Essays on Unconventional Monetary Policy, Inflation Expectations, and Commodity Prices

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Erklärung zur Ko-Autorenschaft

Diese Dissertation besteht aus vier Forschungsspapieren, von denen zwei in Zusammenarbeit mit jeweils zwei Koautoren und eines in Zusammenarbeit mit einem Koautor entstanden sind. Der Eigenanteil an Konzeption, Durchführung und Berichtsabfassung lässt sich folgendermaßen zusammenfassen:

- Pablo Anaya, Michael Hachula und Christian J. Offermanns:

  Spillovers of U.S. Unconventional Monetary Policy to Emerging Markets: The Role of Capital Flows

  Eine ältere Version dieses Kapitels ist als Freie Universität Berlin Discussion Paper 35/2015 erschienen. Veröffentlicht in der Fachzeitschrift *Journal of International Money and Finance* (Vol. 73, Part B, May 2017, pp. 275-295, https://doi.org/10.1016/j.jimonfin.2017.02.008).

  Eigenanteil: 33 Prozent

- Michael Hachula und Malte Rieth:

  Identifying Speculative Demand Shocks in Commodity Futures Markets through Changes in Volatility

  Eine Version dieses Kapitels ist als DIW Berlin Discussion Paper 1646 (2017) erschienen.

  Eigenanteil: 50 Prozent

- Michael Hachula:

  The Anchoring of Inflation Expectations: Evidence from SVARs identified with Macroeconomic News Announcements

  Eigenanteil: 100 Prozent

- Michael Hachula, Michele Piffer und Malte Rieth:

  Unconventional Monetary Policy, Fiscal Side Effects and Euro Area (Im)balances

  Eine Version dieses Kapitels ist als DIW Berlin Discussion Paper 1596 (2016) erschienen.

  Eigenanteil: 33 Prozent

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List of Abbreviations

AEs Advanced Economies
AIC Akaike Information Criterion
GARCH Generalized autoregressive conditional heteroscedasticity
BC "Business Cycle" Model
BEIR Break-even inflation rate
BIC Bayesian Information Criterion
CBOE Chicago Board Options Exchange
CDS Credit default swaps
CFTC Commodities Futures Trading Commission
CFTC SCOT reports CFTC Supplemental Commitments of Traders reports
CPI Consumer price index
CPOs Commodity pool operators
CTAs Commodity trading advisors
DSGE Dynamic stochastic general equilibrium
EC European Commission
ECB European Central Bank
EMEs Emerging market economies
EMU European Monetary Union
Euribor Euro InterBank Offered Rate
FC "Financial Conditions" Model
FED Federal Reserve Bank
FEVD Forecast error variance decomposition
FRED Federal Reserve Bank of St.Louis database
FX Foreign exchange rate
GDP Gross domestic product
GSCI Goldman Sachs Commoditiy Index

GVAR Global vector autoregression
GVECM Global vector error-correction model
i.e. id est
IFS International Financial Statistics database
IMF International Monetary Fund
IP Industrial Production
IRF Impulse response function
LR Likelihood ratio
LTRO Long Term Refinancing Operations
MBS Mortgage backed securities
MFI Monetary financial institution
MNAs Macroeconomic news announcements
OECD Organisation for Economic Co-operation and Development
OMT Outright Monetary Transactions
PMI Purchasing Managers Index
QE Quantitative Easing
SMP Securities Markets Programme
SVAR Structural vector autoregression
TIC Treasury International Capital Reporting System
UMP Unconventional monetary policy
UK United Kingdom
US United States of America
USD US Dollars
VAR Vector autoregression
VARX Vector autoregression with exogenous regressors
VIX CBOE volatility index
VSTOXX Eurostoxx50 volatility index
ZEW Centre for European Economic Research

Overview

This thesis contains four independent papers which empirically address research questions from macroeconomics and finance. A particular focus of the thesis lies on the effects of unconventional monetary policy (UMP). Following the outbreak of the global financial crisis, and with conventional interest rate policy constrained by the zero lower bound, central banks started using unconventional measures like asset purchases or credit easing. The new tools spurred an intense public and academic debate about their effectiveness in stimulating output and price growth, and potential side-effects associated with them.

Two of the papers included in the thesis contribute to this debate. One paper is concerned with spillovers of US UMP to other economies. Specifically, this paper analyzes whether US UMP affects capital flows into emerging market economies and how it impacts on financial and economic conditions in these countries. The other paper focuses on the effectiveness and potential side-effects of UMP in the euro area. It studies whether UMP by the European Central Bank can stimulate price and output growth in the monetary union and its largest member states. Moreover, the paper assesses how fiscal policy reacts to UMP shocks and whether UMP accentuates or attenuates internal trade imbalances within the euro area.

A third paper investigates a topic that is important for the conduct of monetary policy, namely the anchoring of inflation expectations. According to models with forward looking Philips curves, like standard New Keynesian models (see Galí, 2015), inflation expectations are an important determinant of future inflation and central banks, therefore, aim at keeping them well anchored at an inflation target. New empirical evidence on the anchoring of inflation expectations in the US is provided in this part of the thesis.

A fourth paper is concerned with the financialization of commodity markets. The increased presence of financial investors in the markets coincided with drastic boom-bust cycles in commodity prices, which spurred an intense debate about the role of investors in such cycles. The discussion revolves around whether investors are responsible for the large price swings and, more generally, whether they drive prices away from fundamentals. The paper proposes and applies a new strategy to empirically assess the impact of financial speculation on commodity futures returns.

Methodically, in all four papers structural vector autoregressive models (SVARs) are employed to answer the research questions. SVARs have become one of the most widely used approaches for empirical researchers in macroeconomics and finance since the seminal work of Sims (1980). The goal of a structural VAR analysis is to uncover and quantify causal relationships in the data (compare Kilian and Lütkepohl, 2017). In particular, the researcher aims at identifying uncorrelated and economically interpretable structural shocks and at estimating their causal effects on the variables modeled in the VAR.

Identification of economically meaningful shocks is the main challenge in applying SVAR models for econometric analyses. Technically, identifying the model corresponds to decomposing the correlated errors of the reduced-form VAR model, which can be estimated, into the unobservable structural shocks of interest. Different approaches to achieve identification have been proposed and applied since Sims (1980). Most prominently, short-run exclusion restrictions are used which assume that the immediate impact of some structural shocks on selected endogenous variables is zero. Exclusion restrictions can also be placed on the long-run impact of structural shocks on variables within the system. Alternatively, a set of admissible SVAR models can be identified based on sign or shape restrictions on the estimated impulse response functions, or based on magnitude restrictions on the structural parameters of the model or on functions of them.

In this thesis, more recently established methods for identification are employed. One paper combines exclusion and sign restrictions (see Mountford and Uhlig, 2009, Baumeister and Benati, 2013, Arias et al., 2014), another paper uses the heteroskedasticity present in the data (see Sentana and Fiorentini, 2001, Rigobon, 2003, Lanne and Lütkepohl, 2008), while the two remaining papers employ external information from proxy variables (also referred to as external instruments, see Stock and Watson, 2012, Olea et al., 2013, Mertens and Ravn, 2013). The four individual papers constitute the four chapters of this dissertation. Their main results and contributions to the literature are briefly summarized in the following:

- ***Chapter 1:*** *Spillovers of US Unconventional Monetary Policy to Emerging Markets: The Role of Capital Flows*

  The first chapter is based on a joint paper with Pablo Anaya and Christian Offermanns. We employ a structural global VAR model to analyze whether US UMP shocks have an impact on financial and economic conditions in emerging market economies (EMEs), and whether international capital flows are an important channel of shock transmission. The UMP shocks are identified based on changes in the central bank's balance sheet, which is used as a policy indicator. A combination of zero and sign restrictions on their impact allows to disentangle them from other shocks to which the central bank responds by enlarging (or reducing) the balance sheet, like financial risk shocks. We find that an expansionary policy shock significantly increases portfolio flows from the US to EMEs for almost two quarters, accompanied by a persistent movement in real and financial variables in recipient countries. Moreover, EMEs on average respond to the shock with an easing of their own monetary policy stance. The findings appear to be independent of heterogeneous country characteristics like the underlying exchange rate arrangement.

  Our study contributes to the literature on international spillovers of US UMP by employing a global VAR approach that takes interactions between economic variables both between countries and over time into account. Complementing evidence from event studies that detect immediate capital reallocation due to UMP (see, for instance, Fratzscher et al., 2016b), we show that UMP shocks have persistent effects on capital flows, and that these flows are a channel of shock transmission. The paper also relates to studies that find conventional US monetary policy to be an important driver of economic conditions worldwide (see, for instance, Miranda-Agrippino and Rey, 2015), whereas advanced economies with flexible exchange rates are not insulated from the policy (see Passari and Rey, 2015). We show that US UMP is a driver of economic conditions in EMEs, independent of underlying exchange rate arrangements.

- ***Chapter 2:*** *Identifying Speculative Demand Shocks in Commodity Futures Markets through Changes in Volatility*

  The second chapter stems from joint work with Malte Rieth and studies the effects of financial speculation on agricultural commodity futures returns. It uses a VAR model and publicly available weekly positions data from the US Commodity Futures Trading Commission, aggregated by trader groups. To identify exogenous variation in positions of the two most important groups of speculators, index investors and hedge funds, we exploit the heteroskedasticity in the data. We apply this agnostic identification approach as zero restrictions are difficult to defend when working with weekly financial market data, while sign restrictions are not helpful in disentangling the shocks of interest as they should theoretically all exhibit the same sign pattern. The results suggest that idiosyncratic shocks to (net) long demand of both index investors and hedge funds increase futures returns. They further indicate that these shocks are a relevant driver of returns, especially during periods of high speculative demand volatility. Overall, fundamental determinants like demand and supply nevertheless prove to be the main drivers of agricultural commodity prices.

  The paper contributes to a literature that empirically analyzes how financial investment affects commodity futures prices. The key challenge in this literature is to isolate variation in investors' positions due to trades actually initiated by them from variation due to trades initiated by other market participants. Two recent studies address this issue with highly disaggregated and partly private data and detect significant effects of financial investments on commodity futures prices (see Cheng et al., 2015, Henderson et al., 2015). Our findings with publicly available data support their results and provide new evidence on the importance of financial investment for price developments. Moreover, the approach employed in our study has the advantage that it can be replicated and readily updated in the future.

- ***Chapter 3:*** *The Anchoring of Inflation Expectations: Evidence from SVARs identified with Macroeconomic News Announcements*

  The third chapter studies the anchoring of inflation expectations in the US. Well-anchored long-term inflation expectations should not respond to news about short-run macroeconomic developments if the central bank is fully credible and has an implicit or explicit inflation target. The chapter analyzes expectations anchoring in a structural VAR model with daily expectations data based on financial market variables. Specifically, I identify a macro news shock with a proxy variables approach using macroeconomic news announcements (MNAs) as an external instrument. This ensures that the latent shock will be correlated with observable information on macroeconomic surprises, namely the MNAs. Then I study the impact of the shock on long-term expectations. I find that macro news shocks significantly affect expectations in the short run, explaining more than 10 % of their forecast error variance. Their effects, however, fade out eventually, indicating an anchoring of expectations in the long run. Further, a sensitivity analysis suggest that the impact of the shocks on expectations is stronger after the financial crisis than before.

The contribution of the chapter is as follows. The anchoring of expectations is usually evaluated by studying their immediate response to MNAs on announcement days (see, Gürkaynak et al., 2010b, and subsequent studies). Nautz et al. (2016) introduce a SVAR approach to study the dynamic response of expectations to macro news shocks, which they identify with a statistical approach. This chapter, in contrast, proposes a strategy which assures that the identified shock is correlated with observable information on macroeconomic surprises by construction. I show that the results of Nautz et al. (2016) are qualitatively robust to this alteration and that after the financial crisis the dynamic effects of macro news shocks on inflation expectations are more pronounced.

- ***Chapter 4:*** *Unconventional Monetary Policy, Fiscal Side Effects and Euro Area (Im)balances*

  The fourth chapter, based on joint work with Michele Piffer and Malte Rieth, studies the macroeconomic effects of unconventional monetary policy in the euro area using SVARs, identified with an external instrument. The instrument is the common unexpected variation in sovereign bond spreads with different maturities of several euro area countries against Germany on announcement days of UMP measures. Using this high frequency data to identify the policy shock has the advantage of capturing the effects of monetary interventions without restricting them to their implementation, i.e. also accounting for announcement effects. We find that expansionary monetary surprises, which reduce public interest rates, also lower private interest rates and financial market uncertainty, and are effective at increasing output, consumer prices, and inflation expectations. Our results further indicate that the shocks lead to a rise in primary fiscal expenditures in the euro area as a whole and important member countries. The fiscal reaction, however, appears to be heterogenous across countries and also output and prices respond differently. This heterogeneity, in turn, is associated with a divergence of relative prices and a widening of existing trade imbalances within the union.

The contribution of the chapter is twofold. First, we provide new empirical evidence on the macroeconomic effectiveness of unconventional monetary policy in the euro area in a VAR model using an identification strategy based on high frequency data (compare Gertler and Karadi, 2015, for US conventional monetary policy, and Rogers et al., 2016, for US UMP). Second, this is - to the best of my knowledge - the first paper that empirically studies the reaction of fiscal policy and internal imbalances within the union to (unconventional) monetary policy interventions by the European Central Bank.

Überblick

Diese Dissertation besteht aus vier unabhängigen wissenschaftlichen Aufsätzen, die sich empirisch mit Forschungsfragen aus den Bereichen Makroökonomie und Finanzmärkte beschäftigen. Ein besonderer Fokus der Dissertation liegt dabei auf den Effekten von unkonventioneller Geldpolitik (UGP). Nach dem Ausbruch der globalen Finanz- und Wirtschaftskrise, und bei durch die Nullzinsschranke restringierter konventioneller Geldpolitik, haben Zentralbanken zunehmend unkonventionelle Maßnahmen wie Anleihekaufprogramme oder eine Lockerung der Regeln zur Kreditvergabe ergriffen. Diese neu eingeführten Maßnahmen haben zu einer intensiven Debatte darüber geführt, ob sie wirksam zu einer Verbesserung der wirtschaftlichen Situation beitragen und welche potentiellen Nebeneffekte mit ihnen verbunden sind.

Zwei der in der Dissertation enthaltenen Aufsätze tragen zu dieser Debatte bei. Ein Forschungspapier beschäftigt sich mit den Auswirkungen von UGP in den USA auf andere Länder. Es geht der Frage nach, welchen Einfluss US UGP auf Kapitalflüsse in Schwellenländer hat, und inwiefern sie die Makroökonomie und Finanzmärkte in diesen Ländern beeinflusst. Ein anderes Forschungspapier setzt sich mit der Wirksamkeit und potentiellen Nebeneffekten von UGP der Europäischen Zentralbank auseinander. Dabei wird analysiert, ob UGP das Wirtschaftswachstum und die Verbraucherpreise im Euroraum insgesamt und in den größten Mitgliedsländern stimulieren kann. Zudem wird untersucht, wie die Fiskalpolitik auf unkonventionelle geldpolitische Schocks reagiert und ob diese Schocks bestehende Leistungsbilanzungleichgewichte innerhalb der Währungsunion verstärken oder verringern.

Ein dritter Aufsatz beschäftigt sich mit einem Thema, das für die Gestaltung der Geldpolitik von großer Bedeutung ist: Die Verankerung von Inflationserwartungen. Inflationswartungen sind eine zentrale Determinante zukünftiger Inflationsraten, dies ergibt sich beispielsweise aus Neu-Keynesianischen Modellen mit entsprechender Phillips-Kurve (siehe Galí, 2015). Zentralbanken zielen daher darauf ab, die Erwartungen beim Inflationsziel verankert zu halten. Der Aufsatz liefert neue Evidenz hinsichtlich der Verankerung von Inflationserwartungen in den USA.

Ein viertes Forschungspapier befasst sich mit der sogenannten Finanzialisierung der Rohstoffterminmärkte. Die zunehmende Präsenz von Finanzinvestoren auf diesen Märkten ging mit drastischen Schwankungen von Rohstoffpreisen einher, was eine intensive Debatte über die Rolle ausgelöst hat, welche Investoren in der Entwicklung der Preiszyklen spielen. Insbesondere wird debattiert, ob Investoren für die große Preisvolatilität verantwortlich sind und ob sie generell dafür sorgen, dass sich Preise von ihren fundamentalen Determinanten entkoppeln. Im Papier wird eine neue empirische Strategie vorgeschlagen und implementiert, mit welcher der Einfluss von Finanzinvestoren auf Rohstoffterminpreise evaluiert werden kann.

Zur Beantwortung der jeweiligen Forschungsfragen verwenden alle vier Papiere strukturelle Vektorautoregressive Modelle (SVARs). Seit dem einflussreichen Beitrag von Sims (1980) sind SVARs einer der wichtigsten methodischen Ansätze für empirische Forschung in den Bereichen Makroökonomie und Finanzmärkte geworden. Ziel einer SVAR Analyse ist das Aufdecken und Quantifizieren kausaler Zusammenhänge in den Daten (vergleiche Kilian and Lütkepohl, 2017). Insbesondere gilt es dabei, unkorrelierte und ökonomisch interpretierbare strukturelle Schocks zu identifizieren und ihre kausalen Effekte auf die im Modell enthaltenen Variablen zu schätzen.

Die Identifikation ökonomisch interpretierbarer Schocks ist die größte Herausforderung in der Anwendung von SVAR Modellen. Technisch entspricht die Identifikation des Modells einer Zerlegung der Fehlerterme der reduzierten Form des VARs, die aus den Daten geschätzt werden können, in unbeobachtbare strukturelle Schocks von Interesse. Verschiedene Ansätze wurden dafür seit Sims (1980) entwickelt und angewendet. Am bekanntesten sind Kurzfristrestriktionen, bei denen angenommen wird, dass der unmittelbare Effekt ausgewählter struktureller Schocks auf im Modell enthaltene Variablen Null ist. Auch auf langfristige Effekte von strukturellen Schocks auf Variablen im System können Null-Restriktionen auferlegt werden. Alternativ ist es möglich, ein Menge von zulässigen Modellen anhand von Restriktionen auf die Form oder das Vorzeichen der geschätzten Impuls-Antworten des Modells zu identifizieren, oder die Größenordnung von geschätzten Parametern oder deren Einfluss auf die Varianz der im Modell befindlichen Variablen zu beschränken.

In dieser Dissertation werden eher kürzlich etablierte Methoden zur Identifikation verwendet. Ein Aufsatz verbindet Null- und Vorzeichen-Restriktionen (siehe Mountford and Uhlig, 2009, Baumeister and Benati, 2013, Arias et al., 2014), ein anderer Aufsatz nutzt die in den Daten vorhandene Heteroskastizität (siehe Sentana and Fiorentini, 2001, Rigobon, 2003, Lanne and Lütkepohl, 2008), während die zwei übrigen Aufsätze externe Information aus Proxy-Variablen verwenden (in der Literatur auch externe Instrumente genannt, siehe Stock and Watson, 2012, Olea et al., 2013, Mertens and Ravn, 2013). Die vier individuellen Aufsätze sind in vier Kapiteln angeordnet. Die Hauptergebnisse sowie der Beitrag zur Literatur von jedem einzelnen sind im Folgenden kurz zusammengefasst:

- ***Kapitel 1:*** *Spillovers of US Unconventional Monetary Policy to Emerging Markets: The Role of Capital Flows*

  Der erste Aufsatz basiert auf einem wissenschaftlichen Artikel, der gemeinsam mit Pablo Anaya und Christian Offermanns geschrieben wurde, und globale Effekte von UGP in den USA analysiert. Wir untersuchen in einem globalen strukturellen VAR, ob US UGP Schocks signifikante Effekte auf die wirtschaftliche Situation und die Finanzmärkte in Schwellenländern haben, und ob internationale Kapitalflüsse ein wichtiger Transmissionskanal der Schocks sind. Zur Identifikation der geldpolitischen Schocks benutzen wir die Bilanz der US Zentralbank als Indikatorvariable. Eine Kombination aus Null- und Vorzeichenrestriktionen ermöglicht es, den UGP Schock von anderen Schocks zu unterscheiden, auf welche die Zentralbank mit Maßnahmen reagiert, die ihre Bilanzsumme vergrößern, wie beispielsweise Finanzmarktschocks. Unsere Schätzungen ergeben, dass ein expansiver UGP Schock zu einer signifikanten Zunahme von Portfolioströmen aus den USA in Schwellenländer für fast zwei Quartale führt. Dies ist verbunden mit signifikanten Reaktionen von dortigen realwirtschaftlichen Variablen und Finanzmärkten. Zudem reagieren Zentralbanken in Schwellenländern auf den Impuls mit einer Lockerung der eigenen Geldpolitik. Die Ergebnisse scheinen unabhängig vom zugrundeliegenden Wechselkursregime, der Qualität der Institutionen oder dem Grad der finanziellen Offenheit der untersuchten Länder zu sein.

  Unsere Studie trägt zu einer Literatur bei, die sich mit den globalen Effekten von US UGP beschäftigt. Der globale VAR Ansatz erlaubt es, Zusammenhänge von Variablen sowohl über die Zeit als auch zwischen einzelnen Ländern zu berücksichtigten. Unsere Ergebnisse ergänzen bestehende Resultate, die eine sofortige Portfolioreallokation nach UGP Ankündigungen finden (siehe beispielweise Fratzscher et al., 2016b), indem wir zeigen, dass UGP Schocks persistente Effekte auf Kapitalflüsse haben und diese ein wichtiger Transmissionskanal der Schocks sind. Zudem komplementieren unsere Ergebnisse Studien, die finden, dass konventionelle US Geldpolitik globale Effekte hat und auch Länder mit flexiblem Wechselkurs davon betroffen sind (siehe beispielsweise Miranda-Agrippino and Rey, 2015, und Passari and Rey, 2015). Wir zeigen, dass US UGP ein Treiber von wirtschaftlichen Bedingungen in Schwellenländern ist, auch in solchen mit flexiblem Wechselkurs.

