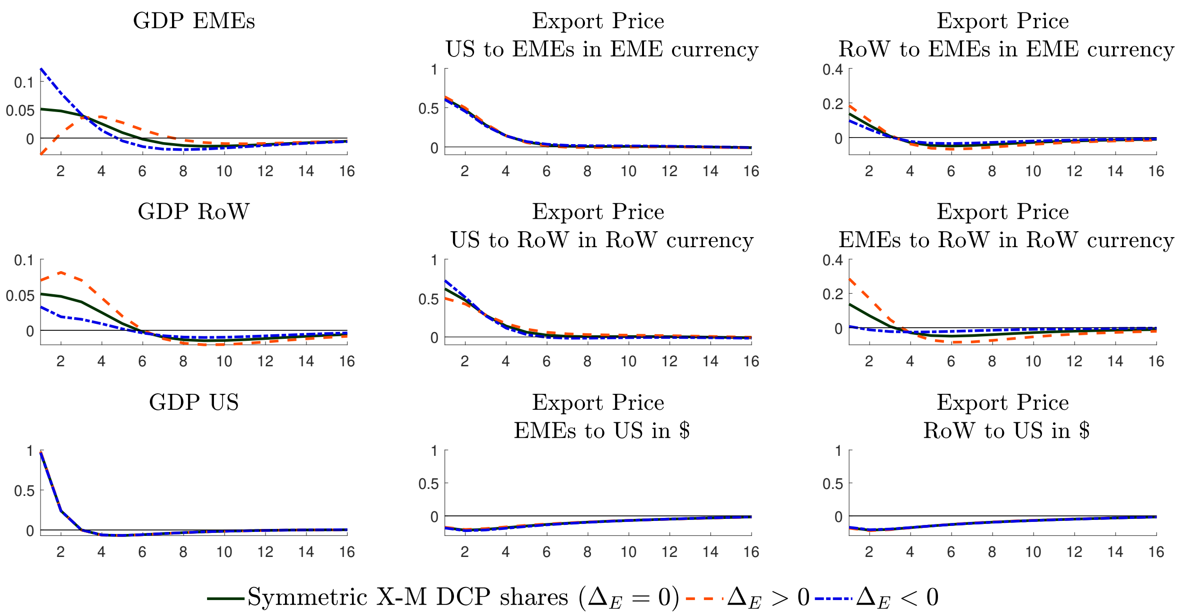
Note: The figure presents the dynamic effects of a UIP shock on real GDP and bilateral export prices denoted in the currency of the importer for the case of partial DCP. See also the note to Figure 4.3.

Figure D.5: Dynamic effects of an US demand shock under partial DCP with trade in intermediate inputs to production



Note: The figure presents the dynamic effects of a positive US demand shock on real GDP and bilateral export prices denoted in the currency of the importer for the case of partial DCP using a model version in which foreign and domestic intermediate inputs are used as means of production (steady state share of intermediate inputs = 0.5). See also the note to Figure 4.3.

Figure D.6: Dynamic effects of an US demand shock under partial DCP in a more elaborate version of the model