- ***Kapitel 2:*** *Identifying Speculative Demand Shocks in Commodity Futures Markets through Changes in Volatility*

  Der zweite Aufsatz ist eine gemeinsam mit Malte Rieth verfasste Studie, die den Effekt von Finanzspekulationen auf die Renditen von Terminkontrakten landwirtschaftlicher Rohstoffe untersucht. Dafür wird ein VAR Modell verwendet mit öffentlich verfügbaren wöchentlichen Positionsdaten der US Commodity Futures Trading Commission, welche nach verschiedenen Gruppen von Händlern aggregiert sind. Um exogene Variation in den Positionen der zwei wichtigsten Gruppen spekulativer Anleger, Indexinvestoren und Hedge Funds, identifizieren zu können, nutzen wir die in den Daten vorhandene Heteroskedastizität. Wir verwenden diese agnostische Identifikationsstrategie da Null-Restriktionen in wöchentlichen Finanzmarktdaten schwierig zu verteidigen sind, während es Vorzeichen-Restriktionen nicht ermöglichen würden, die strukturellen Schocks von Interesse zu unterscheiden. Unsere Schätzergebnisse legen nahe, dass idiosynkratische Schocks, welche jeweils die netto Long-Positionen von Indexinvestoren oder Hedge Funds erhöhen, die Renditen unmittelbar signifikant steigern. Zudem scheinen diese Schocks ein relevanter Treiber der Renditen zu sein, insbesondere wenn die spekulative Nachfrage volatil ist. Alles in allem bestimmen allerdings fundamentale Determinanten, wie Angebot und Nachfrage der zugrundeliegenden Rohstoffe, die Terminpreise.

Die Studie trägt zu einer Literatur bei, die empirisch untersucht, ob und wie die zunehmende Beteiligung von Finanzinvestoren die Preisbildung auf Rohstoffterminmärkten beeinflusst. Eine große Herausforderung ist dabei, die Variation in den Positionen von Finanzinvestoren zu identifizieren, die durch von ihnen initiierte Trades entsteht bzw. bei der sie nicht nur auf Angebote anderer Marktteilnehmer eingehen. Zwei kürzlich verfasste Studien adressieren dieses Problem mit disaggregierten oder privaten täglichen Daten und finden signifikante Preiseffekte von Finanzinvestitionen (siehe Cheng et al., 2015, Henderson et al., 2015). Unsere Resultate mit öffentlich verfügbaren Daten erlauben einen ähnlichen Schluss, darüber hinaus liefert der SVAR Ansatz neue Evidenz hinsichtlich der Bedeutung von Finanzinvestitionen für die historische Entwicklung der Rohstoffpreise. Unser Modell lässt sich zudem replizieren und in der Zukunft mit neuen Daten erweitern.

- ***Kapitel 3:*** *The Anchoring of Inflation Expectations: Evidence from SVARs identified with Macroeconomic News Announcements*

  Der dritte Aufsatz untersucht die Verankerung von Inflationserwartungen in den USA. Langfristige Inflationserwartungen, die gut verankert sind, sollten nicht auf Neuigkeiten über kurzfristige makroökonomische Entwicklungen reagieren, sofern die Zentralbank glaubwürdig ist und ein implizites oder explizites Inflationsziel hat. Meine Studie liefert neue Evidenz hinsichtlich der Verankerung von Inflationserwartungen in einem strukturellen VAR mit täglichen Daten zu Erwartungen basierend auf Finanzmarktvariablen. Ich identifiziere einen sogenannten Makro News Schock, der Neuigkeiten abbildet, basierend auf einem Proxy-Variablen Ansatz. Als Proxies benutze ich dabei beobachtbare makroökonomische Überraschungen, das heißt den Unterschied zwischen dem tatsächlich veröffentlichten Wert einer makroökonomischen Größe und der entsprechenden Erwartung der Marktteilnehmer, gemessen durch Umfragen. Durch den gewählten Ansatz sind die identifizierten strukturellen Schocks per Konstruktion mit den beobachtbaren Überraschungen korreliert. Es zeigt sich, dass die Schocks kurzfristig einen signifikanten Einfluss auf Inflationserwartungen haben und mehr als 10 % von deren Prognosefehlervarianz erklären. Langfristig laufen die Effekte des Schocks jedoch aus, was eine Verankerung der Erwartungen in der langen Frist impliziert. Zudem finde ich, dass die Reaktion der Inflationswartungen auf den Schock nach der Finanzkrise stärker ausfällt als davor.

Das Kapitel trägt wie folgt zur Literatur bei. Üblicherweise wird die Verankerung von Inflationserwartungen untersucht, indem man deren unmittelbare Reaktion auf makroökonomische Überraschungen an Veröffentlichungstagen schätzt (siehe Gürkaynak et al., 2010b, und nachfolgende Studien). Nautz et al. (2016) führen einen SVAR Ansatz ein, um die Reaktion der Erwartungen auf Makro News Schocks über die Zeit zu messen. Die Schocks identifizieren sie mit einem statistischen Verfahren. Im Gegensatz dazu schlage ich vor, die Makro News Schocks so zu identifizieren, dass sie per Konstruktion mit beobachtbarer Information über makroökonomische Überraschungen korreliert sind. Ich zeige, dass die Ergebnisse von Nautz et al. (2016) qualitativ robust gegenüber dieser Modifikation sind und dass die dynamische Reaktion von Inflationserwartungen auf Überraschungen nach der Finanzkrise stärker ist als vorher.

- ***Kapitel 4:*** *Unconventional Monetary Policy, Fiscal Side Effects and Euro Area (Im)balances*

  Der vierte Aufsatz, welcher auf einer gemeinsamen wissenschaftlichen Arbeit mit Michele Piffer und Malte Rieht beruht, analysiert die makroökonomischen Effekte von unkonventioneller Geldpolitik im Euroraum mit strukturellen VARs, die mit einem externen Instrument identifiziert werden. Als Instrument benutzen wir die gemeinsame unerwartete Variation in Renditespannen zwischen Staatsanleihen von verschiedenen Ländern des Euroraums gegenüber Deutschland an Tagen, an denen die Europäische Zentralbank unkonventionelle geldpolitische Maßnahmen angekündigt hat. Das Benutzen dieser hochfrequenten Information, um UGP Schocks zu identifizieren, hat den Vorteil, dass die Effekte von geldpolitischen Maßnahmen erfasst werden, ohne diese auf ihre Implementierung zu beschränken. In anderen Worten, auch die Effekte der bloßen Ankündigung der Maßnahmen werden in Betracht gezogen. Wir finden, dass expansive geldpolitische Schocks, welche die Renditen auf Staatsanleihen im Euroraum verringern, auch Renditen privater Anleihen und Finanzmarktunsicherheit im Euroraum reduzieren und zu einem Anstieg der Produktion sowie der Verbraucherpreise und Inflationserwartungen führen. Dies impliziert, dass unkonventionelle Geldpolitik wirksam sein kann. Wir finden auch, dass expansive Schocks zu einem Anstieg der primären Staatsausgaben im Euroraum als Ganzes und wichtigen Mitgliedsländern führen. Die fiskalische Reaktion ist allerdings heterogen und auch Produktion und Preise reagieren nicht in allen Ländern gleich. Diese Heterogenität ist wiederum mit einer Divergenz der relativen Preise in der Währungsunion und einer Zunahme von Leistungsbilanzdefiziten einzelner Länder gegenüber Deutschland verbunden.

Das Forschungspapier trägt zwei Aspekte zur Literatur bei. Erstens liefern wir neue empirische Evidenz hinsichtlich der makroökonomischen Wirksamkeit der unkonventionellen Geldpolitik der Europäischen Zentralbank in einem VAR Modell, in dem der UGP Schock mit Hilfe von hochfrequenter Information identifiziert wird (vergleiche Gertler and Karadi, 2015, für konventionelle US Geldpolitik, and Rogers et al., 2016, für US UGP). Zweitens ist dies - nach meinem besten Wissen - die erste Studie, die empirisch die Reaktion von Fiskalpolitik und Leistungsbilanzen innerhalb der Eurozone auf (unkonventionelle) geldpolitische Maßnahmen der Europäische Zentralbank untersucht.

CHAPTER 1

Spillovers of US Unconventional Monetary Policy to Emerging Markets: The Role of Capital Flows[1]

1.1 Introduction

Since the onset of the global financial crisis, large and volatile capital flows into emerging market economies (EMEs) have triggered a renewed interest in the determinants and consequences of such cross-border flows. A growing literature perceives a "global financial cycle" to be a key determinant of capital flows into EMEs (see Nier et al., 2014). This cycle is described as co-movement in gross capital flows, credit conditions, and asset prices across countries (see Rey, 2013, or Passari and Rey, 2015).[2] Rey (2013, 2016) argues that one of the main drivers of the cycle is monetary policy by the US Federal Reserve, whose interest rate decisions are transmitted to EMEs' financial conditions through international capital flows. Moreover, she reasons that, as a consequence, US interest rate decisions influence the conduct

[1]This chapter is based on a research paper that is joint work with Pablo Anaya and Christian Offermanns. We thank Benjamin Beckers, Kerstin Bernoth, Ambrogio Cesa-Bianchi, Sandra Eickmeier, Marcel Fratzscher, Georgios Georgiadis, João Tovar Jalles, Dieter Nautz, Mathias Trabandt, two anonymous referees and participants of the DIW Macroeconometric Workshop 2015, Berlin, the FMM Annual Conference 2015, Berlin, the International Conference on Macroeconomic Analysis and International Finance 2016, Crete, the Conference of the Royal Economic Society 2016, Brighton, the EC^2 Conference 2016, Edinburgh, and seminar participants at the Bank of England, at the Deutsche Bundesbank, at the University of Bamberg, and at the Freie Universität Berlin for helpful comments and suggestions.

[2]Passari and Rey (2015) and Miranda-Agrippino and Rey (2015) provide evidence for the existence of a global financial cycle by showing that prices of stocks and other risky assets as well as credit, leverage, and gross capital flows around the world are related to a common global factor.

of monetary policy in EMEs. Responding to the financial crisis and the subsequent sluggish recovery, however, the Federal Reserve (Fed) repeatedly engaged in unconventional monetary policy measures such as large-scale asset purchases. While many studies analyze how conventional interest rate policy by the Fed can drive financial conditions globally (see, among others, Miranda-Agrippino and Rey, 2015, or Bruno and Shin, 2015), results on unconventional measures are rather scarce.

In this paper, we provide new evidence on the global role of US monetary policy by empirically investigating the effects of US unconventional monetary policy (UMP) shocks on real and financial key indicators in EMEs. In particular, we analyze whether UMP shocks are a driver of capital flows into EMEs and whether capital flows are an important channel of shock transmission. In this regard, Bruno and Shin (2015) establish a link between US monetary policy and international banking flows for the period 1995–2007. However, Shin (2013) presents evidence that banking flows strongly diminished since the beginning of the financial crisis. Instead, bond and equity flows to EMEs increased heavily. Accounting for this change in the capital flow composition, we evaluate whether portfolio bond or equity flows play a pivotal role in transmitting US UMP shocks to EMEs.

We structure our analysis into the following consecutive questions. First, is US UMP related to portfolio flows into EMEs? Second, does US UMP affect asset prices and exchange rates (henceforth: financial conditions) in EMEs? If so, are portfolio flows an important channel of shock transmission? Third, is the conduct of monetary policy in EMEs in this way influenced by US UMP? To answer these questions, we estimate a structural global vector autoregressive (GVAR) model incorporating both real and financial variables for 39 advanced and emerging market economies over the period 2008–2014 and evaluate the dynamic responses of these variables to a US UMP shock. Identification of the UMP shock is based on the approach of Gambacorta et al. (2014): we use the size of the Fed balance sheet as an unconventional monetary policy instrument and employ a mixture of zero and sign restrictions to distinguish exogenous policy changes from endogenous reactions to other shocks.

To provide a general assessment of these questions, we turn to the panel dimension of the EMEs included in our sample and average over the estimated impulse response functions. Given that existing studies on international capital flows (see below) often find a substantial degree of heterogeneity in the responses of individual countries to various pull and push factors, we then also investigate whether specific economic country characteristics are associated with diverging responses to the UMP shock across country groups.[3]

Our work is related to two lines of empirical studies that address similar questions. First, it is linked to studies that analyze the determinants of capital flows to EMEs (see, among many others, Fratzscher, 2012, or Forbes and Warnock, 2012). More closely, it relates to those studies that explicitly analyze the impact of the Fed's unconventional monetary policy on international capital flows. One set of papers analyzes this question in cross-country panel frameworks (Moore et al., 2013, Koepke, 2014, Lim et al., 2014, Ahmed and Zlate, 2014, Lo Duca et al., 2016). Other studies employ high frequency and event study approaches assessing the relationship in selected windows around policy events (International Monetary Fund, 2013, Rai and Suchanek, 2014, Fratzscher et al., 2016b).[4] Second, our work is connected to studies that investigate the effect of US unconventional monetary policy on EMEs financial and real conditions in event studies or comparable frameworks (Eichengreen and Gupta, 2015, Bowman et al., 2015, Aizenman et al., 2016a).

We contribute to these two strands of the literature by employing a global VAR approach that takes interactions between financial and real variables both between countries and over time into account. Given that the international transmission of shocks via financial and trade linkages takes time, this is an important feature that existing event studies or panel approaches do not account for. Moreover, our study provides a systematic assessment of US unconventional monetary policy on the global financial cycle and international portfolio flows in a VAR framework, where

[3]Thereby, a special emphasis is given to how responses differ with respect to the countries' exchange rate regimes. This is motivated by the classical "trilemma", which states that countries can only have two of the following three policy options: independent monetary policy, free capital flows, or a fixed exchange rate. Hence, countries that let their currency float freely might have potentially more leeway to set interest rates independently of the US and, in turn, might be able to buffer foreign shocks better. For further discussion, references, and empirical evidence on the trilemma, see Obstfeld et al. (2005), Klein and Shambaugh (2015), or Aizenman et al. (2016b).

[4]Only a few studies employ VAR models, and in those studies capital flows are usually aggregated across countries (see Dahlhaus and Vasishtha, 2014, or Tillmann, 2016). Our approach differs from these models as we analyze the effect of US UMP on capital flows and real and financial variables for individual economies, accounting for trade and financial relations in a global model.

the policy shock is structurally identified. It thus allows quantifying the persistent effects of UMP shocks on capital flows and other variables for individual countries.

Conceptually, the paper compares directly to the fast growing literature on the global effects of monetary policy shocks in structural VAR or GVAR frameworks (see, among others, Miranda-Agrippino and Rey, 2015, Passari and Rey, 2015, Georgiadis, 2015, Georgiadis, 2016, or Chen et al., 2016). For instance, Miranda-Agrippino and Rey (2015) show that conventional US monetary policy is linked to a global factor in risky asset prices and cross-border credit flows. Passari and Rey (2015) and Rey (2016) provide evidence for an effect of conventional US monetary policy on financial conditions in a few small open advanced economies. Closest to our paper is a study by Chen et al. (2016) who also investigate spillovers from US unconventional monetary policy on a number of variables in several AEs and EMEs in a GVAR framework. We complement their work by analyzing the role of capital flows[5] as an important channel of transmission of US monetary policy shocks and by addressing the debate on the global financial cycle as we consider a broad set of EME financial indicators. Moreover, we differ from their approach by directly using an policy instrument, the size of the balance sheet, to identify unconventional monetary policy innovations.

Our main results can be summarized as follows. First, we find that an expansionary US monetary policy shock, associated with an exogenous innovation to the Fed balance sheet, significantly increases portfolio outflows from the US for almost two quarters. In the EMEs this is, on average, associated with a rise in portfolio inflows. Second, this increase in inflows is accompanied by significant real and financial effects in EMEs. In response to the UMP shock, real output growth and equity returns increase, and the real exchange rate appreciates. Importantly, portfolio flows prove to be a key channel of transmission between the US and the EMEs in the GVAR specification. Third, with regard to domestic monetary policy, we find that on average, the short-term (policy) interest rate in EMEs declines in response to an expansionary US shock. This result indicates that US policy innovations have an influence on the conduct of monetary policy in EMEs. Regarding potential differences of our results across countries, we find that the magnitude of portfolio inflows to EMEs appears to vary with the proximity to the US. However, economic coun-

[5]Henceforth we will use the terms “capital flows” and “portfolio flows” interchangeably.

try characteristics like the degree of financial openness or the quality of domestic institutions do not affect the countries' response to the UMP shock. In particular, a floating exchange rate arrangement does not appear to provide a better insulation from US spillovers.

The remainder of the paper is structured as follows. In Section 1.2, we outline transmission channels through which US UMP can affect international portfolio flows and structure our empirical analysis into three hypotheses along the research questions. Section 1.3 comprises a description of the data and of the GVAR specifications. In Section 1.4, we present our main results and a sensitivity analysis, and study the role of country characteristics. Section 1.5 concludes.

1.2 UMP, portfolio flows, and financial conditions in EMEs

In our analysis, we are interested in how US UMP shocks affect financial conditions in EMEs, i.e. we analyze the time period during and after the financial crisis. Moreover, we investigate whether international capital flows are an important channel of shock transmission. Shin (2013) presents evidence that during and after the financial crisis direct (portfolio) bond and equity flows play a pivotal role in capital flows to EMEs. In contrast, the formerly dominant international banking flows strongly diminished since the beginning of the crisis. Therefore, we focus our analysis on the role of portfolio flows as an important transmission channel of UMP shocks.[6]

From a theoretical perspective, there are several channels through which the Fed's UMP can affect portfolio allocation decisions by asset managers with a global reach and hence, international portfolio flows (see, e.g., Fratzscher et al., 2016b). In principle, they are all linked to the potential domestic transmission channels of UMP brought forward by the literature (for a review, see Joyce et al., 2012). First and foremost, a portfolio balance channel is often emphasized. A Fed purchase of e.g. US Treasury bills crowds out private investment from this market segment. In turn,

[6] The arguments of Shin (2013) are in detail laid out in Azis and Shin (2015). A summary of the evidence given by Shin (2013) can also be found in the working paper version of this chapter. We do not additionally consider banking flows in the estimations for two reasons. First, the reviewed stylized facts indicate that the importance of banking flows has strongly diminished since the beginning of the global financial crisis. Second, banking flows are available only at a quarterly frequency. Given the short time period covering the financial crisis and UMP, we do not have enough quarterly observations for a meaningful estimation of their role after the crisis.

investors rebalance their portfolio and move to close substitute assets. Ultimately, a chain of rebalancing is set in motion which may affect the allocation of assets across countries. Second, the UMP measures can affect the risk appetite of investors, often termed "confidence or risk-taking" channel. While theoretical descriptions of this channel are often rather informal, Bekaert et al. (2013) provide empirical evidence that conventional monetary policy in the US significantly impacts on financial market risk taking. An increased risk appetite might drive investors into more risky high-yield EME assets. Third, UMP might work through a signaling channel. If the Fed's measures are understood by markets as keeping future interest rates low for a period longer than previously expected, this can affect asset prices by lowering the risk-neutral price component of (future) interest rates. This induced fall in yields on securities in the US can, in turn, lead to an increase in capital flows to EMEs due to a "search-for-yield".[7]

Foreshadowing our UMP shock identification strategy (see Section 1.3.3), our approach is linked more closely to transmission channels which imply effects not only from the policy announcement, but also from the actual policy implementation. This feature is particularly relevant for the portfolio balance channel, where the available supply of assets is changed through actual purchases. Regarding the other channels, the consideration of actual purchases is often criticised as an incomplete view on the effects of UMP. This objection claims that, since amount and timing of the Fed's UMP – in the form of large-scale asset purchase programs – were transparently communicated, efficient markets should have fully processed the information when it was announced. However, as Fratzscher et al. (2016b) argue, there are two important points against this notion. First, many UMP measures were installed pre-

[7]It has to be emphasized that these channels are not mutually exclusive and may rather be at work simultaneously (see Fratzscher et al., 2016b) and we do not aim at disentangling them in this paper. Moreover, all of them imply to some extent falling yields in the US and subsequent capital outflows related to a search for yields and an increase in risk appetite. The signaling channel, however, potentially also allows for a negative effect of UMP on capital flows. If new UMP measures are understood as indicating that economic conditions are worse than expected, this might induce a flight to safety and a decrease in capital outflows to EMEs as investment in these countries is perceived as being risky. Hence, in our empirical implementation, we do not impose a sign restriction on the effect of UMP on US bond and equity outflows. As pointed out by an anonymous referee, also note that the different channels potentially induce a different timing in the response of capital flows and financial markets. In our empirical analysis, we therefore also conduct a counterfactual exercise that showcases the importance of capital flows for the transmission of the UMP shock to other variables in the estimated model.

cisely because markets were not functioning. Hence, mere announcements may have been less important than actual purchases because the latter ones restored liquidity and allowed investors to adjust their exposure. Second, even if market expectations were fairly precise about the actual amount and timing of the purchases, still the expectations about effectiveness of the actual UMP measures might not have been accurate. In their empirical analysis, Fratzscher et al. (2016b) show that Fed operations, such as purchases of Treasury securities, indeed had larger effects on the portfolio decisions than Fed announcements of these programs.

Based on the discussion outlined in the introduction and this section, we summarize our research questions for the empirical analysis in three hypotheses in logical order. We will then interpret our estimation results along these hypotheses.

Hypothesis 1: *US UMP affects portfolio outflows from the US, and, in turn, portfolio inflows into EMEs.*

If Hypothesis 1 holds true, our estimated model should show an increase in portfolio outflows from the US and in portfolio inflows to EMEs after an expansionary UMP shock as implied by the three channels discussed above.

The second hypothesis links the response of US portfolio flows to the debate on whether and how US monetary policy determines financial conditions globally. In order to test both for effects and their cause, we split the hypothesis into two parts:

Hypothesis 2: *US UMP drives financial conditions in EMEs (**2a**) and portfolio flows are an important channel of transmission (**2b**).*

The first part, Hypothesis 2a, is supported by our results if we find that variables which reflect financial conditions in EMEs significantly react in response to a US UMP shock. Thereby, we assess financial conditions in our analysis using equity returns, to capture general financial market developments, and the foreign exchange rate. In response to an expansionary US UMP shock, equity returns should increase and the currency should appreciate with respect to the US dollar. For the second part, Hypothesis 2b, to hold true portfolio flows should constitute an important channel of transmission of US UMP to EMEs in the empirical specification. In this regard, estimation results in the GVAR should not solely be driven by other transmission channels potentially related to a US UMP shock like demand effects e.g. by an increase in trade with the US, or an increase in global growth. We address this issue in estimating the GVAR allowing for different channels of transmission, see Sections 1.3.3 and 1.3.5.

Finally, the third hypothesis focuses on how domestic monetary policy in EMEs reacts to the UMP shock. It connects our estimations to the discussion on whether and how US monetary policy influences the conduct of monetary policy in other countries (see, e.g., Rey, 2013, or Klein and Shambaugh, 2015). By similarly splitting the hypothesis into two parts, we investigate the role of portfolio flows for this effect:

Hypothesis 3: *US UMP has a significant impact on EMEs' monetary policy (**3a**) and portfolio flows are an important channel of transmission (**3b**).*

The first part, Hypothesis 3a, is supported by our results if we find that policy rates in EMEs significantly react in response to a US UMP shock. In particular, we expect policy rates in EMEs to be lowered despite the easing of real and financial conditions, mirroring the expansionary US UMP shock as a consequence of the increase in portfolio flows from the US to EMEs (Hypothesis 3b).

In the empirical analysis, we first generally assess these hypotheses for our panel of countries by averaging the estimated impulse response function across all EME countries. We then also analyze whether economic characteristics of the countries, like the exchange rate regime or the quality of institutions, lead to heterogeneity in the results.

1.3 Data and empirical methodology

1.3.1 Data

We use monthly data on US portfolio bond and equity asset outflows and EMEs financial and real conditions from January 2008 to December 2014. The sample captures the period in which the Fed conducted measures of unconventional monetary policy. Our source of bond and equity flows are the monthly estimates of changes in US holdings of foreign securities provided by the Federal Reserve Board. The dataset is based on estimations that combine two different types of data reported by the Treasury International Capital Reporting System (TIC). On the one hand, data is based on annual benchmark surveys of US holdings of foreign securities. On the other hand, transaction data from mandatory monthly TIC surveys, filed by US banks, securities dealers, and other entities that report net purchases of foreign assets by US residents, is used.[8] The dataset covers portfolio investment in long-term securities, specifically debt instruments with greater-than-one-year original maturity (bonds), and equities. It is widely used in the literature on studying foreign flows into US securities and US flows into foreign securities and considered as highly accurate (see, e.g., Curcuru et al., 2010, or Hanlon et al., 2015).

Such a large and comprehensive dataset on bilateral capital flows is available only for flows out of the US. For the purpose of our analysis, this is not a major concern as we are interested in the effects of US monetary policy that should affect flows from the US more heavily than bilateral flows involving third countries. Moreover, US portfolio flows are of major relevance for the emerging market economies that we study. According to data from the 2012 Coordinated Portfolio Investment Survey, US investors account for more than a third of all cross-border investment in bonds and equities of emerging market economies (see Bertaut and Judson, 2014). The dataset assigns foreign investment of US investors to the country where the entity issuing the security is legally a resident. Hence, it accounts for the so-called ‘transaction bias’, where the investment is assigned wrongly to the financial center in which the transaction takes place. However, the dataset cannot account for the issuance of securities by EME firms through subsidiaries residing in financial centers. Empirical evidence shows that emerging market firms have increasingly issued debt through foreign subsidiaries in the UK, Ireland, Luxembourg, and other (off-shore) financial centers (see Department of the Treasury, Federal Reserve Bank of New York and Board of Governors of the Federal Reserve System, 2016). This implies that our approach yields estimates which should be interpreted as a lower bound of the effects of US UMP on EMEs.

Further data on EMEs’ real and financial conditions is obtained from Thomson Reuters Datastream and the IMF’s International Financial Statistics (IFS) database,

[8] Bertaut and Tryon (2007) for the period 1994–2010 and Bertaut and Judson (2014) for the period 2011–2014 combine the monthly transaction data with the yearly survey data to obtain estimates on monthly levels for both flows and valuation changes. The data can be accessed at the following website: http://www.federalreserve.gov/pubs/ifdp/2014/1113/default.htm. The estimated monthly flows are subject to several adjustments to reduce the noise and biases in the underlying monthly TIC flow data. The necessary steps are in detail described in Bertaut and Tryon (2007) and Bertaut and Judson (2014). We employ only the data on flows and not on valuation changes, since we are only interested in investors’ decisions following the UMP shock.

in particular gross domestic product (GDP), industrial production (IP), consumer price index (CPI), equity prices, exchange rates, interest rates, commodity prices and foreign exchange reserves. Data on US GDP, IP and CPI as well as the Fed balance sheet and the option-implied volatility index VIX is taken from the Federal Reserve Bank of St. Louis database (FRED). Time series with monthly observations for real GDP are constructed by interpolating the quarterly figures with the monthly index of industrial production, using the method of Chow and Lin (1971). All variables in nominal terms have been converted to real terms prior to the estimation using domestic CPI. Detailed information on data sources and transformations is given in Table 1.A.1 (Appendix 1.A).

Selection of emerging market countries into the sample is driven by different considerations. First, we do not include countries that have limited access to global financial markets and hardly issue securities and equities globally. Second, we only add countries with comprehensive monthly data on economic and financial conditions available. Third, we exclude China as it plays a distinct role through its trade and financial linkages to the US and its particular institutional settings. Nevertheless, the dataset covers a broad range of EMEs and closely resembles the sample of many studies on emerging markets (see, for instance, Bowman et al., 2015, or, Aizenman et al., 2016a). Ultimately, our sample contains the following 19 countries: Argentina, Brazil, Chile, Colombia, Hungary, India, Indonesia, Israel, Korea, Mexico, Malaysia, Peru, the Philippines, Poland, Russia, Singapore, South Africa, Thailand, and Turkey. In addition, we add 16 advanced economies to the model to account for trade and financial linkages between them and the included EMEs.[9]

1.3.2 The GVAR model

To analyze the effects of US UMP on a broad set of EMEs while taking all potential cross-country interlinkages into account, in principle, a large scale vector autoregressive (VAR) model of the following form would be adequate:

$$\boldsymbol{y}_t = \boldsymbol{\mu} + \boldsymbol{\lambda} t + \boldsymbol{\Gamma}_1 \boldsymbol{y}_{t-1} + \boldsymbol{\Gamma}_2 \boldsymbol{y}_{t-2} + ... + \boldsymbol{\Gamma}_p \boldsymbol{y}_{t-p} + \boldsymbol{\nu}_t \tag{1.1}$$

[9] The added AEs are Australia, Belgium, Canada, Germany, Denmark, Finland, France, Hong Kong, Italy, Japan, the Netherlands, Norway, Spain, Sweden, Switzerland, and the UK. Our model is specified and estimated for the AEs (except the US) in the same way as for the EMEs.

where $\boldsymbol{y}_t = (\boldsymbol{y}'_{1t}\ \boldsymbol{y}'_{2t}\ \dots\ \boldsymbol{y}'_{Nt})'$ denotes the vector of k endogenous variables stacked across N countries for $t = 1, 2, ..., T$ time periods, $\boldsymbol{\Gamma}_s$, $s = 1, 2, ..., p$ denotes an $(Nk \times Nk)$ matrix of coefficients, $\boldsymbol{\mu}$ and $\boldsymbol{\lambda}$ denote $Nk \times 1$ vectors of constant and trend coefficients, and $\boldsymbol{\nu}_t = (\boldsymbol{\nu}'_{1t}\ \boldsymbol{\nu}'_{2t}\ \dots\ \boldsymbol{\nu}'_{Nt})'$ represents shocks. While the model can accommodate a rich structure of cross-country interrelations, estimation of this model under a fully flexible parametrization is not feasible due to the large number of parameters, even for moderate sizes of N.

Dees et al. (2007) employ a parsimonious way of re-specifying this model by modeling the relations between countries via their bilateral trade linkages (the "global VAR" (GVAR) approach). For estimation, the GVAR model is represented by a set of linked country VAR models with exogenous regressors (VARX) for each country $i = 1, 2, ..., N$

$$\boldsymbol{y}_{it} = \sum_{s=1}^{p} \boldsymbol{A}_{is}\boldsymbol{y}_{it-s} + \sum_{s=0}^{p} \boldsymbol{B}_{is}\boldsymbol{y}^*_{it-s} + \sum_{s=0}^{p} \boldsymbol{C}_{is}\boldsymbol{d}_{t-s} + \boldsymbol{\lambda}_i t + \boldsymbol{\mu}_i + \boldsymbol{\varepsilon}_{it}, \tag{1.2}$$

where $\boldsymbol{y}_{it}$ is a vector of k endogenous variables and $\boldsymbol{y}^*_{it}$ is a vector of k country-specific (weakly exogenous) "foreign variables".[10] In order to take account of potential observed common factors (in addition to unobserved common factors captured by the foreign variables), the models include the k_d-dimensional vector $\boldsymbol{d}_t$ of "global" variables affecting every country. The coefficient matrices $\boldsymbol{A}_{is}$, $\boldsymbol{B}_{is}$ and $\boldsymbol{C}_{is}$, $s = 0, 1, 2, ..., p$, as well as the coefficient vectors $\boldsymbol{\lambda}_i$ and $\boldsymbol{\mu}_i$ are of suitable dimension.

The key feature of this approach is to define the so-called foreign variables, $\boldsymbol{y}^*_{it}$, as weighted averages of other countries' variables with bilateral weights w_{ij}:

$$\boldsymbol{y}^*_{it} = \sum_{j=1}^{N} w_{ij}\boldsymbol{y}_{jt}, \qquad \sum_{j=1}^{N} w_{ij} = 1, \quad w_{ij} \geq 0\ \forall\ i, j, \quad w_{ii} = 0. \tag{1.3}$$

The weights capture the exposure of country i to country j based on trade linkages. The foreign variables $\boldsymbol{y}^*_{it}$ are assumed to be weakly exogenous with respect

[10]For notational simplicity, we present the model with the number of endogenous and foreign variables to be the same and homogeneous across countries. In the empirical implementation, these dimensions might vary. Also the common lag order is presented for convenience of processing of the coefficient matrices. It does not imply a restriction to the estimation of the model since the coefficient matrices can be filled with with zeros after country-wise estimation.

to the parameters of the VARX model in Equation (1.2). This assumption appears admissible given that N in our case is 39.

Summarizing the influence of global variables and deterministics in the vector $\boldsymbol{h}_{it} = \sum_{s=0}^{p} \boldsymbol{C}_{is}\boldsymbol{d}_{t-s} + \boldsymbol{\lambda}_i t + \boldsymbol{\mu}_i$ for notational simplicity, Equation (1.2) can be rewritten as

$$\boldsymbol{\Phi}_{i0}\boldsymbol{z}_{it} = \sum_{s=1}^{p} \boldsymbol{\Phi}_{is}\boldsymbol{z}_{it-s} + \boldsymbol{h}_{it} + \boldsymbol{\varepsilon}_{it}, \tag{1.4}$$

where $\boldsymbol{z}_{it} = (\boldsymbol{y'}_{it}, \boldsymbol{y'}^{*}_{it})'$, $\boldsymbol{\Phi}_{i0} = (\boldsymbol{I}_k, -\boldsymbol{B}_{i0})$, and $\boldsymbol{\Phi}_{is} = (\boldsymbol{A}_{is}, -\boldsymbol{B}_{is})$. Hence, $\boldsymbol{z}_{it}$ is linked to the endogenous variables $\boldsymbol{y}_t = (\boldsymbol{y}'_{1t}\ \boldsymbol{y}'_{2t}\ \dots\ \boldsymbol{y}'_{Nt})'$ through the link matrix $\boldsymbol{W}_i$ in the following way

$$\boldsymbol{z}_{it} = \boldsymbol{W}_i\boldsymbol{y}_t, \quad \boldsymbol{W}_i = \begin{pmatrix} \boldsymbol{0} & \dots & \boldsymbol{I}_k & \dots & \boldsymbol{0} \\ w_{i1}\boldsymbol{I}_k & \dots & w_{ii}\boldsymbol{I}_k & \dots & w_{iN}\boldsymbol{I}_k \end{pmatrix}. \tag{1.5}$$

Using this relation, Equation (1.4) is equivalent to

$$\boldsymbol{\Phi}_{i0}\boldsymbol{W}_i\boldsymbol{y}_t = \sum_{s=1}^{p} \boldsymbol{\Phi}_{is}\boldsymbol{W}_i\boldsymbol{y}_{t-s} + \boldsymbol{h}_{it} + \boldsymbol{\varepsilon}_{it}. \tag{1.6}$$

Stacking the individual country VARX models yields the following equation for y_t

$$\boldsymbol{G}_0\boldsymbol{y}_t = \sum_{s=1}^{p} \boldsymbol{G}_s\boldsymbol{y}_{t-s} + \boldsymbol{h}_t + \boldsymbol{\varepsilon}_t, \tag{1.7}$$

where

$$\boldsymbol{G}_0 = \begin{pmatrix} \boldsymbol{\Phi}_{10}\boldsymbol{W}_1 \\ \boldsymbol{\Phi}_{20}\boldsymbol{W}_2 \\ \dots \\ \boldsymbol{\Phi}_{N0}\boldsymbol{W}_N \end{pmatrix}, \boldsymbol{G}_s = \begin{pmatrix} \boldsymbol{\Phi}_{1s}\boldsymbol{W}_1 \\ \boldsymbol{\Phi}_{2s}\boldsymbol{W}_2 \\ \dots \\ \boldsymbol{\Phi}_{Ns}\boldsymbol{W}_N \end{pmatrix}, \boldsymbol{h}_t = \begin{pmatrix} \boldsymbol{h}_{1t} \\ \boldsymbol{h}_{2t} \\ \dots \\ \boldsymbol{h}_{Nt} \end{pmatrix}, \boldsymbol{\varepsilon}_t = \begin{pmatrix} \boldsymbol{\varepsilon}_{1t} \\ \boldsymbol{\varepsilon}_{2t} \\ \dots \\ \boldsymbol{\varepsilon}_{Nt} \end{pmatrix} \sim iid(0, \boldsymbol{\Sigma}_{\varepsilon}).$$

Equation (1.7) has to be pre-multiplied by $\boldsymbol{G}_0^{-1}$ to obtain the autoregressive representation of the GVAR model, the so-called global solution,

$$\boldsymbol{y}_t = \sum_{s=1}^{p} \boldsymbol{F}_s\boldsymbol{y}_{t-s} + \tilde{\boldsymbol{h}}_t + \boldsymbol{u}_t, \quad \boldsymbol{u}_t \sim iid(0, \boldsymbol{\Sigma}_u) \tag{1.8}$$

with $\boldsymbol{F}_s = \boldsymbol{G}_0^{-1}\boldsymbol{G}_s$, $s = 1, ..., p$, $\tilde{\boldsymbol{h}}_t = \boldsymbol{G}_0^{-1}\boldsymbol{h}_t$ and $\boldsymbol{u}_t = \boldsymbol{G}_0^{-1}\boldsymbol{\varepsilon}_t$ such that $\boldsymbol{\Sigma}_u = \boldsymbol{G}_0^{-1}\boldsymbol{\Sigma}_\varepsilon\boldsymbol{G}_0^{-1\prime}$. Estimates for the parameters of the global solution can be constructed based on the estimated individual-country VARX models. The global solution is equivalent to the reduced-form VAR representation of Equation (1.1), but with numerous within- and cross-equation restrictions. It can thus be used to perform standard VAR analysis and obtain structural impulse response functions (IRFs), as we do. In order to retain a parsimonious specification, the lag order p is set to one in the baseline case. The sensitivity analysis of our estimation results evaluates the implications of choosing higher lag orders.

1.3.3 Specification of the GVAR model

In specifying the underlying VARX models for the individual countries, we treat the US equations differently than the EMEs equations. On the one hand, we include a different set of weakly exogenous variables, similar to what is commonly done in GVAR applications due to the dominant role of the US in global financial markets (see, for instance, Eickmeier and Ng, 2015, Georgiadis, 2016, or Chen et al., 2016). More importantly, as we are interested in an unconventional monetary policy shock, we set up a model that allows identifying such a shock.

Hence, the VARX model for the US resembles VAR specifications from the literature on identifying conventional monetary policy shocks, usually containing output, inflation, and the Fed funds rate (see, among many others, Christiano et al., 1999). However, we replace the Fed funds rate as the monetary policy instrument by the size of the Fed balance sheet as in Gambacorta et al. (2014). Following the beginning of the financial crisis in 2008, the Federal funds rate soon reached its effective lower bound and stayed there for most of our sample period. Instead, the Fed introduced a number of (new) policy tools, most of which have altered both the size and the composition of its balance sheet, commonly referred to as unconventional monetary policy. The purpose of these tools has been to stabilize the functioning of financial markets, especially during the crisis, and to provide support to the economy during the recession and the subsequent sluggish recovery.

The balance sheet of the Fed more than quadrupled between 2007 and 2014 (see Figure 1.B.2 in Appendix 1.B). By using the size of the Federal Reserve balance sheet as the monetary policy instrument, the identified unconventional monetary policy shocks will be linked to all measures that increased the balance sheet. Notable policies that affected the balance sheet size start after the failure of Lehman Brothers in September 2008, as the Fed immediately provided credit to intermediaries and key markets. Other notable expansions are associated with the three programs of Quantitative Easing (QE) that were conducted. First, QE1, announced in November 2008 and expanded in March 2009, included purchases of mortgage backed securities and Treasury securities. In November 2010, it was succeeded by QE2 that focused on buying long-term Treasury securities. Lastly, QE3 was initiated in September 2012 and again included both mortgage backed securities and Treasury securities. Although these programs also affected the composition of the Fed balance sheet, the main change in the monetary policy stance was initiated through its expansion.[11] Hence, the size of the balance sheet should be a suitable instrument to measure the Fed's unconventional policy stance over our sample period.

Changes in the balance sheet size, however, do not only capture exogenous innovations to UMP, but also the endogenous reaction of the Fed to the state of the economy and, importantly, to financial market turmoil as e.g. in the immediate aftermath of the Lehman collapse. To identify an exogenous innovation to the balance sheet, we add the volatility index VIX to the model to capture financial market uncertainty. Lastly, we add total US portfolio outflows as the variable of interest for our research questions. In sum, the following definition of endogenous and foreign variables is used for the US model:[12]

$$
\begin{aligned}
\boldsymbol{y}_{US,t} &= (\text{output}, \text{inflation}, \text{VIX}, \text{Fed balance sheet}, \text{portfolio outflows})', \\
\boldsymbol{y}^*_{US,t} &= (\text{foreign output}, \text{foreign inflation})',
\end{aligned}
$$

where the balance sheet is included in its logarithm for the ease of interpretation. To distinguish between an exogenous innovation to the Fed balance sheet and an endogenous reaction of the central bank to the state of the economy or financial

[11] In contrast, two unconventional monetary policy measures that did only alter the composition of the balance sheet are on the one hand the Fed's response to the run on Bear Stearns in March 2008, when it increased its lending to investment banks and other stressed financial intermediaries, but at the same time lowered its holding of short-term Treasury securities. On the other hand, between 2011 and 2012, the Fed ran a maturity extension program in which short- and medium-term Treasury securities were sold and proceeds used to purchase long-term Treasury securities to flatten the yield curve.

[12] All models for the US and for the EMEs include a constant and a linear time trend.

market turmoil, we follow Gambacorta et al. (2014) and impose a mixture of zero and sign restrictions on the structural impulse response functions. First, in accordance with standard assumptions in the literature, we assume that a shock to the policy instrument, in our case the Fed balance sheet, only has a lagged impact on output and inflation. The Fed itself reacts instantaneously to innovations to output and inflation as commonly assumed in VAR analysis of monetary policy transmission. Second, to account for the endogenous reaction to financial market turmoil, we use the sign restrictions displayed in Table 1.3.1. On the one hand, we assume that an expansionary UMP shock does not increase the VIX. This reflects the notion that UMP had the effect of mitigating concerns about financial instability. It is also in line with results by Bekaert et al. (2013) who show that an expansionary conventional monetary policy shock has a lowering effect on the VIX.[13] On the other hand, we define a shock that affects both the VIX and the Fed balance sheet in the same direction as a financial market risk shock to which the Fed responds, most notably seen in the immediate Lehman aftermath. The sign restrictions are imposed on impact and in the first month after the shock. Note that, as outlined in Section 1.2, the shock will primarily capture the actual implementation of UMP, namely measures that enlarge the balance sheet.[14]

The VARX specification for the EMEs is restricted by the rather short period of US UMP. Given the limited number of observations, we cannot include all variables of interest into one large model. Instead, we consider two different models for different economic concepts of interest. First, we estimate a model which is focused on the response of real economic conditions to a US UMP shock ("business cycle"

[13]In theory, it is also possible that unconventional monetary policy expansions increase the volatility by increasing the uncertainty about the future path of monetary policy or if the expansion is perceived as a harbinger of less encouraging prospects (compare Section 1.2). Fratzscher et al. (2016a), however, show that, on average, press announcements by the Fed regarding unconventional monetary policy lowered the VIX on impact. Unfortunately, there is no comprehensive assessment of the impact of QE purchases on the VIX available, which is a subject potentially worth examining in future studies.

[14]Identifying the UMP shock using a Cholesky ordering and sign restrictions might pose a problem given the inclusion of financial variables, in this application most importantly US portfolio outflows. Therefore, we have assessed the sensitivity of the results towards a different ordering of portfolio flows and using an identification strategy based on a shadow interest rate (see Section 1.4.3). Another approach would be to identify the VAR using external instruments as pioneered by Gertler and Karadi (2015) for conventional monetary policy. For a pure UMP shock and our sample, however, we did not find a valid instrument, most notably due to the zero lower bound.

Table 1.3.1: Sign restrictions to identify UMP shock

	Output	Inflation	VIX	Fed Balance Sheet
Unconv. Monetary Policy Shock	0	0	≤ 0	> 0
Financial Market Risk Shock	0	0	> 0	> 0

The table shows the sign and zero restrictions on the endogenous variables (columns) applied to identify the unconventional monetary policy shock and distinguish it from a financial market risk shock (rows).

(BC) model):

$$\boldsymbol{y}_{it} = (\text{output}, \text{real exchange rate change}, \text{portfolio inflows}, \text{real interest rate})',$$
$$\boldsymbol{y}^*_{it} = (\text{US portfolio outflows}, \text{foreign output})'.$$

Second, we estimate a model to analyze the effect of a US UMP shock on financial conditions in EMEs (FC model):

$$\boldsymbol{y}_{it} = (\text{portfolio inflows}, \text{real interest rate}, \text{real exchange rate change}, \text{equity returns})',$$
$$\boldsymbol{y}^*_{it} = (\text{US portfolio outflows}, \text{foreign output})'.$$

The two models share the same VARX model for the US to ensure that the UMP shock is the same across models. In the baseline specification, we include US total portfolio outflows, the transmission channel of interest, and foreign output, as standard in the literature, as foreign variables $\boldsymbol{y}^*$ in the EMEs VARX models. Following Georgiadis (2015), we include the potentially non-stationary level variables real GDP and real exchange rate in first difference form in all models.[15]

[15] This has mainly two reasons. First, using differences ensures stability of the model across all different specifications. Estimating the different models in (log) levels using co-integrated global vector error-correction models (GVECMs) often yields explosive dynamics which makes a meaningful interpretation of the results difficult. Second, given our short sample period, estimations of the long-run relationship between the variables and, hence, the co-integration rank of GVECM have to be treated with caution. Therefore, for our application using a GVECM does not provide insurance against model mis-specification. Furthermore, all the variables are checked for stationarity in a panel unit root test, and are confirmed to be stationary. The results of the test are available in Appendix 1.D.