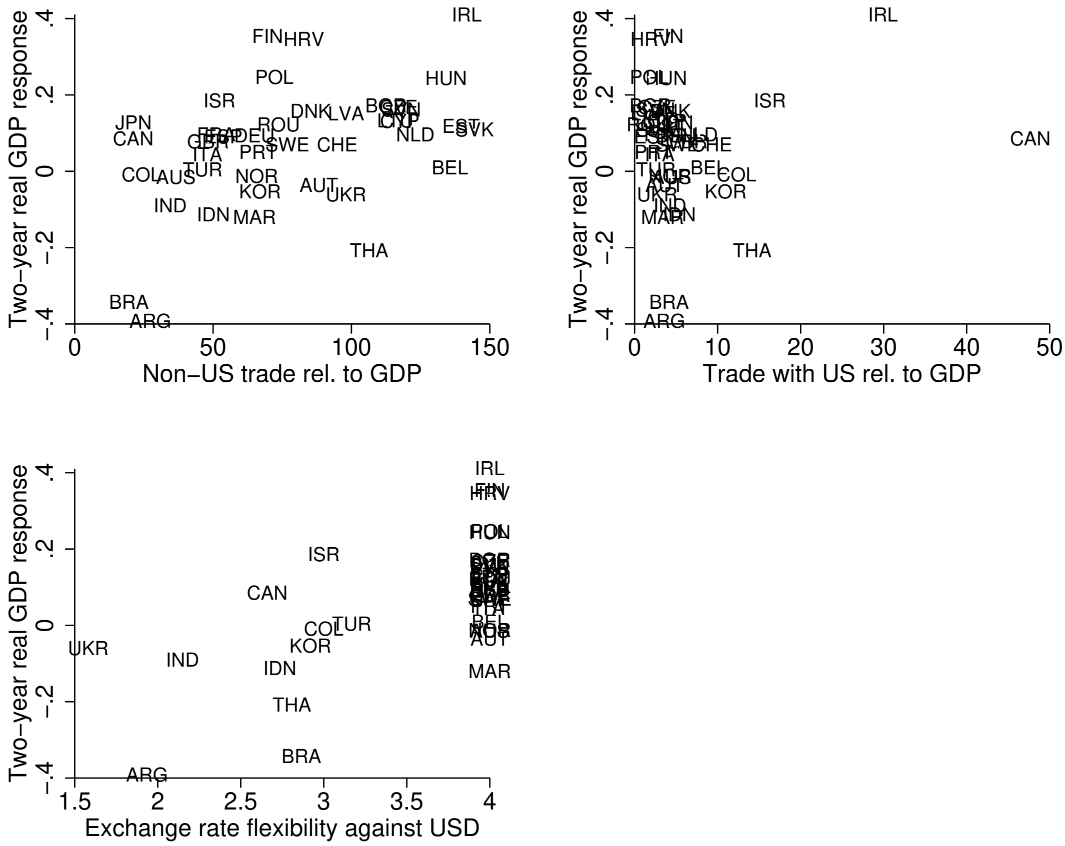


Note: The figure presents the dynamic effects of a positive US demand shock on real GDP and bilateral export prices denoted in the currency of the importer for the case of partial DCP using a more elaborate version of the baseline model with capital, sticky wages and financial frictions. See also the note to Figure 4.3.

Figure D.7: Estimated GDP responses to a positive US demand shock, expanded country sample

Note: The figure shows the averages over eight quarters of the point estimates of the real GDP effects of a positive US demand shock. See also the note to Figure 4.5.

Figure D.8: Relationship between spillovers from a positive US demand shock and country characteristics



Note: The figure depicts the unconditional correlation between the estimated spillovers from a positive US demand shock and control variables included in the regression in Equation (4.27).

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Erklärungen

Erklärung gem. §4 PO

Hiermit erkläre ich, dass ich mich noch keinem Promotionsverfahren unterzogen oder um Zulassung zu einem solchen beworben habe, und die Dissertation in der gleichen oder einer anderen Fassung bzw. Überarbeitung einer anderen Fakultät, einem Prüfungsausschuss oder einem Fachvertreter an einer anderen Hochschule nicht bereits zur Überprüfung vorgelegen hat.

Berlin, den 05. April 2024

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Ben Alexander Schumann

Erklärung gem. §10 Abs. 3

Hiermit erkläre ich, dass ich für die Dissertation folgende Hilfsmittel und Hilfen verwendet habe:

- MATLAB
- **R**-Studio
- Excel
- STATA
- LaTeX
- DeepL
- Siehe References, für die verwendete Literatur

Auf dieser Grundlage habe ich die Arbeit selbstständig verfasst.

Berlin, den 05. April 2024

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Ben Schumann