1.3.4 Implementation of identifying restrictions

To implement the identifying restrictions in the GVAR, we proceed as follows. First, we estimate the model and retrieve its global solution given in Equation (1.8) that includes all endogenous variables across all countries. We then carry out a Cholesky decomposition of the covariance matrix of the residuals of the global model, $\boldsymbol{u}_t$, obtaining a lower triangular matrix $\boldsymbol{P}$. Multiplying this matrix with the corresponding moving-average coefficients of the global solution of the GVAR, $\boldsymbol{\Psi}_h$, $h = 0, 1, 2, ...$, yields the preliminary, standard orthogonalized IRF at horizon h, $\boldsymbol{\Psi}_h \boldsymbol{P}$. Since the effect of a shock identified from a Cholesky decomposition depends on the ordering of the variables, we implement the following concept-based scheme: we order the variables such that real GDP growth for the US and all other countries (in the business cycle model) is followed by US inflation and then, by the US UMP variables (Fed balance sheet and VIX). These variables are then followed by all others, including interest rates, portfolio flows etc.[16]

The IRF obtained from this identification scheme does not yet satisfy the sign restrictions of Table 1.3.1 as the pure Cholesky scheme implies a recursive structure also in the contemporaneous relation between the VIX and the Fed balance sheet. In the next step, we multiply the preliminary IRF by an orthonormal rotation matrix $\boldsymbol{Q}(\theta)$ that takes the following form:

$$\boldsymbol{Q}(\theta) = \begin{pmatrix} \boldsymbol{I}_{m_1} & \boldsymbol{0}_{m_1,2} & \boldsymbol{0}_{m_1,m_2} \\ \boldsymbol{0}_{2,m_1} & \boldsymbol{Q}^{UMP}(\theta) & \boldsymbol{0}_{2,m_2} \\ \boldsymbol{0}_{m_2,m_1} & \boldsymbol{0}_{m_2,2} & \boldsymbol{I}_{m_2} \end{pmatrix}, \tag{1.9}$$

where $\boldsymbol{Q}^{UMP}(\theta) = \begin{pmatrix} \cos(\theta) & -\sin(\theta) \\ \sin(\theta) & \cos(\theta) \end{pmatrix}$.

$\boldsymbol{I}_r$ is an identity matrix of dimension $r \times r$, $\boldsymbol{0}_{r,s}$ is a zero matrix of dimension $r \times s$, and the corresponding dimensions are given by $(m_1, m_2) = (N+1, 2N-1)$ in the BC model and by $(m_1, m_2) = (2, 3N-2)$ in the FC model. The matrix $\boldsymbol{Q}(\theta)$ is thus set up in such a way that $\boldsymbol{Q}^{UMP}(\theta)$ affects the responses of the Fed balance

[16]The specific ordering of the variables after a shock does not have any consequence on their or other variables' responses to this shock.

sheet and the VIX to shocks to these two variables. The impulse response function for each θ is then given by $\text{IRF}(h) = \boldsymbol{\Psi}_h \boldsymbol{P}\boldsymbol{Q}(\theta)$. We draw rotation matrices $\boldsymbol{Q}(\theta)$ by drawing θ from a uniform distribution over the interval $[0, \pi]$ until we obtain the $\text{IRF}(h)$ that satisfies the sign restrictions in Table 1.3.1. The realized UMP shock is then given by the corresponding element in $\boldsymbol{\eta}_t = \boldsymbol{Q}'(\theta)\boldsymbol{P}^{-1}\boldsymbol{u}_t$ that raises the Fed balance sheet and, at the same time, does not increase the VIX.

In principle, there is a variety of models with different rotation matrices $\boldsymbol{Q}(\theta)$ that satisfy the sign restriction. We account for this fact by drawing rotation matrices $\boldsymbol{Q}(\theta)$ until we obtain 1000 admissible IRFs. Then, we apply the "median targeting" approach by Fry and Pagan (2007) and select among admissible models the one that yields IRFs that are closest to the median response across models and horizons. This approach has the advantage that the reported final IRFs are generated by one particular model and that the shocks are orthogonal. For statistical inference, we carry out 500 bootstrap replications of the same setup, with 1000 draws of the rotation matrix in each replication.

1.3.5 Evaluation of portfolio flows as a channel of shock transmission

Even if all the variables in the US and in the EMEs, including portfolio flows, respond to the UMP shock identified above, this is not sufficient evidence that portfolio flows are indeed an important channel of shock transmission in line with Hypotheses 2b and 3b. In order to evaluate the role of portfolio flows in the transmission process, we document their quantitative contribution to the impulse responses of the EME variables by means of the following scenario analysis: given the estimated parameter matrices $\hat{\boldsymbol{G}}_s$, $s = 0, 1, ..., p$, we obtain corresponding matrices $\tilde{\boldsymbol{G}}_s$, $s = 0, 1, ..., p$, where the effects of all variables on both portfolio outflows from the US and portfolio inflows into EMEs are counterfactually set to zero while all other effects are equal to their estimated values. The same scenario is applied to the estimated covariance matrix $\hat{\boldsymbol{\Sigma}}_\varepsilon$ such that we obtain a corresponding matrix $\tilde{\boldsymbol{\Sigma}}_\varepsilon$ where all covariances between the portfolio flow variables and all other variables are counterfactually set to zero while all other variances and covariances are equal to their estimated values.

Using these counterfactual effect matrices in the same algorithm as described in Section 1.3.4, we compute counterfactual impulse response functions $\widetilde{\text{IRF}}(h) = \tilde{\boldsymbol{\Psi}}_h \tilde{\boldsymbol{P}}\boldsymbol{Q}(\theta)$, where $\tilde{\boldsymbol{\Psi}}_h$ are the moving-average coefficients corresponding to $\tilde{\boldsymbol{F}}_s =$

$\tilde{\boldsymbol{G}}_0^{-1}\tilde{\boldsymbol{G}}_s$, $s = 1, ..., p$, and $\tilde{\boldsymbol{P}}$ is the lower triangular Cholesky matrix to $\tilde{\boldsymbol{\Sigma}}_u = \tilde{\boldsymbol{G}}_0^{-1}\tilde{\boldsymbol{\Sigma}}_\varepsilon\tilde{\boldsymbol{G}}_0^{-1\prime}$. Comparing the resulting impulse response functions with the ones obtained from the estimated coefficients uncovers the part of the reaction of each variable that is generated through the responses of portfolio flows. Note that this scenario analysis should not be interpreted as representing alternative outcomes from a policy experiment, but simply as a summary statistic on the magnitudes of estimated coefficients. If the IRFs based on counterfactual matrices showed substantially smaller responses than the original ones, we would infer that portfolio flows constitute an important channel of transmission of US UMP to EMEs' real and financial conditions and monetary policy in accordance with our Hypotheses 2b and 3b. Also note that the counterfactual exercise is conducted using the estimated parameter matrices and, thus, under the assumption of the estimated model being correctly specified.

1.4 Empirical results

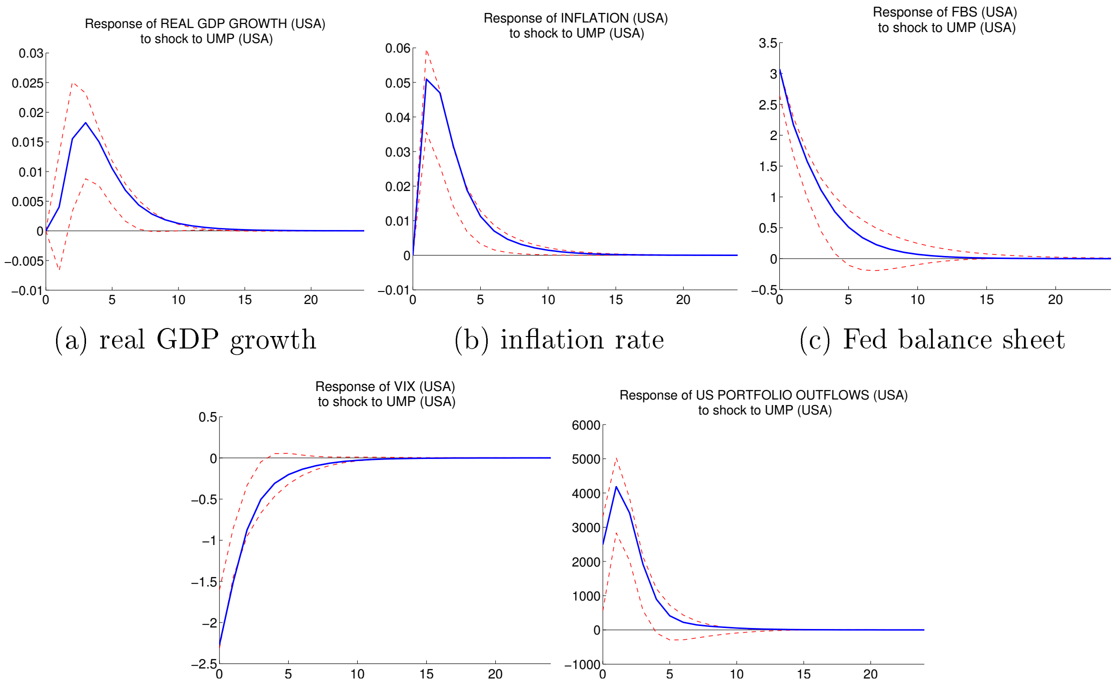
In this section, we first present the main results regarding Hypotheses 1 to 3 that arise from the panel dimension of the sample. We also assess their sensitivity to various alterations of the model. Lastly, we analyze whether and how results from the panel dimension differ with respect to underlying country characteristics.

Before turning to the results, two facts regarding the estimated US UMP shocks have to be noted that follow from a detailed examination of their time series representation (see Appendix 1.B). First, the identified balance sheet shocks capture the monetary policy stance of the Fed over the sample period well. The different phases of QE are, on average, associated with expansionary UMP shocks and ending the respective program of QE implies contractionary impulses (relative to the trend increase). Second, while shocks are overwhelmingly estimated to be positive for QE2 and QE3, this is not the case for QE1 as the increase in the balance sheet for QE1 has been less steep than for the other two programs. Hence, we will bear in mind that expansionary shocks in our specification are mainly associated with QE2 and QE3 when comparing our results to existing evidence on the individual programs.

1.4.1 US reaction to UMP shock

We start by looking at the US part of the estimated GVAR models for a first indication regarding Hypothesis 1 and assess if a UMP shock is related to an increase in capital outflows. IRFs for the US variables to a one standard deviation expansionary UMP shock are depicted in Figure 1.4.1. Exemplarily, we present impulse response functions from the FC model. However, the reaction in the US part of the model is virtually identical in the BC model. The standard deviation shock corresponds to an enlargement of the Fed balance sheet of roughly three percentage points (pp) on impact. As in all the following figures, the solid line represents the median response, and the red dotted lines represent bootstrapped 16% and 84% quantiles.

Figure 1.4.1: Responses of US variables to UMP shock



(a) real GDP growth (b) inflation rate (c) Fed balance sheet

(d) VIX (e) portfolio outflows

Note: The figure shows the estimated impulse responses, along with 68 percent confidence bands, of the US variables to the UMP shock in the financial conditions model. Confidence bands are based on 1000 bootstrap replications with 500 draws of the rotation matrix each.

The response of real GDP growth is estimated to be significantly positive for a period of more than six months, reaching its peak after around four months at an increase of 0.02 pp. The response of inflation is similarly positive, peaks at a 0.05 pp increase, and remains significant for around five months. Moreover, the VIX is significantly reduced for four months after the UMP shocks while the balance sheet is significantly increased for around five months. These results are qualitatively very similar to the ones obtained by Gambacorta et al. (2014) in a comparable setting. However, the magnitude by which GDP and inflation respond is smaller which most likely reflects the larger parameter space of the GVAR. Finally, panel (e) of Figure 1.4.1 shows the response of US portfolio outflows. Following a UMP shock, outflows increase immediately, reaching a peak at around four billion USD after two months. This finding is a first indication in favor of Hypothesis 1. Given how the estimated innovations relate to UMP measures (see above), this result is also in line with, for instance, Fratzscher et al. (2016b), who find that QE2 triggered an immediate re-balancing of assets from the U.S into EMEs.[17]

1.4.2 EMEs' mean response to UMP shock

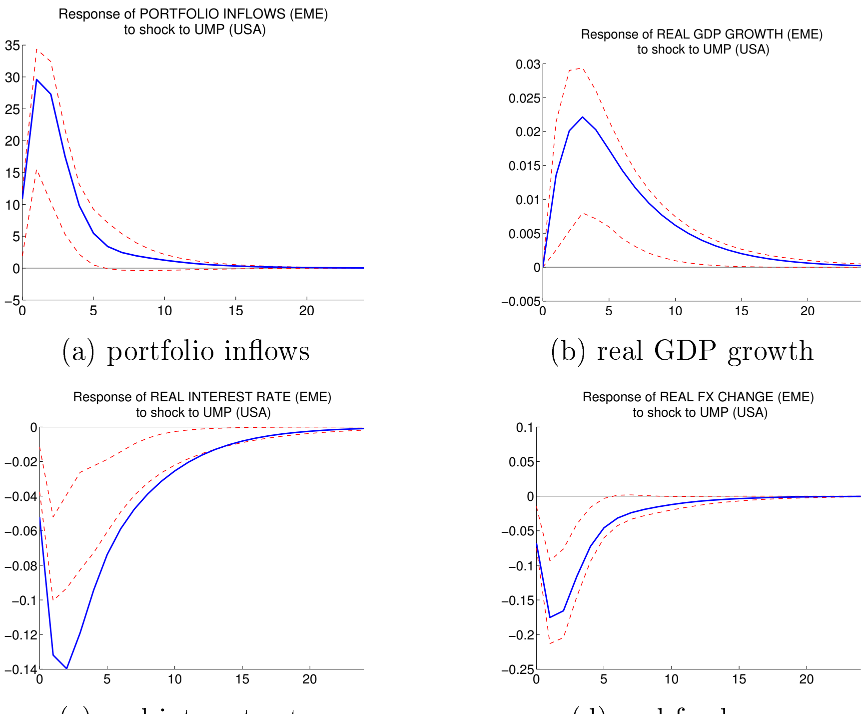
We then examine the average impulse responses of EMEs to the US UMP shock in our two models, starting with the one capturing real economic conditions (see Figure 1.4.2).[18] As panel (a) shows, US portfolio flows to EMEs increase significantly after a UMP shock, consistent with the increase in outflows found in the US part of the model. Regarding the magnitude of the effect, however, the mean increase in inflows has its peak at around 30 million USD. Thus, around 600 million USD are estimated to flow into all EMEs, representing only about 15% of the total response of US outflows. This finding is not necessarily surprising, as the data capture only

[17] Our results show that the effect does not only appear on impact, but also entails a large degree of persistence as the UMP shock leads to a significant increase in portfolio outflows for almost half a year.

[18] The mean impulse response is calculated as the mean of the IRFs across all EME countries except Argentina. Argentina is included in the estimation, but it is excluded from the calculation of the mean because the estimated reaction of the real interest rate is implausibly large. Most likely, this is due to the data capturing the repeated debt restructuring and default that Argentina experienced over the sample period, which was associated with strong movements in both nominal interest rates and inflation. The estimated IRFs for Argentina can be found in Figures 1.C.4 and 1.C.6 in Appendix 1.C. As outlined in Section 1.4.3, all our panel results are robust to dropping Argentina from the sample for the estimation of the model. Figures 1.C.5 and 1.C.7 in Appendix 1.C present the mean response including Argentina. Qualitatively and quantitatively, mean responses for the variables other than the real interest rate are very similar to those computed without Argentina.

flows going directly from the US to EMEs. The data do not account for possibly substantial indirect flows from the US to EMEs through subsidiaries of EME debtors residing in financial centers like the United Kingdom or Hong Kong, where the inflows are indeed estimated to be very large.[19] In this regard, results can be interpreted as a lower bound of actual portfolio inflows. In the model for financial conditions, portfolio flows are also found to increase, showing an almost identical reaction in shape and magnitude (see Figure 1.4.3). Taken together, these results provide support for Hypothesis 1.

Figure 1.4.2: Responses of EME variables to US UMP shock in BC model



(a) portfolio inflows (b) real GDP growth

(c) real interest rate (d) real fx change

Note: The figure shows the estimated mean impulse responses, along with 68 percent confidence bands, of the EME variables to the UMP shock in the business cycle (BC) model. Confidence bands are based on 1000 bootstrap replications with 500 draws of the rotation matrix each.

The other variables in the business-cycle model are also found to respond significantly to a US UMP shock. Real GDP growth in the EMEs increases by around 0.02 percentage points and stays above trend for almost one year. The short-term

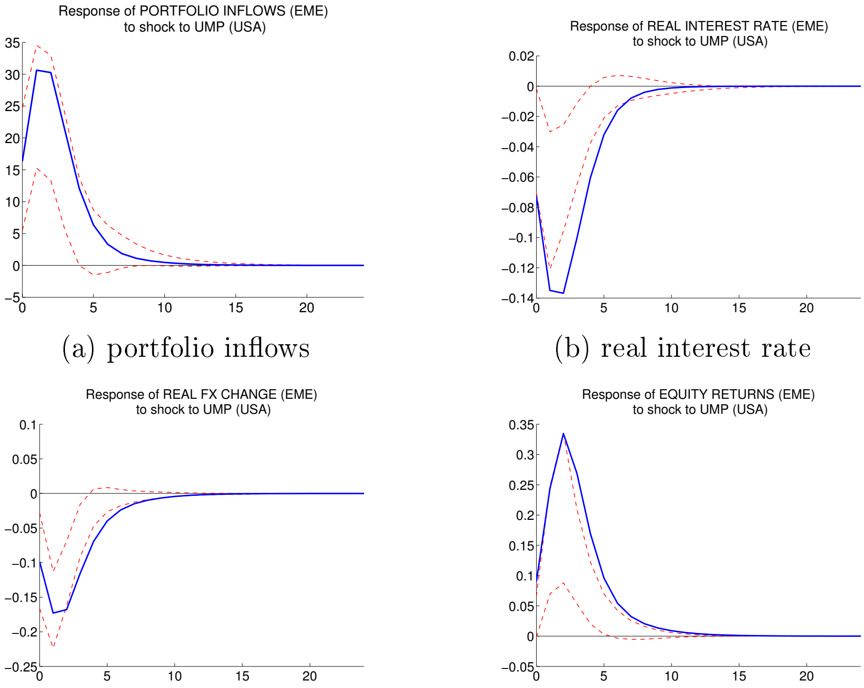
[19] Figure 1.C.9 in Appendix 1.C show that portfolio inflows into the UK and Hong Kong alone amount to roughly 2.2 billion US dollar during the first quarter after the shock.

real interest rate, on average, decreases in response to the expansionary shock. This indicates pro-cyclicality in the monetary policy reaction as interest rates decrease while output increases, and that monetary policy is expansionary in response to an expansionary US impulse, which can be interpreted as first support for Hypothesis 3a. Finally, the response of the real exchange rate shows that EMEs experience a real appreciation of the currency against the US dollar after the UMP shock: the response peaks at an appreciation of around 0.17 pp and persists for roughly 5 months. Hence, one possible explanation for the decrease in the real interest rate is that monetary policy authorities to some extent try to deflect the inflows by lowering the policy rate, in an attempt to avoid an even larger appreciation.

Turning to EMEs financial conditions, Figure 1.4.3 shows the results for the model that includes the responses of portfolio inflows, real exchange rates, real interest rates, and equity returns to an expansionary US UMP shock. Following the shock, portfolio inflows increase in the same magnitude as in the model for the business cycle. Similarly, both the real exchange rate and real interest rate response closely mirror the responses of these variables in the business cycle model. The surge in inflows, the real appreciation, and the lower interest rate are accompanied by a significant increase in equity returns, which lasts for around five months. Hence, these responses suggest that financial conditions, proxied by the exchange rate and equity returns, are indeed affected by an expansionary UMP shock in the US, thus providing support for Hypothesis 2a.

Next, we evaluate the role of portfolio flows as an important channel of shock transmission as implied by Hypotheses 2b and 3b. So far, we have presented results from empirical models where portfolio outflows and foreign output are the transmission channels of US shocks in the GVAR specifications (compare Section 1.3.3). For a first indication on the importance of capital flows in the transmission, we re-estimate the FC model with various alterations regarding the transmission vector $\boldsymbol{y}_{it}^{*}$ for the EMEs. The results of this exercise can be found in Appendix 1.C. First, we drop foreign output from $\boldsymbol{y}_{it}^{*}$. Doing this leaves the IRFs qualitatively and quantitatively almost unchanged compared to the baseline specification, suggesting that foreign output is not the primary transmission channel that drives the results for financial conditions (see Figure 1.C.8 in Appendix 1.C). On the other hand, if we use only foreign output as a transmission channel in $\boldsymbol{y}_{it}^{*}$, spillovers from US policy

Figure 1.4.3: Responses of EME variables to US UMP shock in FC model



(a) portfolio inflows

(b) real interest rate

(c) real fx change

(d) equity returns

Note: The figure shows the estimated mean impulse responses, along with 68 percent confidence bands, of the EME variables to the UMP shock in the financial conditions (FC) model. Confidence bands are based on 1000 bootstrap replications with 500 draws of the rotation matrix each.

shocks are estimated to be considerably smaller. We also add other financial variables to capital flows in $\boldsymbol{y}^*_{it}$ (foreign interest rates, foreign lending rates and foreign equity returns), replacing foreign output one-by-one, to check whether this alters our main results. For the first two alternative specifications, we find that results are qualitatively and also quantitatively similar to the baseline model, suggesting that transmission is already captured by capital flows. Only the inclusion of foreign equity returns produces larger total spillovers from UMP shocks.

This underlines the importance of financial markets in transmitting monetary policy shocks and highlights that portfolio flows are not the only channel of transmission. Overall, the results from the different specifications regarding $\boldsymbol{y}^*_{it}$ are first support for Hypotheses 2b and 3b: capital flows are an important channel of transmission of US UMP shocks.

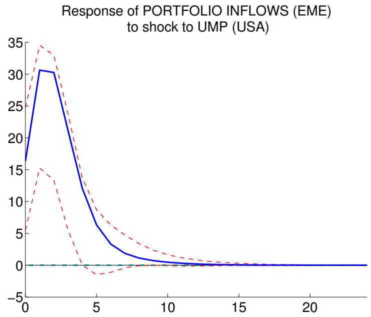
To further investigate this hypothesis, we present a counterfactual exercise from our main specification of model FC where the transmission through capital flows is artificially turned off after estimation (compare Section 1.3.5). In particular, we set all coefficients that load on US portfolio outflows or on EME portfolio inflows to zero, leaving all other coefficients at their estimated values. This exercise should not be interpreted as a policy experiment (e.g. the implementation of capital account restrictions), but simply as a summary statistic on the magnitude of the effects through portfolio flows within the estimated model. Also note that the counterfactual does not aim at isolating the effect going through one of the channels of transmission outlined in Section 2, but captures the overall importance of portfolio flows for the reaction of the variables in the GVAR to the UMP shock. The resulting IRFs from this exercise for the EMEs are presented in Figure 1.4.4. They show that responses of real interest rates, real exchange rate changes and equity returns are substantially smaller compared to the baseline IRFs and thus, underline the quantitative importance of portfolio flows for the transmission of the UMP shock to EMEs.[20] In sum, we interpret the results from this section as providing strong evidence in favor of our Hypothesis 2a and 2b, that US UMP drives financial conditions in EMEs, and that portfolio flows are an important channel of transmission.

1.4.3 Sensitivity to alternative specifications

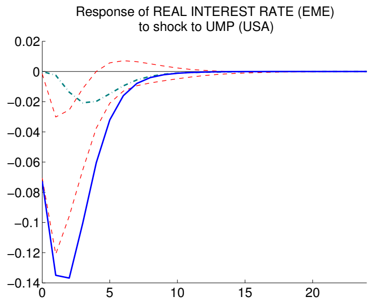
Next, we evaluate the sensitivity of the results from the panel dimension to changes in the specification of the model.[21] We start by assessing the sensitivity of the results to the ordering of variables in the identification scheme. Instead of the concept-based ordering implemented in the baseline specification, we choose a country-based ordering, where the US are placed first. Within the US part of the model, the variables follow the ordering as specified in Section 1.3.3, i.e. output and inflation are followed by the UMP variables, the VIX and the Fed balance sheet. All other variables of the model follow accordingly. This scheme features a stronger isolation of the US from the rest of the world since US monetary policy does not react to any contemporaneous innovations abroad. The qualitative results are robust to this

[20] Feedback effects through portfolio flows to the US economy seem to be small, such that the IRFs for US variables to the UMP shock are basically unaffected, see Figure 1.C.11 in Appendix 1.C.
[21] IRF graphs from this subsection are available in Appendix 1.C.

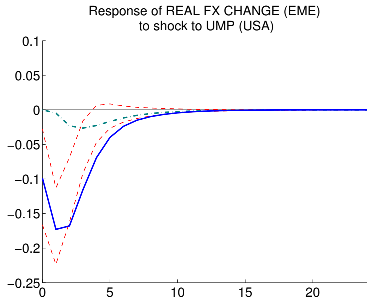
Figure 1.4.4: Counterfactual responses of EME variables to US UMP shock



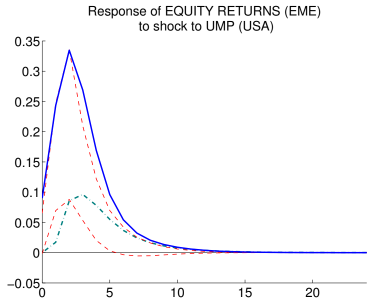
(a) portfolio inflows



(b) real interest rate



(c) real fx change



(d) equity returns

Note: The figure shows the estimated mean impulse responses (blue solid line) and the 68 percent confidence bands (red dashed line) of the EME variables to the UMP shock in the financial conditions (FC) model along with the mean responses in the same model where the transmission through capital flows is counterfactually turned off (green dash-dotted line).

change in ordering. Another alternative ordering scheme that we apply focuses on the timing of portfolio flows. In particular, different to the baseline specification, we try an ordering scheme where all capital flow variables in the model are placed in front of the UMP block. This is done to allow for US monetary policy to react instantaneously to portfolio outflows. On the other hand, this alternative ordering implies that the immediate response of portfolio flows to the UMP shock is zero. Despite the difference in the impact reaction, the overall response of US portfolio outflows and EME portfolio inflows is still very similar to the baseline specification. This holds qualitatively and quantitatively.

Returning to our baseline ordering, we perform a number of additional robustness checks. For instance, we drop several countries one-by-one from the sample that display large reactions in one or more variables and find that mean responses change slightly quantitatively, but not qualitatively. We also perform the estimation using only a post-Lehman sample, i.e. starting in November 2008 instead of January 2008, and find no large qualitative changes resulting from this alteration. Likewise, in the baseline estimation we parsimoniously only include one lag in each country-specific VARX model. As a robustness exercise, we allow for a second or a third lag in the country-specific VARX models based on information criteria. Our findings are very robust to this exercise. A further robustness analysis is to include commodity price inflation as an exogenous (global) variable given that commodity price developments potentially play an important role for several EMEs. The results are qualitatively similar, but the inclusion of this exogenous variable produces smaller - but still significant - reactions of the the EMEs' variables, except for portfolio inflows, which stay almost unchanged.

We also assess whether only one kind of portfolio assets, equity assets or bonds, drives the reaction of flows after a UMP shock. To do so, we replace the portfolio flow variables in our baseline model with equity flows and bond flows, respectively. We find that both equity and bond inflows into EMEs increase, with the rise in bond flows being larger on average. Also, from a cross-country comparison no obvious pattern arises which would indicate the relative importance of one type of flows over the other for single countries.

Further, to assess our identification strategy, we investigate how other US variables react to the UMP shocks in our model and compare the results to evidence from the literature. As found by Wright (2012) and Rogers et al. (2014), US UMP causes a depreciation of the dollar and an increase in US equity prices, among others. Hence, we add those variables separately to the US part of the model and study the estimated response functions. Both responses look as expected given existing evidence, but do not affect the responses of the baseline variables, and thus provide further support for our identification strategy.

As an alternative (unconventional) monetary policy instrument to identify the structural UMP shocks, we use the shadow federal funds rate constructed by Wu and Xia (2016). Specifically, we replace the UMP block, namely the VIX and the balance sheet, in our baseline model with the shadow rate and then apply a recursive ordering scheme to identify the UMP shock. Doing this yields qualitative similar responses for the EME variables as our baseline model with the identification scheme based on Gambacorta et al. (2014).

Lastly, we replace our preferred measures of financial conditions from the main specification one-by-one by other proxies that potentially similarly capture financial developments, and assess whether the reaction of the other variables is robust to this alterations. First, in the model FC, we replace the real interest rate, our proxy for domestic monetary policy, by the real lending rate. The responses of capital inflows, the real exchange rate and equity returns remain unaffected compared to the baseline FC model. The lending rate, on the other hand, displays a reaction very similar to the response of the interest rate in our baseline model, reflecting a close connection of the two rates in our sample. Second, following Bowman et al. (2015), we replace the real interest rate in the model FC by the long-term government bond yield. As before, we find that the reaction of flows, exchange rates, and equity prices is robust to this alteration. The mean responses show that long-term government bond yields significantly decrease, in line with findings by Bowman et al. (2015).

Third, we re-estimate the model FC replacing the interest rate by foreign exchange reserves growth. We do so as the accumulation of reserves is a policy tool actively used by central banks in EMEs, for instance to alleviate exchange rate appreciation pressures. Again we find that the reaction of portfolio flows, equity returns, and exchange rates is robust to this alteration. Moreover, reserve accumulation increases in response to the UMP shock. Similar to the real interest rate response in the baseline model, this indicates that monetary policy in EMEs reacts to the US expansion. In particular, the increase in reserves could be related to policies aimed at mitigating a currency appreciation that arises from the capital inflows.

1.4.4 The role of economic characteristics of countries

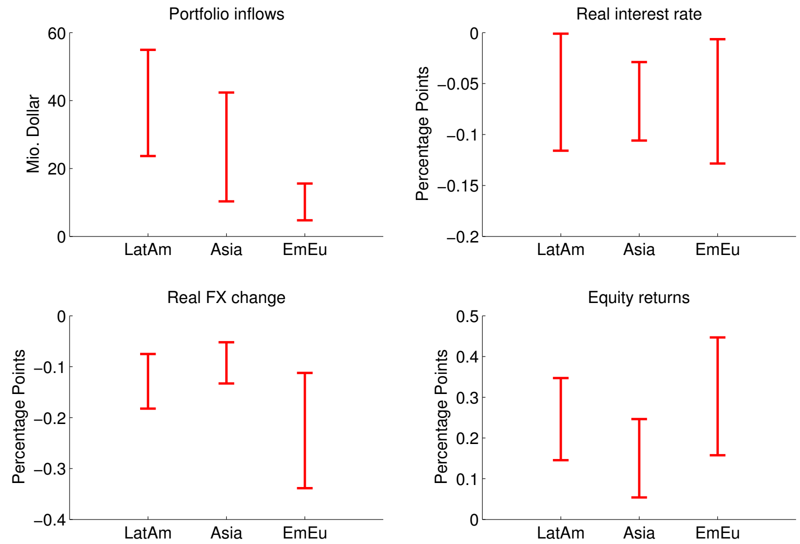
Finally, we go into detail and analyze if and to what extent EMEs are affected differently by the US UMP shock. For this purpose, we split our sample of countries in different ways according to economic characteristics and analyze whether the responses of portfolio flows, domestic financial conditions, and monetary policy rates to the shock differ between groups. In particular, we look at differences in the estimated peak responses of the variables from model FC across groups with different country characteristics. The differences in peak responses correspond very closely to the differences in cumulated IRFs for the country groups.

First, we consider the role of country geography. US capital flows towards Europe, for instance, may not be as large as towards Latin America, since the latter has tighter economic and geographical linkages to the US The EMEs we analyze are grouped by geographic region as follows: Latin America (Brazil, Chile, Colombia, Mexico, Peru), Europe (Hungary, Poland, Russia, Turkey) and Asia (India, Indonesia, Korea, Malaysia, the Philippines, Singapore, Thailand). 68% confidence bands of estimated peak effects in response to the US UMP shock for the different regions are displayed in Figure 1.4.5. The full set of responses are depicted in Figures 1.C.1 – 1.C.3 in Appendix 1.C.

We indeed observe the strongest reaction of capital inflows after an expansionary shock for Latin American countries, whereas the smallest reaction is found for European countries. In principle, comparisons of total flows have to be treated with caution as they do not take into account that groups may vary by economic size. The group of Latin American countries and the group of European countries, however, are of similar size regarding their total GDP, while GDP of the Asian countries is even larger. The larger response of portfolio flows to Latin America compared to Asia is reflected in slightly stronger responses of equity returns and exchange rates. On the other hand, European interest rates and equity returns react with similar magnitude as their American counterparts whereas the foreign exchange reaction is even more pronounced. This can be explained by effects being enhanced by close trade linkages to advanced European countries, and in particular by currencies being tied to the euro. In sum, there is evidence for limited regional heterogeneity in quantitative respect, indicating that the capital flow channel is of particular importance for Latin America, whereas for Asia and Emerging Europe also other channels play an important role. Qualitatively, however, responses across regions are very similar.

Next, we study the role of institutional quality, as Fratzscher et al. (2016b) find evidence that countries with better institutions are less affected by Fed policies. We follow their approach and proxy the institutional quality by the 2007 average of the following four indicators of governance: 'Political Stability', 'Rule of Law', 'Control of Corruption' and 'Regulatory Quality'; all from Kaufmann et al. (2010). The country characteristics are predetermined to account for the possibility that they

Figure 1.4.5: Peak responses across geographic regions

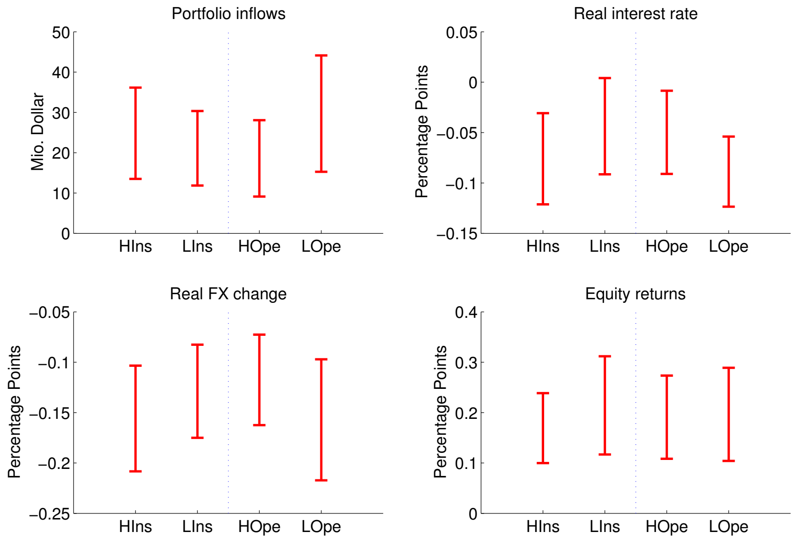


Note: The figure shows 68 percent confidence bands of estimated peak effects for the model FC across different geographical regions. LatAm: Latin America, EmEu: Emerging Europe.

might be contemporaneously related to US UMP. Then, we split our sample into two groups, one with institutional quality above the cross-country median and one with institutional quality below the cross-country median. Figure 1.4.6 (left half of each graph) shows that peak responses are very similar for both groups of countries. Hence, we do not find evidence that the EMEs in our sample are affected differently due to differences in institutional quality. The difference to the results by Fratzscher et al. (2016b) can be explained by the fact that their sample also contains advanced economies, who – on average – have a higher institutional quality than EMEs.

We also analyze whether countries with a lower degree of financial openness are better insulated from US UMP shocks. Financial openness is proxied by the Chinn-Ito coefficient for capital account openness (Chinn and Ito, 2006). Again, we use the 2007 value of the country characteristic to avoid problems of endogeneity and split our sample into two groups, one with financial openness above the median and one with openness below. Peak responses for the two groups are displayed in Figure 1.4.6 (right half of each graph). There is no evidence that countries with a higher

Figure 1.4.6: Peak responses across country groups with different degree of institutional quality / openness



Note: The figure shows 68 percent confidence bands of estimated peak effects for the model FC across country groups with different characteristics. HIns: institutional quality above cross-country median, LIns: institutional quality below cross-country median, HOpe: financial openness above cross-country median, LOpe: financial openness below cross-country median.

degree of financial openness at the start of our sample are more strongly affected by the UMP shock.[22]

Finally, we study how responses to the UMP shock differ with respect to the countries' exchange rate arrangements. This is motivated by the classical "trilemma", which states that countries can only have two of the following three policy options: independent monetary policy, free capital flows, or a fixed exchange rate. Hence, countries that let their currency float freely might have potentially more leeway for an independent monetary policy, meaning that they can set interest rates independently of the US and, in turn, achieve a better buffering of US shocks. Empirical studies often find evidence for this kind of reasoning (see Klein and Shambaugh, 2015, or Aizenman et al., 2016b). However, Rey (2013) argues that the trilemma

[22]It is important to keep in mind that this result cannot be interpreted as a test of how efficient the countries' micro- and macroprudential measures were during and after the crisis. As the introduction of measures like capital controls is highly endogenous to the degree of capital inflows following a monetary policy shock, this question cannot be addressed in the current framework and constitutes an important avenue for future research.

is often reduced to a dilemma: a floating exchange rate regime alone might not be sufficient to insulate countries from the global financial cycle and US monetary policy as one of its drivers.

Ideally, we would thus split our sample countries in those that are situated in a floating exchange rate regime ("floaters") and those that fix there exchange rate ("peggers"). This, however, is not a straightforward task given our sample period in which most of the countries are *de jure* considered to be floaters. Therefore, we turn to Klein and Shambaugh (2008)'s *de facto* classification of countries' prevailing exchange rate regimes for this task, which is based on observable exchange rate variation and is commonly used in the literature. Following the strategy outlined above, we split our sample into two groups, one with exchange rate flexibility above the cross-country median and one with exchange rate flexibility below the cross-country median. We measure exchange rate flexibility in two ways, namely by the number of years that a country was considered a floater in the years 2003–2007 and 2008–2014, respectively. While the latter classification potentially suffers from an endogeneity bias between exchange rate stability and the amount of capital inflows, it serves as a useful benchmark for the results based on pre-crisis data. Both classifications yield similar, but slightly different groups of countries.[23]

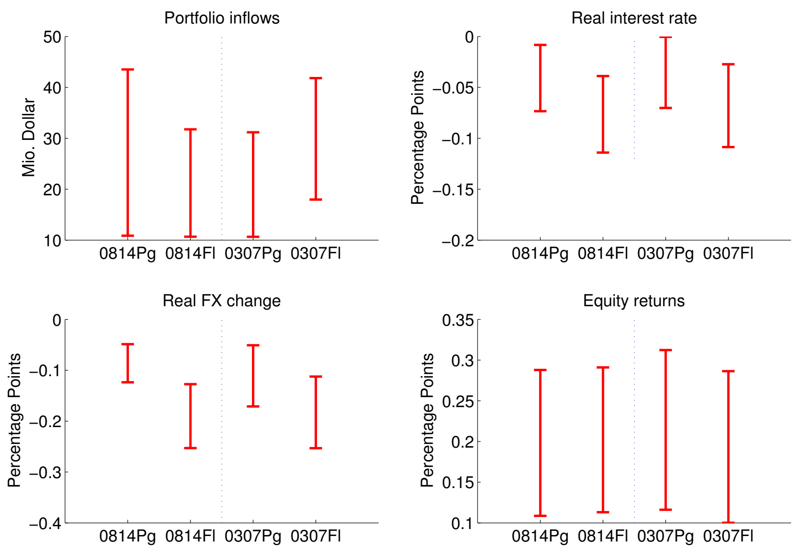
Figure 1.4.7 shows the peak responses for the variables in model FC for the group of floaters (above median) and the group of peggers (below median) for both classifications. It indicates that, despite small changes in the group compositions, results between the two classifications are very similar. Both floaters and peggers receive capital inflows of similar magnitude and also experience a similar increase in asset prices. Moreover, currencies in both groups are found to appreciate after the shock, reflecting the fact that the group of peggers does not constitute hard pegs in our sample. Not surprisingly, the appreciation is stronger in countries that float freely. Of particular interest in the sense of the trilemma is the reaction of monetary policy

[23] Different to the other country characteristics, we do not only use one year of data to group the countries since *de facto* exchange rate classifications are far less persistent than the degree of institutional quality of financial openness considered above. The classification of Klein and Shambaugh (2008) can be found on Jay C. Shambaugh's homepage. As alternative specifications we have also classified countries using only their 2007 observations, which yielded qualitatively similar results. As an alternative dataset, we have used the coarse classification by Reinhart and Rogoff (2004) and applied a similar classification strategy. The resulting groups are slightly different, but the results regarding the responses to the policy shock in both country groups are very similar.

in both groups as floating countries should experience a higher degree of policy independence and thus, should display less of a mirroring interest rate reaction to the expansionary US shock. However, we find that both groups react to the expansionary shock by decreasing their own monetary policy rates, indicating a procyclical reaction of monetary policy as found in the panel results.

Together with the evidence obtained in Section 1.4.2 (Figure 1.4.4 panel b)) on the importance of portfolio flows as a transmission channel for the UMP shock, we interpret this result as strongly supporting our Hypothesis 3: via portfolio flows, US UMP significantly influences monetary policy in EMEs.

Figure 1.4.7: Peak responses across country groups with different exchange rate flexibility



Note: The figure shows 68 percent confidence bands of estimated peak effects for the model FC across country groups with different degrees of *de facto* exchange rate flexibility. 0814Pg: flexibility in 2008–2014 below cross-country median, 0814Fl: flexibility in 2008–2014 above cross-country median, 0307Pg: flexibility in 2003–2007 below cross-country median, 0307Fl: flexibility in 2003–2007 above cross-country median.

Furthermore, our results imply that a floating exchange rate regime alone is not sufficient to insulate the EMEs in our model from the US UMP shock. This is in line with the findings of Passari and Rey (2015) and Rey (2016) for a group of small open advanced economies and US interest rate surprises. However, our results are not necessarily at odds with existing evidence that stresses the importance of exchange rate regimes and the classical trilemma. These studies often cover a broader range of shocks, countries as well as time periods, and allow for a more detailed analysis on how the varying degrees of exchange rate flexibility and also capital control intensity are related to the reaction of domestic monetary policy.

There are a number of caveats to the analysis offered in this subsection that need to be mentioned. Our results most likely constitute lower bounds for the capital inflows into EMEs after a UMP shock as they do not take the issuance of debt through subsidiaries into financial centers into account (compare Section 1.3.1). The likelihood that corporations in a country issue debt through an offshore subsidiary could be related to the underlying country characteristics, in particular the exchange rate regime or the degree of financial openness. Should this be the case, results for the individual groups could differ if it would be possible to account for offshore subsidiaries. Moreover, it is important to note that the different country characteristics might not be independent from each other. Lastly, due to a lack of data availability, the influence of other micro- and macroprudential measures that were taken during the crisis is not taken into account. Nevertheless, overall our results are indicative that neither a floating exchange rate nor a higher degree of institutional quality or a lower degree of financial openness are sufficient to insulate countries from the spillovers of US UMP.

1.5 Conclusion

In this paper, we investigate empirically whether US unconventional monetary policy has an impact on financial and real variables in emerging market economies, and examine whether portfolio flows are an important channel of transmission. In contrast to existing studies, we use a structural global vector autoregressive approach that takes economic interlinkages between countries and across time into account and allows to assess the persistence of the effects of US UMP shocks.

We find that an expansionary UMP shock significantly increases portfolio outflows from the US for almost six months. In the EMEs, this is equivalently associated with portfolio inflows. Along with the increase in inflows, real output growth and equity returns rise, the real exchange rate appreciates and the real lending rate decreases. Importantly, portfolio flows prove to be an important channel of transmission in the GVAR specification. We also find that EMEs, on average, react pro-cyclically by decreasing their short-term interest rate in response to the US shock, indicating a monetary policy response that mirrors the expansionary US impulse. All our findings appear to be independent of economic characteristics like the degree of financial openness, institutional quality, or the underlying exchange rate regime of a country.

Our results complement existing evidence along different dimensions. Fratzscher et al. (2016b), for instance, find that UMP in the US had a direct effect on portfolio reallocation between advanced economies and EMEs. We show that US shocks have persistent effects on portfolio flows, and that these flows are a channel of transmission to EMEs. Miranda-Agrippino and Rey (2015) find that US monetary policy had an effect on financial conditions worldwide for the period 1980–2010. We obtain a comparable result for the particular period of unconventional monetary policy. Lastly, Passari and Rey (2015) and Rey (2016) show for a group of open advanced economies that countries with a flexible exchange rate regime are not insulated from US monetary policy. We document that also EMEs with a flexible exchange rate arrangement are affected by US UMP shocks.

Finally, some limitations of our analysis are particularly interesting for future research. Given our identification strategy and sample period, we did not analyze the effects of uncertainty about the end of UMP in the US (for instance, the so-called "taper tantrum"). Future research might explore how such policy uncertainty affects portfolio flows between the US and EMEs. Furthermore, the dataset on capital flows does not take potential flows through subsidiaries in financial centers into account. It might be interesting to study in future research whether the likelihood of issuing bond and equity assets by EME firms through offshore subsidiaries is related to cross-country characteristics, like the degree of financial openness or the exchange rate regime.

Appendix

1.A Data and sources

Table 1.A.1: Data construction and sources

Variable	Construction and source
Inflation	First difference of log of monthly consumer price index (CPI), in percent. Source: Datastream (US: St. Louis FRED).
Equity return	First difference of log MSCI equity, in percent. Source: Datastream (US CPI: St. Louis FRED).
Real output	Monthly real GDP is obtained by interpolating quarterly log real GDP with log index of industrial production using the method of Chow and Lin (1971). Real GDP growth is first difference of log real GDP, in percent. Source: Datastream (US: St. Louis FRED).
Fed balance sheet	Sum of assets held by Federal Reserve System. Source: St. Louis FRED.
VIX	Option-implied volatility index. Source: Chicago Board Options Exchange, retrieved from St. Louis FRED.
Real exchange rate	Calculated from log nominal US dollar exchange rate e_t using the following formula: $e_t + \ln \mathrm{CPI}_t^{US} - \ln \mathrm{CPI}_t$. Real exchange rate change is first difference of log real exchange rate, in percent. Source: Datastream (US CPI: St. Louis FRED).
Real interest rate	Calculated from nominal short-term interest rate i_t using the following formula: $i_t - 12 \cdot \Delta \ln \mathrm{CPI}_t \cdot 100$. As nominal short-term rate, we choose the monetary policy rate (exact definition depends on policy measures of the respective country ("target rate", "policy rate", ...). Source: Datastream (US CPI: St. Louis FRED).
Real lending rate	Calculated from nominal lending rate l_t using the following formula: $l_t - 12 \cdot \Delta \ln \mathrm{CPI}_t \cdot 100$. Source: IMF International Financial Statistics (IFS), (US CPI: St. Louis FRED).
Real effective exchange rate	Log of index of real effective exchange rate. Change is first difference, in percent. Source: Bank of International Settlements, retrieved from Datastream.
Foreign exchange reserves	First difference of log foreign exchange reserves, in percent. Source: IFS.

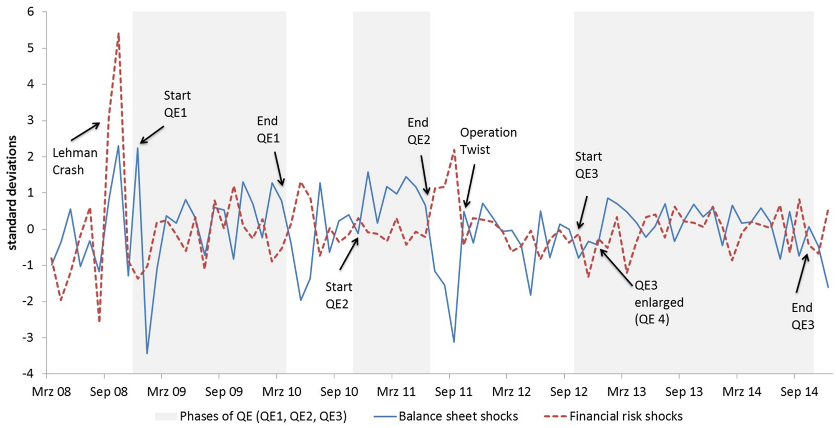
Government bond yield	Source: Datastream.
Commodity price inflation	First difference of log commodity price index, in percent. Source: IFS.

1.B Identified exogenous balance sheet innovations

For a better interpretation of the identified US UMP shocks, we examine their time series representation. This inspection allows us to assess whether the major policy measures taken by the Fed during and after the financial crisis (see Section 3.3 in the paper) are reflected in the estimated innovations. Given that most of the measures have to some extent an unexpected component, this is a useful check of the identification approach. In doing so, we follow Boeckx et al. (2017) who apply a similar balance sheet driven identification strategy for UMP by the ECB.

Figure 1.B.1 shows the time series of the identified UMP innovations in the balance sheet as well as the time series of identified financial market risk shocks. The phases of Quantitative Easing (QE) are depicted as shaded areas.[24] The sum of the shocks is, by construction as a white noise residual, zero over the sample period and the scale is standard deviations. A positive shock reflects an expansionary shock while a negative impulse is associated with a tightening of the balance sheet relative to the average endogenous response to other shocks hitting the economy.

Figure 1.B.1: Time series of balance sheet shocks and financial market risk shocks



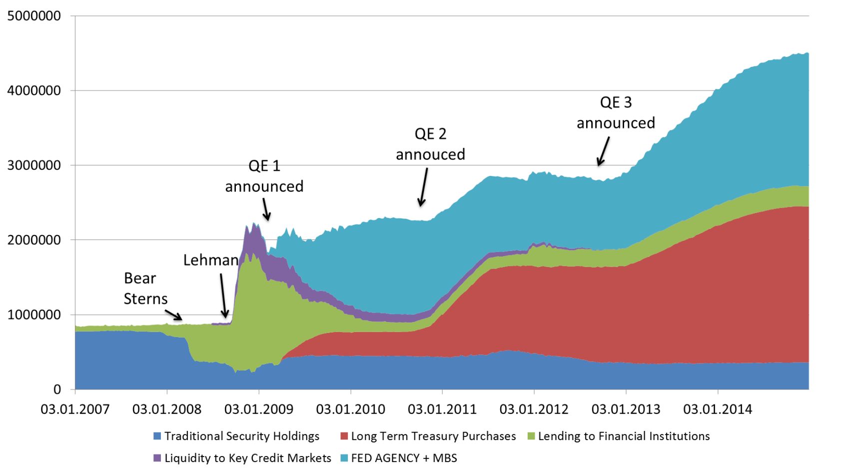
Note: The figure shows the estimated unconventional monetary policy shock, reflecting an unexpected extension of the balance sheet, and financial market risk shock, capturing financial market turbulence to which the Fed responds, as well as phases of quantitative easing (shaded areas) and important monetary policy decisions over the sample period. The scale is standard deviations.

[24] The shock series shown in the figure correspond to the model that yields the median IRF in the baseline specification with equity prices, i.e. the financial conditions model. The shocks, however, are virtually identical for the two specifications.

The figure shows that the identified balance sheet shocks capture the monetary policy stance of the Fed over the sample period well. The different phases of QE are, on average, associated with expansionary UMP shocks. On the other hand, ending the respective program of QE implies contractionary impulses. This is in line with the notion that the first two programs were ended despite no major improvement in economic conditions and financial stability. Nevertheless, there are differences between the identified UMP shocks over the different phases of QE. While shocks are overwhelmingly estimated to be positive for QE2 and QE3 (after its enlargement in December 2012), this is not the case for QE1. The main reasons for this is that the increase in the balance sheet for QE1 has been less steep than for the other two programs (see Figure 1.B.2). Hence, the estimated policy reaction function in the GVAR, which also includes a linear trend, perceives QE1 as less expansionary than QE2 and QE3. This result has to be borne in mind when comparing our results to existing evidence on the individual programs.

The financial market risk shock, in contrast, most notably captures the turmoil associated with the collapse of Lehman Brothers in 2008. Hence, most of the Fed balance sheet enlargement directly after Lehman is regarded by the model as an endogenous reaction to the collapse and not an exogenous policy innovation. This reflects the identification strategy which exactly aims at disentangling an endogenous reaction to financial market turmoil from expansionary policy.

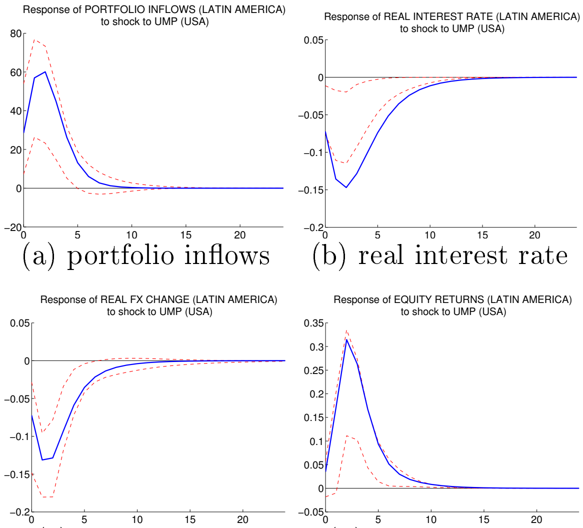
Figure 1.B.2: Federal Reserve assets by type, 2007–2014



Note: In Million of US dollars. Source: Federal Reserve Bank of Cleveland

1.C Additional figures

Figure 1.C.1: Responses of Latin America's variables to US UMP shock – model: FC

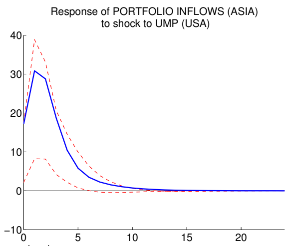


(a) portfolio inflows (b) real interest rate

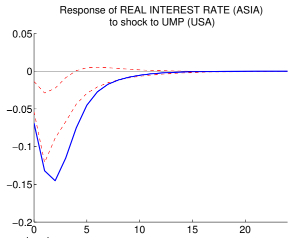
(c) real fx change (d) equity returns

Note: The figure shows the estimated impulse responses, along with 68 percent confidence bands, of the depicted variables to a one standard deviation UMP shock in the FC model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

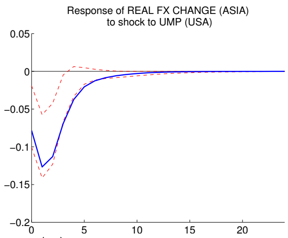
Figure 1.C.2: Responses of Asia's variables to US UMP shock – model: FC



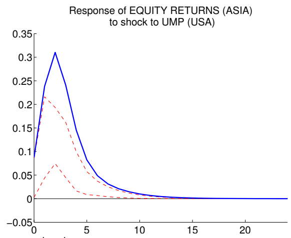
(a) portfolio inflows



(b) real interest rate



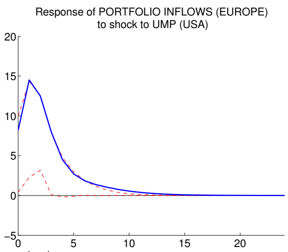
(c) real fx change



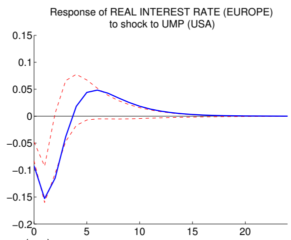
(d) equity returns

Note: The figure shows the estimated impulse responses, along with 68 percent confidence bands, of the depicted variables to a one standard deviation UMP shock in the FC model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

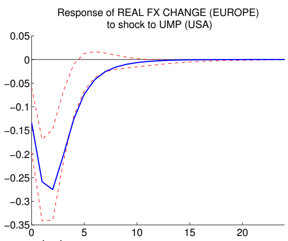
Figure 1.C.3: Responses of Europe's variables to US UMP shock – model: FC



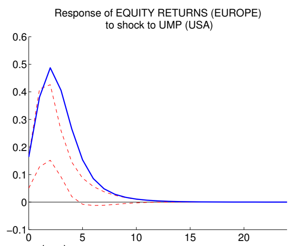
(a) portfolio inflows



(b) real interest rate



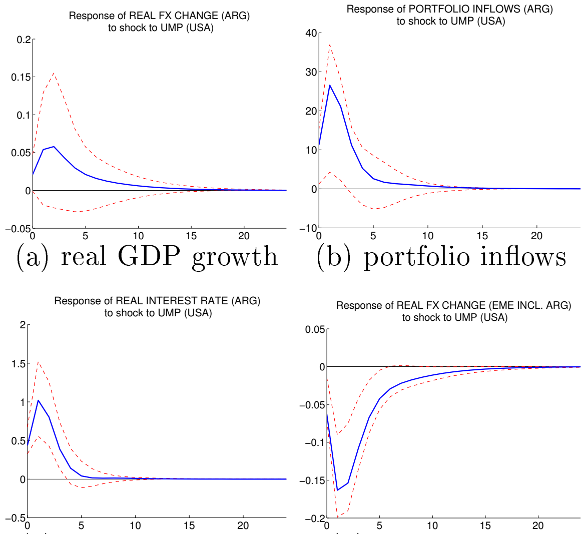
(c) real fx change



(d) equity returns

Note: The figure shows the estimated impulse responses, along with 68 percent confidence bands, of the depicted variables to a one standard deviation UMP shock in the FC model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

Figure 1.C.4: Responses of Argentina's variables to US UMP shock – model: BC

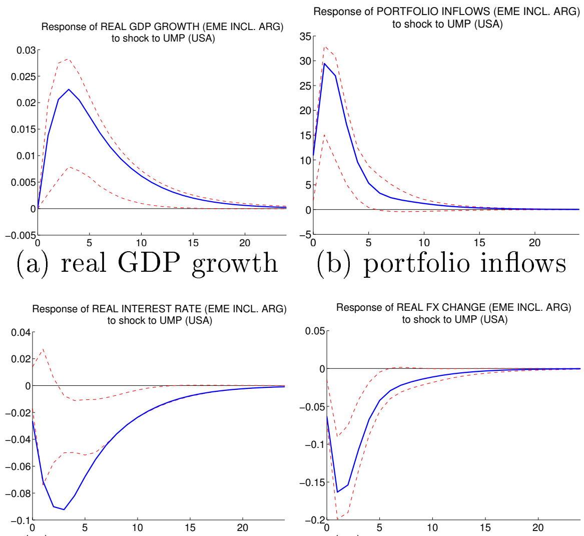


(a) real GDP growth (b) portfolio inflows

(c) real interest rate (d) real fx change

Note: The figure shows the estimated impulse responses, along with 68 percent confidence bands, of the depicted variables to a one standard deviation UMP shock in the BC model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

Figure 1.C.5: Responses of EME variables (mean including Argentina) to US UMP shock – model: BC

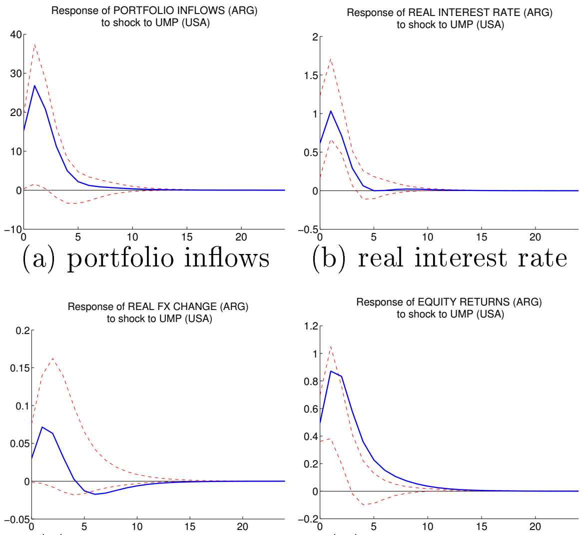


(a) real GDP growth (b) portfolio inflows

(c) real interest rate (d) real fx change

Note: The figure shows the estimated impulse responses, along with 68 percent confidence bands, of the depicted variables to a one standard deviation UMP shock in the BC model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

Figure 1.C.6: Responses of Argentina's variables to US UMP shock – model: FC

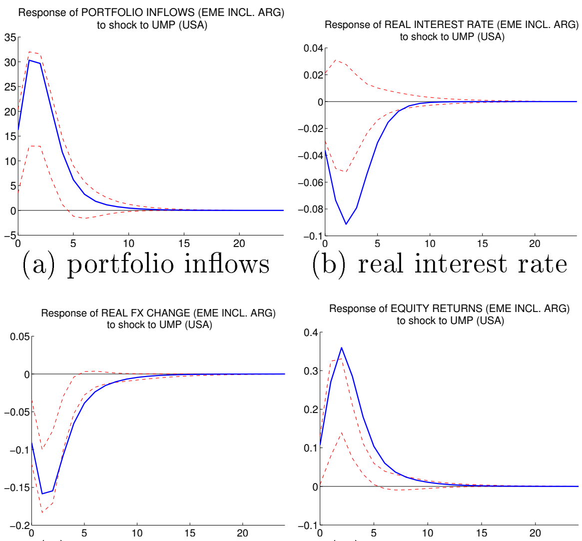


(a) portfolio inflows (b) real interest rate

(c) real fx change (d) equity returns

Note: The figure shows the estimated impulse responses, along with 68 percent confidence bands, of the depicted variables to a one standard deviation UMP shock in the FC model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

Figure 1.C.7: Responses of EME variables (mean including Argentina) to US UMP shock – model: FC

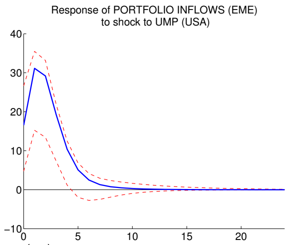


(a) portfolio inflows (b) real interest rate

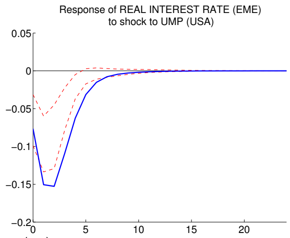
(c) real fx change (d) equity returns

Note: The figure shows the estimated impulse responses, along with 68 percent confidence bands, of the depicted variables to a one standard deviation UMP shock in the financial conditions (FC) model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

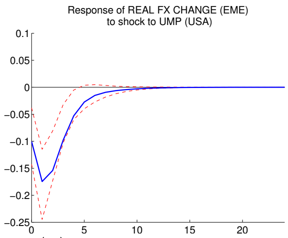
Figure 1.C.8: Responses of EME variables (mean) to US UMP shock – model: FC with portfolio flows as only cross-country transmission channel



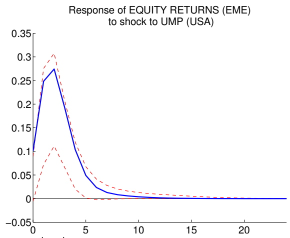
(a) portfolio inflows



(b) real interest rate



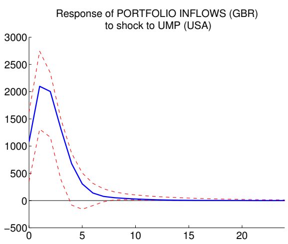
(c) real fx change



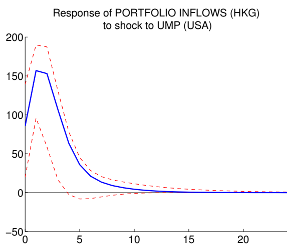
(d) equity returns

Note: The figure shows the estimated impulse responses, along with 68 percent confidence bands, of the depicted variables to a one standard deviation UMP shock in the financial conditions (FC) model with portfolio flows as the only foreign variable. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

Figure 1.C.9: Responses of portfolio inflows to the United Kingdom and to Hong Kong – model: FC



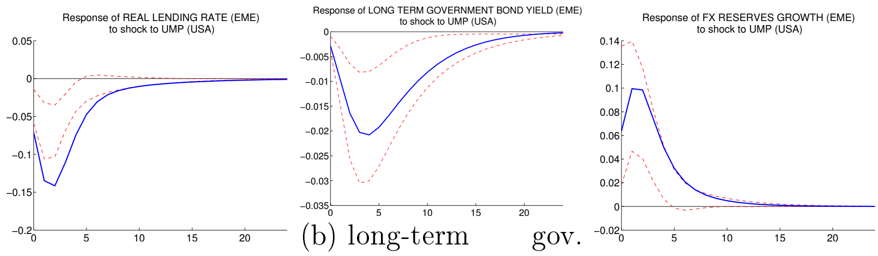
(a) portfolio inflows - United Kingdom



(b) portfolio inflows - Hong Kong

Note: The figure shows the estimated impulse responses, along with 68 percent confidence bands, of the depicted variables to a one standard deviation UMP shock in the financial conditions (FC) model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

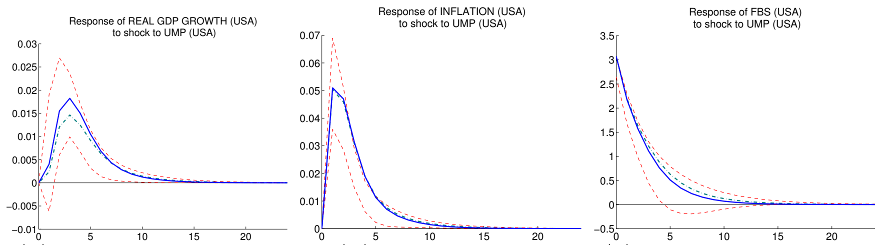
Figure 1.C.10: Responses of alternative measures of financial conditions



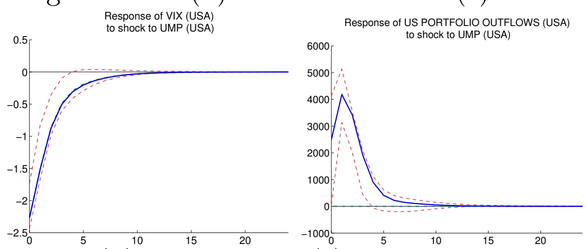
(a) real lending rate (b) long-term gov. bond yield (c) fx reserve growth

Note: The figure shows the estimated impulse responses, along with 68 percent confidence bands, of the depicted variables to a one standard deviation UMP shock in the financial conditions (FC) model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

Figure 1.C.11: Responses of US variables to US UMP shock – model: FC, counterfactual exercise with restriction to portfolio flows



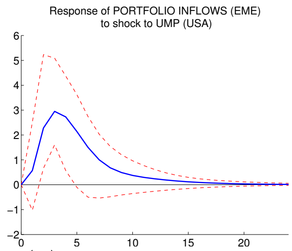
(a) real GDP growth (b) inflation rate (c) Fed balance sheet



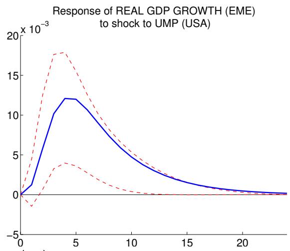
(d) VIX (e) portfolio outflows

Note: The figure shows the estimated impulse responses (blue solid line) and the 68 percent confidence bands (red dashed line) of the EME variables to the UMP shock in the financial conditions (FC) model of the depicted variables to a one standard deviation UMP shock in the financial conditions (FC) model, along with the corresponding impulse responses in the same model where the transmission through capital flows is counterfactually turned off (green dash-dotted line, see Section 1.3.5).

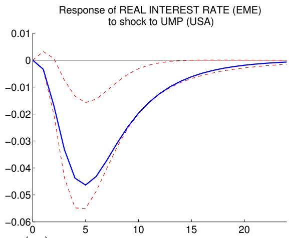
Figure 1.C.12: Responses of EME variables (mean) to US UMP shock - model: BC with GDP as only foreign variable



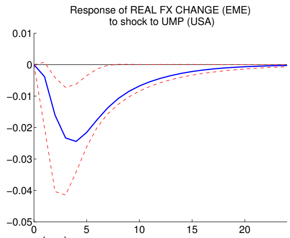
(a) portfolio inflows



(b) real GDP growth



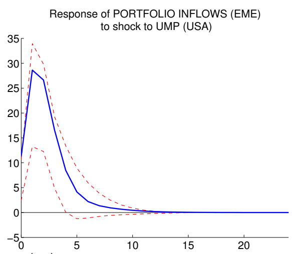
(c) real interest rate



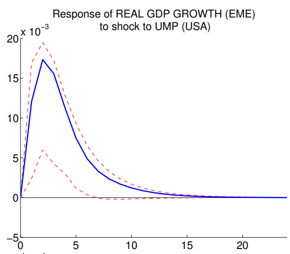
(d) real fx change

Note: The figure shows the estimated mean impulse responses, along with 68 percent confidence bands, of the EME variables to the UMP shock in the business cycle (BC) model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

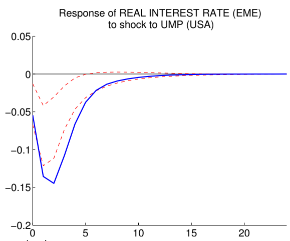
Figure 1.C.13: Responses of EME variables (mean) to US UMP shock - model: BC model with portfolio flows as only foreign variable



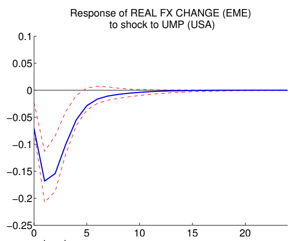
(a) portfolio inflows



(b) real GDP growth



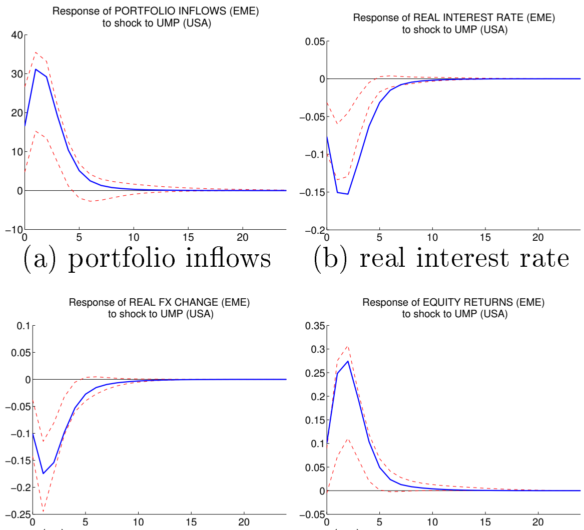
(c) real interest rate



(d) real fx change

Note: The figure shows the estimated mean impulse responses, along with 68 percent confidence bands, of the EME variables to the UMP shock in the business cycle (BC) model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

Figure 1.C.14: Responses of EME variables (mean) to US UMP shock - model: FC with portfolio flows as only foreign variable

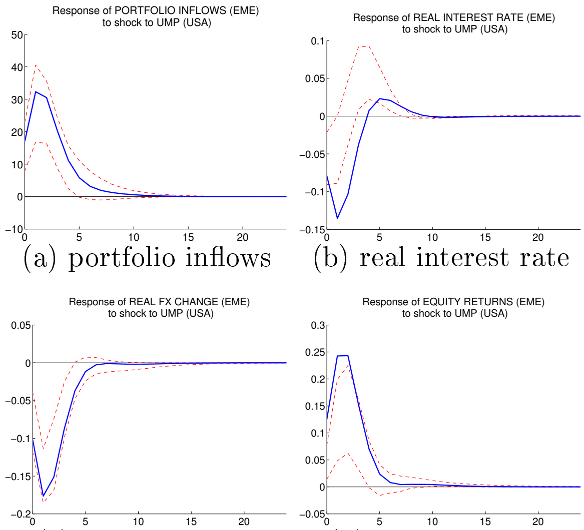


(a) portfolio inflows (b) real interest rate

(c) real fx change (d) equity returns

Note: The figure shows the estimated mean impulse responses, along with 68 percent confidence bands, of the EME variables to the UMP shock in the financial conditions (FC) model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

Figure 1.C.15: Responses of EME variables (mean) to US UMP shock - model: FC with portfolio flows and foreign interest rates as foreign variables

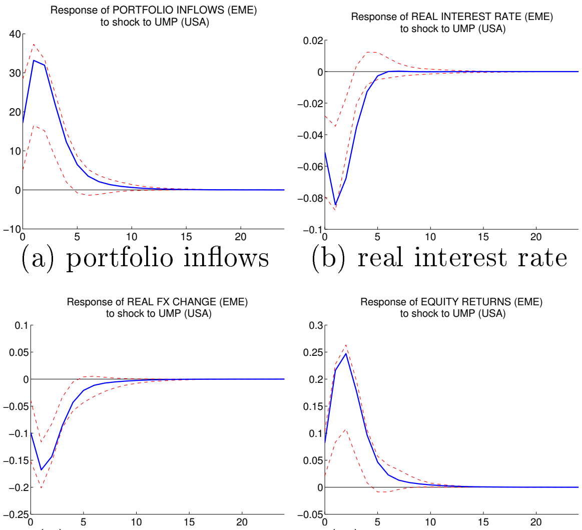


(a) portfolio inflows (b) real interest rate

(c) real fx change (d) equity returns

Note: The figure shows the estimated mean impulse responses, along with 68 percent confidence bands, of the EME variables to the UMP shock in the financial conditions (FC) model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

Figure 1.C.16: Responses of EME variables (mean) to US UMP shock - model: FC with portfolio flows and foreign lending rates as foreign variables

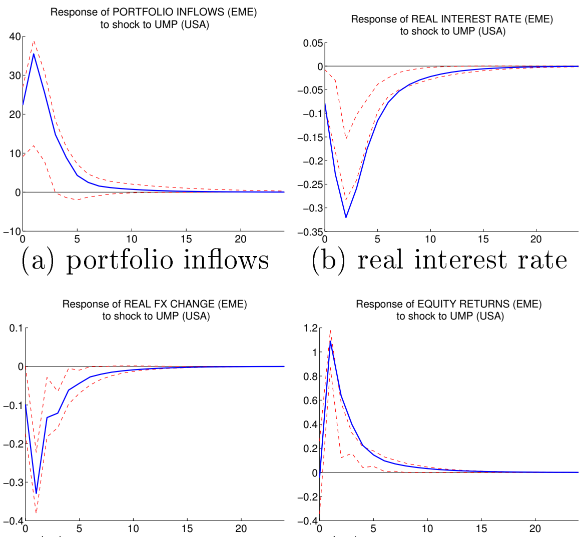


(a) portfolio inflows (b) real interest rate

(c) real fx change (d) equity returns

Note: The figure shows the estimated mean impulse responses, along with 68 percent confidence bands, of the EME variables to the UMP shock in the financial conditions (FC) model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

Figure 1.C.17: Responses of EME variables (mean) to US UMP shock - model: FC with portfolio flows and foreign equity returns as foreign variables

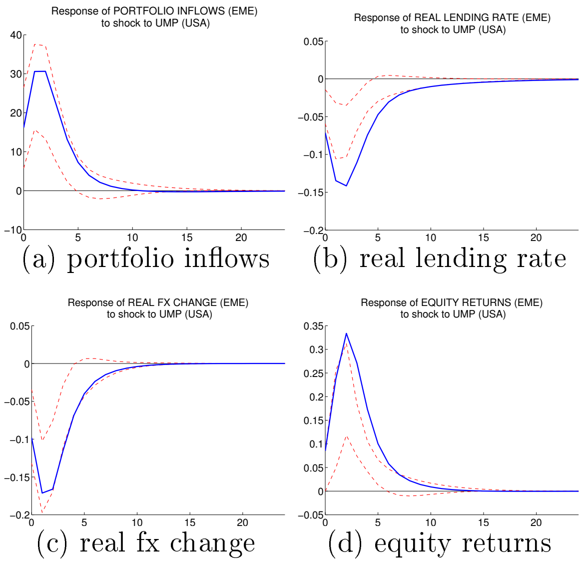


(a) portfolio inflows (b) real interest rate

(c) real fx change (d) equity returns

Note: The figure shows the estimated mean impulse responses, along with 68 percent confidence bands, of the EME variables to the UMP shock in the financial conditions (FC) model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

Figure 1.C.18: Responses of EME variables (mean) to US UMP shock - model: FC with lending rate

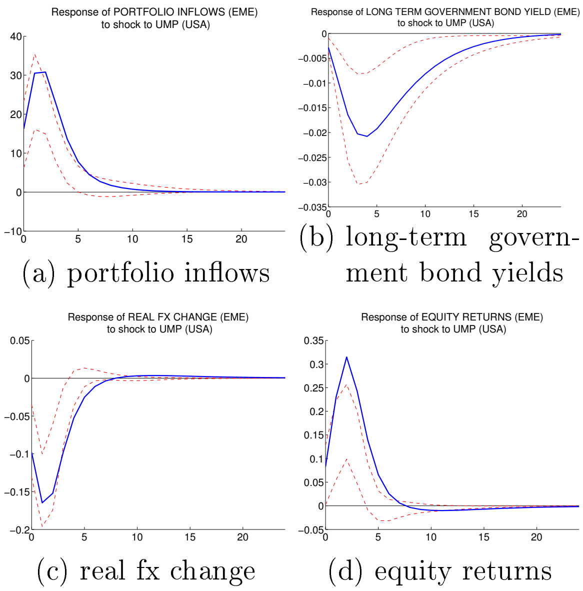


(a) portfolio inflows (b) real lending rate

(c) real fx change (d) equity returns

Note: The figure shows the estimated mean impulse responses, along with 68 percent confidence bands, of the EME variables to the UMP shock in the financial conditions (FC) model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

Figure 1.C.19: Responses of EME variables (mean) to US UMP shock - model: FC with long term government bond yields

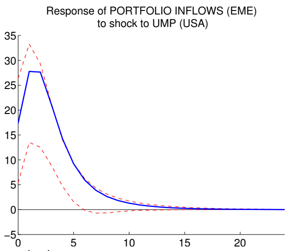


(a) portfolio inflows (b) long-term government bond yields

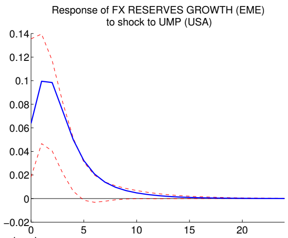
(c) real fx change (d) equity returns

Note: The figure shows the estimated mean impulse responses, along with 68 percent confidence bands, of the EME variables to the UMP shock in the financial conditions (FC) model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

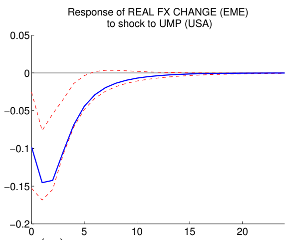
Figure 1.C.20: Responses of EME variables (mean) to US UMP shock - model: FC, fx reserves as monetary policy instrument in EMEs



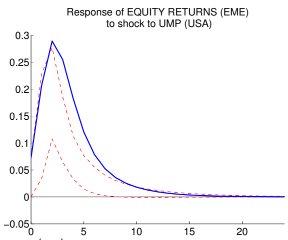
(a) portfolio inflows



(b) fx reserves growth



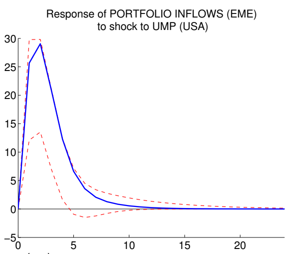
(c) real fx change



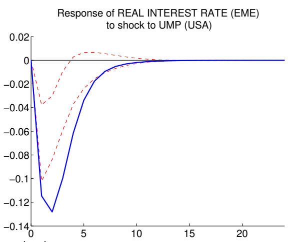
(d) equity returns

Note: The figure shows the estimated mean impulse responses, along with 68 percent confidence bands, of the EME variables to the UMP shock in the financial conditions (FC) model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

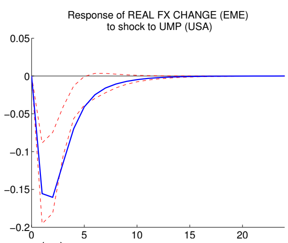
Figure 1.C.21: Responses of EME variables (mean) to US UMP shock - model: FC, US portfolio outflows ordered before monetary policy variables



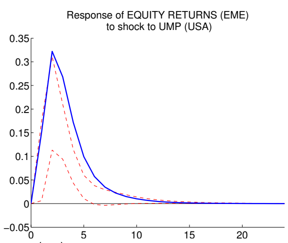
(a) portfolio inflows



(b) real interest rate



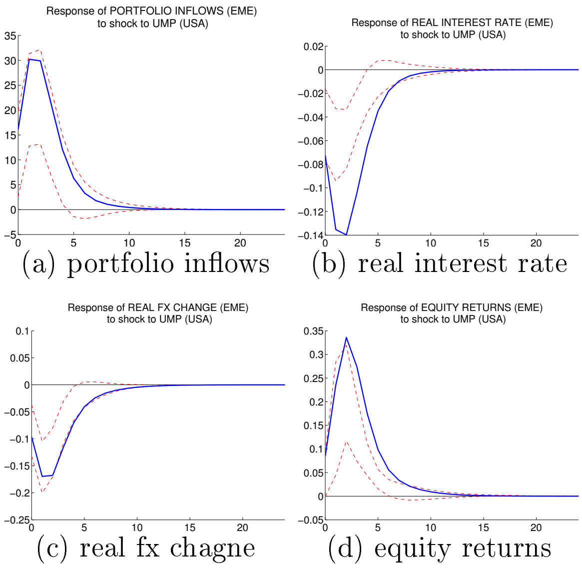
(c) real fx change



(d) equity returns

Note: The figure shows the estimated mean impulse responses, along with 68 percent confidence bands, of the EME variables to the UMP shock in the financial conditions (FC) model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

Figure 1.C.22: Responses of EME variables (mean) to US UMP shock - model: FC, Argentina is dropped from the estimation

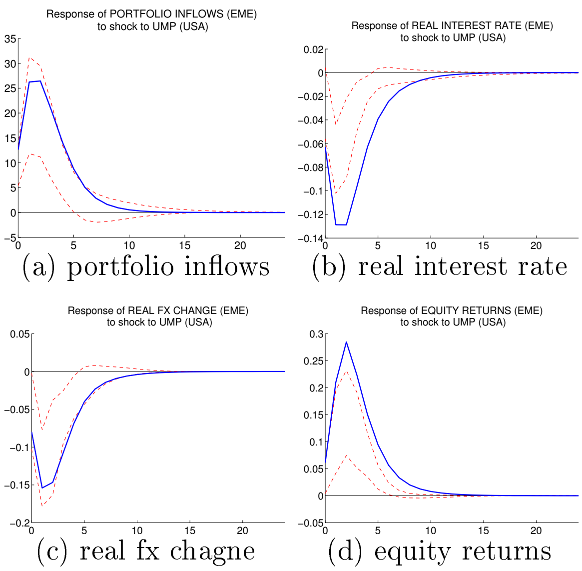


(a) portfolio inflows (b) real interest rate

(c) real fx chagne (d) equity returns

Note: The figure shows the estimated mean impulse responses, along with 68 percent confidence bands, of the EME variables to the UMP shock in the financial conditions (FC) model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

Figure 1.C.23: Responses of EME variables (mean) to US UMP shock - model: FC, alternative lag length

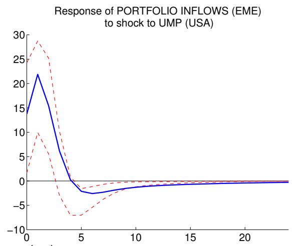


(a) portfolio inflows (b) real interest rate

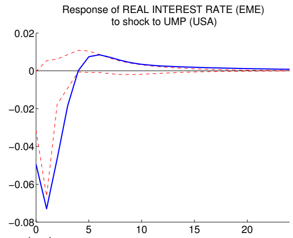
(c) real fx chagne (d) equity returns

Note: The figure shows the estimated mean impulse responses, along with 68 percent confidence bands, of the EME variables to the UMP shock in the financial conditions (FC) model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

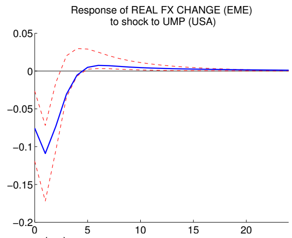
Figure 1.C.24: Responses of EME variables (mean) to US UMP shock - model: FC, post-Lehman sample



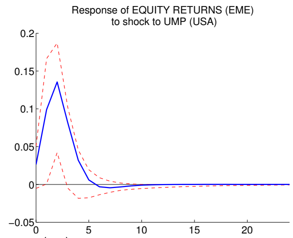
(a) portfolio inflows



(b) real interest rate



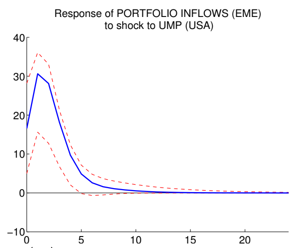
(c) real fx change



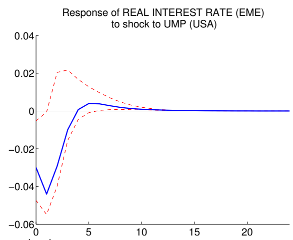
(d) equity returns

Note: The figure shows the estimated mean impulse responses, along with 68 percent confidence bands, of the EME variables to the UMP shock in the financial conditions (FC) model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

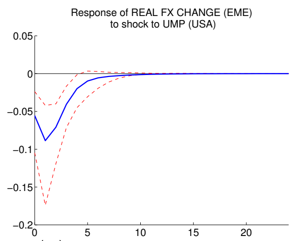
Figure 1.C.25: Responses of EME variables (mean) to US UMP shock - model: FC, commodity price inflation & portfolio flows as transmission channels



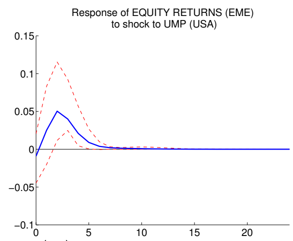
(a) portfolio inflows



(b) real interest rate



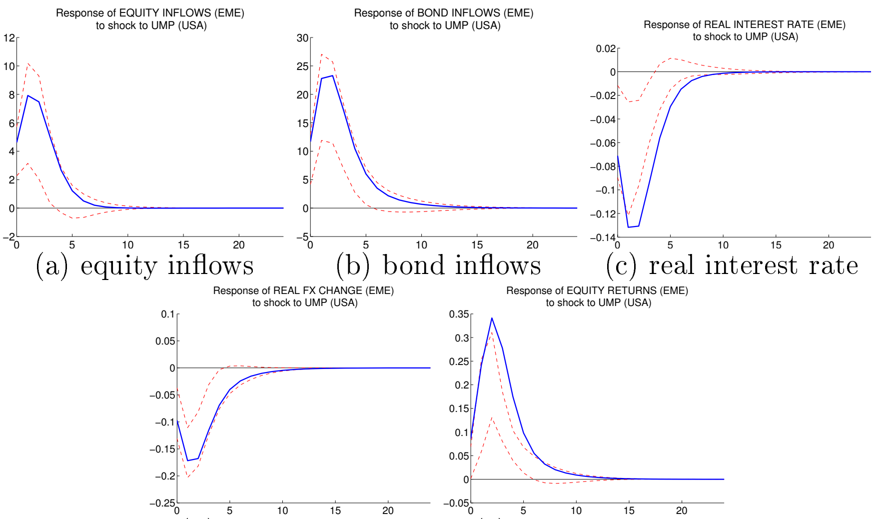
(c) real ifx change



(d) equity returns

Note: The figure shows the estimated mean impulse responses, along with 68 percent confidence bands, of the EME variables to the UMP shock in the financial conditions (FC) model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

Figure 1.C.26: Responses of EME variables (mean) to US UMP shock - model: FC, equity and bond flows

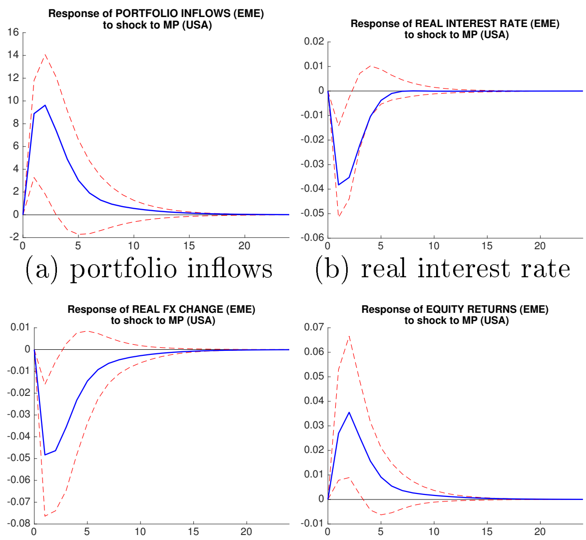


(a) equity inflows (b) bond inflows (c) real interest rate

(d) real fx change (e) equity returns

Note: The figure shows the estimated mean impulse responses, along with 68 percent confidence bands, of the EME variables to the UMP shock in the financial conditions (FC) model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

Figure 1.C.27: Responses of EME variables (mean) to US UMP shock - model: FC, shadow federal funds rate as monetary policy instrument in the US

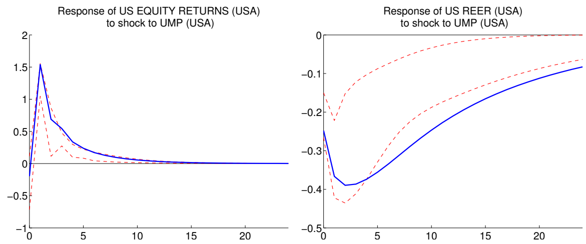


(a) portfolio inflows (b) real interest rate

(c) real fx change (d) equity returns

Note: The figure shows the estimated mean impulse responses, along with 68 percent confidence bands, of the EME variables to the UMP shock in the financial conditions (FC) model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

Figure 1.C.28: Responses of US equity returns and US REER to US UMP shock



(a) US equity returns (b) US reer

Note: The figure shows the estimated mean impulse responses, along with 68 percent confidence bands, of the EME variables to the UMP shock in the financial conditions (FC) model. Confidence bands are based on 500 bootstrap replications with 1000 draws of the rotation matrix each.

1.D Additional tables

Table 1.D.1: Panel unit root tests

	t-stat	critical values (1% / 5% / 10%)	Type	Det
IP	-3.84	-2.73 / -2.63 / -2.56	2	3
D(IP)	-11.35	-2.24 / -2.13 / -2.06	2	2
Real GDP	-1.77	-2.73 / -2.63 / -2.56	2	3
D(Real GDP)	-5.59	-2.25 / -2.13 / -2.06	2	2
CPI	-1.98	-2.73 / -2.63 / -2.56	2	3
D(CPI)	-8.21	-2.25 / -2.13 / -2.06	2	2
Portfolio Inflows	-8.01	-2.25 / -2.13 / -2.06	2	2
D(Portfolio Inflows)	-8.11	-1.70 / -1.56 / -1.47	2	1
US Portfolio Outflows	-6.14	-3.52 / -2.90 / -2.58	0	2
D(US Portfolio Outflows)	-11.25	-2.58 / -1.95 / -1.62	0	1
VIX	-2.62	-3.52 / -2.90 / -2.58	0	2
D(VIX)	-8.39	-2.60 / -1.95 / -1.62	0	1
Fed Balance Sheet	-3.74	-4.09 / -3.47 / -3.17	0	3
D(Fed Balance Sheet)	-1.79	-3.54 / -2.90 / -2.59	0	2
Real Interest Rate	-8.09	-2.25 / -2.13 / -2.05	2	2
D(Real Interest Rate)	-14.90	-1.70 / -1.56 / -1.47	2	1
Equity Price Index	-2.49	-2.73 / -2.63 / -2.56	2	3
D(Equity Price Index)	-10.01	-2.25 / -2.13 / -2.06	2	2
Real FX	-2.10	-2.25 / -2.13 / -2.06	2	2
D(Real FX)	-12.05	-1.70 / -1.56 / -1.47	2	1

Type: Type of panel unit root test,
2 = P-CEA test controlling for cross-section dependence of errors,
1 = IPS test under the assumption of no cross-section dependence,
0 = simple ADF test for cross-section invariant data.
Det: Deterministics case for the test regression,
3 = with const and trend, 2 = with constant, 1 = without deterministics.

CHAPTER 2

Identifying Speculative Demand Shocks in Commodity Futures Markets through Changes in Volatility[1]

2.1 Introduction

The recent drastic boom-bust cycles in commodity prices spurred an intense debate about the increased presence of financial investors in commodity markets. The discussion revolves around whether investors are responsible for the large price swings and, more generally, whether they drive prices away from fundamentals, distort price signals, and reduce welfare. Growing concerns among policy makers already led to initiatives of stronger futures market regulation.[2] The empirical literature, on the other hand, has reached no consensus on whether and how financial investment affects commodity prices. Many studies use publicly available data on futures market positions aggregated by groups of traders, provided by the US Commodity Futures Trading Commission (CFTC) at a weekly or lower frequency. Only a few

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[2]In the US, the Dodd-Frank act granted the Commodity Futures Trading Commission (CFTC) the responsibility for additional regulations of commodity derivative markets. In the European Union, the European Commission set up an expert group on the regulation of commodity derivatives.

of these papers document positive effects of investor flows on futures returns for specific sample periods and markets (see Singleton, 2014, and Gilbert and Pfuderer, 2014). Most of them, however, find no effect of speculators' position changes on futures prices (see, among others, Stoll and Whaley, 2010, Büyükşahin and Harris, 2011, Irwin and Sanders, 2012, Aulerich et al., 2013, Hamilton and Wu, 2015).[3]

The main challenge in this literature is identification. Specifically, it is necessary to isolate variation in investors' positions due to trades actually initiated by speculators from variation due to trades initiated by other market participants, such as producers, to which speculators only respond by taking the counter-side. This distinction is important because only the former trades induce a positive correlation between speculators' long positions and futures prices, whereas the latter trades imply a negative correlation as producers need to compensate speculators for taking the risk by setting the futures price at a discount. A lack of identification might thus imply an insignificant correlation, as both types of trades are averaged.

Two recent studies address the identification issue using daily proprietary or disagreggated data and find significant positive price effects of financial investments. Henderson et al. (2015) use detailed issuance data on commodity-linked notes and show that futures prices increase when the financial institutions issuing the notes hedge their short exposure vis-à-vis the holders of the securities through long positions in the futures market. Cheng et al. (2015) have access to the CFTC's Large Trader Reporting System which provides private account-level data on individual traders' positions. The authors show that increases in the VIX, that are associated with lower futures prices, lead to a reduction in financial traders' exposure, and to an increase in producers' net long positions. This is consistent with financial traders initiating the trades.

In this paper, we provide new evidence on the price effects of financial investments in commodity futures markets by proposing an approach to address the identification issue in the publicly available aggregated weekly CFTC data. Specifically, we identify a system of simultaneous equations, modeled as a vector autoregression (VAR), through the heteroskedasticity that is present in the weekly data to isolate exogenous variation in speculators' net long positions. Following Sentana and Fiorentini (2001) and Rigobon (2003), the approach exploits the fact that changes in the volatility of

[3]See also Fattouh et al. (2013) and Cheng and Xiong (2013) for overviews of the literature.

the structural shocks in the system contain additional information on the relation between the endogenous variables. For example, in a period of high speculative demand volatility, we learn more about the response of returns to positions as the covariance between both variables temporarily increases. Then, speculative demand shocks are more likely to occur and can be used as a 'probabilistic instrument' (see Rigobon, 2003).

The model includes three endogenous variables: commodity futures returns and net long positions of 'index investors' and 'hedge funds', respectively, who are both financial speculators.[4] We use position data from the CFTC Supplemental Commitments of Traders (SCOT) reports, which contain a proper category for 'index investors'. Both groups are important in terms of market share and have received considerable attention in the academic debate (see Büyüksahin and Robe, 2014, Singleton, 2014, Cheng et al., 2015, Basak and Pavlova, 2016). The reports cover eleven agricultural markets, but exclude energy and metal markets. For the core analysis, we compute an aggregate index over all markets for each endogenous variable and apply a statistical approach to the reduced-form residuals of the model to detect changes in the volatility of the structural shocks. These changes in volatility, together with the assumption of time-invariant impact effects, are central to achieving identification. Formal tests support the necessary assumptions and indicate that identification has been achieved from a statistical point of view.

Our results suggest that the identified exogenous position changes of speculators have significant contemporaneous price effects and that they are a relevant driver of futures returns. In particular, we find that demand shocks of both index investors and hedge funds impact positively on returns. A one standard deviation shock to index investors' net long positions increases futures returns significantly by 0.15 standard deviations on impact. The contemporaneous effect of hedge funds' demand shocks on returns is 0.39 standard deviations. These results are qualitatively and quantitatively robust to various alterations of the model and the data. Specifically, we assess the sensitivity of the estimates to changing the definition of volatility regimes, to adding another trader group to the model, and to splitting the sample.

[4] The latter group actually contains positions of hedge funds, commodity pool operators, and commodity trading advisors. For brevity, we refer to this category as hedge funds in the following.

Our results also hold on the single markets underlying the aggregate indexes used in the main specification.

We further assess the economic importance of the identified speculative demand shocks for commodity price fluctuations with variance and historical decompositions. The variance decompositions suggest that the shocks account for roughly one fifth of the variation in returns on average. Moreover, their importance increases during periods of high speculative demand volatility. Then, demand shocks of hedge funds account for 30 percent of the variation in futures returns, and demand shocks of index investors explain up to 10 percent. By means of historical decompositions, we also quantify the relevance of fundamental demand and supply conditions as well as changes in the VIX and oil prices for explaining agricultural futures prices. The results suggest that these forces account for the largest part of commodity price fluctuations, and in particular explain their secular dynamics. Speculative demand shocks, on the other hand, seem to mainly contribute to short-run price movements.

Overall, the results support existing studies that detect significant impacts of financial investments on commodity returns based on highly disaggregated or proprietary data. Using a structural VAR approach allows quantifying the statistical significance of speculative demand shocks and their economic importance - both on average and during specific time periods - with publicly available aggregated data. The documented price effects are consistent with two recent strands of theoretical models. The first strand emphasizes the existence of limits to arbitrage. When financial intermediaries are funding constrained, position changes of other market participant can have price effects (see He and Krishnamurthy, 2013, Acharya et al., 2013, Hamilton and Wu, 2014). The second strand stresses the role of informational frictions. Under asymmetric information, trades can transmit private signals to the market and thereby affect prices (see Goldstein and Yang, 2015, 2016, Sockin and Xiong, 2015).

Methodologically, our paper connects to a fast-growing line of research that investigates the role of financial investors in commodity markets using time-series models. Irwin and Sanders (2012), Aulerich et al. (2013), or Gilbert and Pfuderer (2014) rely on bivariate Granger-causality tests or similar techniques. Other authors use structural VAR models. Ederer et al. (2013) and Bruno et al. (2017) employ Cholesky identification schemes. Zero restrictions, however, seem difficult to defend when working with weekly or lower frequency financial market data. Alternatively, Kilian and Murphy (2014) use sign restrictions that allow for an instantaneous response of all endogenous variables. Different to our focus, they analyze the impact of speculation tied to physical inventories. Moreover, sign restrictions do not allow us to disentangle the main shocks of interest, as theory gives similar predictions regarding the sign of the impact of several of the model's structural shocks on the endogenous variables. Therefore, we apply an agnostic identification approach using changes in volatility, without additional sign or zero restrictions.

The remainder of the paper is structured as follows. Section 2.2 outlines a simple theoretical framework to develop a notion about the structural shocks driving the systems of equations and to derive testable hypotheses. Then, we describe the data and the identification strategy in Section 2.3. Section 2.4 contains the main results, while their sensitivity and robustness is evaluated in Section 2.5. The last section concludes.

2.2 Theoretical framework

Our model of simultaneous equations contains three endogenous variables: the commodity futures return and net long positions of index investors and hedge funds, respectively. The variables are assumed to be contemporaneously driven by three uncorrelated structural shocks as well as exogenous variables. To develop a notion about the three structural shocks, we employ a simple theoretical model of the futures market which also allows us to derive some hypotheses about the contemporaneous impacts of the shocks on the endogenous variables. As in Cheng et al. (2015), we consider a one period model with different groups of market participants, hedgers and–in our case–two groups of financial investors. The hedgers are commodity producers (pr) who need to hedge their price risk in the futures market. Financial investors are speculators without an interest in the physical delivery of the commodity, consisting of index investors and hedge funds ($f1$ and $f2$).

Speculative demand of the two groups of financial investors is driven by two idiosyncratic shocks, υ_{f1} and υ_{f2}, that motivate them to change their positions. Additionally, there is a shock η which commonly affects demand of all three groups. As we will exogenously control for physical supply conditions and financial market risk in the model (see Section 2.3.3), the common shock is unrelated to these two driving forces. Instead, it can be thought of as capturing changes in demand in the spot market which are transmitted to the futures market. The demand curves for producers and financial investors are

$$\begin{aligned}\Delta x_{pr} &= -\beta_{pr}\Delta F - \gamma_{pr}\eta,\\ \Delta x_{f1} &= -\beta_{f1}\Delta F + \gamma_{f1}\eta + \upsilon_{f1},\\ \Delta x_{f2} &= -\beta_{f2}\Delta F + \gamma_{f2}\eta + \upsilon_{f2},\end{aligned}$$

where ΔF is the change in the futures price, Δx_{pr}, Δx_{f1}, and Δx_{f2} is the change in net long demand of hedgers and financial investors, respectively, and it is assumed that $\beta_{pr}, \beta_{f1}, \beta_{f2} \geq 0$ and that $\gamma_{pr}, \gamma_{f1}, \gamma_{f2} \geq 0$. The first assumption implies that all demand curves are downward sloping. The second assumption relates to the common shock. To meet higher physical demand, commodity producers increase their output, which in turn raises their hedging needs. The common shock therefore causes a decline in net long demand of producers in the futures market. We further assume that the common shock increases speculative net long demand of financial investors as the physical demand for commodities rises. This reaction can be motivated by, for example, trend-following behavior as speculators expect further price increases (see Rouwenhorst, 1998, Bhardwaj et al., 2014, Kang et al., 2014).

Market clearing imposes that $\Delta x_{pr} + \Delta x_{f1} + \Delta x_{f2} = 0$ where the equilibrium price balances the three groups' net demand. Solving the model with respect to the underlying shocks yields the following equation for the change in the futures price:

$$\Delta F = \frac{1}{\beta_{pr} + \beta_{f1} + \beta_{f2}}\upsilon_{f1} + \frac{1}{\beta_{pr} + \beta_{f1} + \beta_{f2}}\upsilon_{f2} + \frac{\gamma_{f1} + \gamma_{f2} - \gamma_{pr}}{\beta_{pr} + \beta_{f1} + \beta_{f2}}\eta$$

According to the price equation $\partial\Delta F/\partial\upsilon_{f1} > 0$ and $\partial\Delta F/\partial\upsilon_{f2} > 0$ if $\beta_{pr}+\beta_{f1}+\beta_{f2} < \infty$. For the empirical model this implies the testable hypothesis:

Hypothesis 1. *Positive speculative demand shocks lead to an increase in net long positions of financial investors and contemporaneously increase commodity futures returns.*

The alternative is to find no significant effect of speculative demand shocks on futures returns. This result would indicate that some or all of the β_i are so large that $1/(\beta_{pr} + \beta_{f1} + \beta_{f2})$ is statistically indistinguishable from zero. Economically, this means that idiosyncratic position changes by financial investors are absorbed by other market participants with nearly infinitely elastic demand curves and have no price effects.

While Hypothesis 1 is derived from a highly stylized model, it is consistent with more sophisticated asset pricing models. Shleifer and Summers (1990) and Shleifer and Vishny (1997), for example, show that large position changes can influence prices through an effect on the order book if the instantaneous supply of counter-party orders is low. Such problems of illiquidity might arise if there are limits to arbitrage which deter risk averse arbitrageurs from taking the counter-side. Positions changes can also influence futures market risk premia and thereby drive up prices (see Acharya et al., 2013, and Hamilton and Wu, 2014, 2015). If producers want to hedge their price risk, the futures price needs to include a risk premium and, hence, to be set at a discount to induce speculative traders to take the price risk. The higher is the provision of hedging liquidity through speculators, the lower is the risk premium and, hence, the higher the futures price.[5] Additionally, financial investors could affect prices through informational channels. If some investors possess private information, their trades might communicate this information to the market and change the price (see Grossman and Stiglitz, 1980, Hellwig, 1980, Goldstein and Yang, 2015). Private information could be due to better forecasting abilities, different costs of private information production, or divergent interpretations of public information (see Singleton, 2014).

The effect of the common shock on the futures price depends on the relative size of γ_{pr}, γ_{f1}, and γ_{f2}. If long demand of investors increases by more than short demand of producers in response to the shock, that is, if $\gamma_{f1} + \gamma_{f2} > \gamma_{pr}$, then $\partial \Delta F / \partial \eta > 0$. Solving the model yields the following equation for changes in net long positions of

[5]Following the theory of normal backwardation, going back to Keynes (1930), the spot and the futures price are related according to $F_{t,T} - S_t = [\mathbb{E}(S_T) - S_t] - \pi_{t,T}$, where S_t and S_T are the spot price at t and T, respectively, $F_{t,T}$ the T-periods ahead futures price and $\pi_{t,T}$ the risk premium. If short hedging demand exceeds long supply, the risk premium will be positive. Hamilton and Wu (2014, 2015) show that the same mechanism is at work if the market is characterized by long-pressure of speculators and not by short-pressure of producers. If speculators cannot find a counter-party to take the short side, the futures contract needs to include a risk premium on the short side. Therefore, an increase in speculators long exposure can lead to an increase in futures prices if they affect risk premia.

financial investor group $i = 1, 2$

$$\Delta x_{fi} = \frac{\beta_{pr} + \beta_{fj}}{\beta_{pr} + \beta_{f1} + \beta_{f2}} \upsilon_{fi} - \frac{\beta_{fi}}{\beta_{pr} + \beta_{f1} + \beta_{f2}} \upsilon_{fj} + \frac{(\beta_{pr} + \beta_{fj})\gamma_{fi} + \beta_{fi}(\gamma_{pr} - \gamma_{fj})}{\beta_{pr} + \beta_{f1} + \beta_{f2}} \eta,$$

where j denotes the other investor group. The sign of the effect of physical demand shocks on financial investors' demand, $\partial \Delta x_{fi}/\partial \eta$, depends on the relative sizes of the parameters. However, as long as no group reacts extremely to the common shock (γ_{fi} very large) and no group reacts extremely to price changes (β_{fi} very large), it follows that $\partial \Delta x_{f1}/\partial \eta > 0$ and $\partial \Delta x_{f2}/\partial \eta > 0$. For the empirical model these observations can be translated to

Hypothesis 2. *Positive physical demand shocks have a positive contemporaneous effect on commodity futures returns and drive up net long positions of financial speculators.*

The alternative is that physical demand shocks have no significant effect on or even lead to a decrease in speculators' net long positions. This could be the case if, for example, $\gamma_{fj} \gg \gamma_{pr}$ and $\beta_{fi} \gg 0$. In the next section, we outline how we specify the empirical model to test the two hypotheses.

2.3 Empirical model, data, and estimation methodology

2.3.1 Empirical model

The structural VAR model is given by

$$Ay_t = \tilde{c} + \tilde{A}_1 y_{t-1} + ... + \tilde{A}_p y_{t-p} + \tilde{\Lambda} x_t + \epsilon_t, \tag{2.1}$$

with the vector of endogenous variables

$$y_t = \begin{pmatrix} \Delta log(\text{Agricultural futures price})_t \\ \Delta(\text{Net long positions index investors})_t \\ \Delta(\text{Net long positions hedge funds})_t \end{pmatrix},$$

x_t a vector of exogenous variables, and $\tilde{c}$, $\tilde{A}_p$, and $\tilde{\Lambda}$ parameter matrices. The vector ϵ_t contains the structural shocks with regime-dependent diagonal covariance matrix in regime k

$$\Sigma_{\epsilon,k} = E(\epsilon_t \epsilon_t') = \begin{pmatrix} \sigma_k^F & 0 & 0 \\ 0 & \sigma_k^I & 0 \\ 0 & 0 & \sigma_k^H \end{pmatrix}.$$

In its reduced-form, the model in equation (2.1) can be re-written as follows

$$y_t = c + \Pi_1 y_{t-1} + ... + \Pi_p y_{t-p} + \Lambda x_t + u_t, \tag{2.2}$$

where $\Pi_p = A^{-1}\tilde{A}_p$, $\Lambda = A^{-1}\tilde{\Lambda}$, and $u_t = (u_t^F, u_t^I, u_t^H)'$. The vector of reduced-form residuals, u_t, is related to the structural shocks through matrix A^{-1}: $u_t = A^{-1}\epsilon_t$.

The focus of the empirical analysis is on the impact matrix A^{-1} that contains the contemporaneous effects of the structural shocks on the endogenous variables. Specifically, the hypotheses outlined in Section 2.2 can be assessed based on the estimated A^{-1}. Assuming that the identified structural shocks in the two equations with investors positions are speculative demand shocks of the different investor groups, Hypothesis 1 comes down to testing $\alpha_{1,2}, \alpha_{1,3} > 0$, where $\alpha_{j,k}$ is the corresponding element in A^{-1}. Similarly, if the structural shock in the futures price equation of the estimated structural VAR model is the physical demand shock, Hypothesis 2 can be tested by analyzing whether $\alpha_{2,1}, \alpha_{3,1} > 0$. Therefore, after the estimation, we first assess how the estimated structural shocks can be interpreted with the outlined theoretical model in mind, before evaluating the estimated parameters in A^{-1}. For our baseline model, we will also assess the *direct* effects of structural shocks captured in A. They differ from the *overall* effects in A^{-1} as they do not take instantaneous feedback among endogenous variables into account. Instead, parameters in A can be interpreted as effects of shocks keeping all other variables constant and showing both is thus indicative of shock amplification among endogenous variables.

2.3.2 Identification

Equation (2.2) and the regime-dependent covariance matrix of the reduced-form shocks, $\Sigma_{u,k}$, can be estimated consistently by ordinary least squares. Specifically, we specify the model in first (log) differences to account for the non-stationarity of the data.[6] Moreover, we standardize all variables prior to the estimation.[7] We include two lags of the endogenous variables to obtain residuals free from autocorrelation. From (2.1) and (2.2), it follows that $\Sigma_{u,k} = A^{-1}\Sigma_{\epsilon,k}(A^{-1})'$. This relation illustrates how different volatility regimes contain additional information that can be exploited to identify the impact matrix A (or equivalently A^{-1}). With $k = 1$ we would only have six moments on the LHS that can be estimated but nine parameters that need to be determined on the RHS (three structural shock variances and six off-diagonal elements in A, with the main diagonal normalized to unity). For $k \geq 2$, however, the system has at least as many moments that can be estimated (for instance, twelve if $k = 2$) as unkowns (six shock variances and six off-diagonal elements if $k = 2$).

The approach of identification through heteroskedasticity has been developed by Sentana and Fiorentini (2001) and Rigobon (2003) and applied in the context of financial markets and asset price co-movements by, among others, Bouakez and Normandin (2010) and Ehrmann et al. (2011). The idea is that changes in the relative variances of the structural shocks over time, that is, changes in $\sigma_k^S/\sigma_k^{S'}$ across k with $S = F, I, H$, contain additional information which allows determining the entries in A. If, for example, the variance of index investor position changes increases in a certain period ($\sigma_k^I/\sigma_k^F > \sigma_{k'}^I/\sigma_{k'}^F$), speculative demand shocks coming from that group help tracing out the demand curve of other market participants, and thereby the price effect, because large speculative demand shocks of index investors are more likely to occur during this period. Rigobon (2003) refers to these relative changes in volatility as 'probabilistic instruments'. The identification strategy relies on two assumptions. First, the structural shocks are uncorrelated. This is commonly assumed in the structural VAR literature. Second, the matrix of contemporaneous impacts A is constant across volatility regimes. This is a standard assumption

[6] Augmented Dickey-Fuller and Phillips-Perron tests on the level of the agricultural futures price and investors' net long positions do not reject the null hypothesis of a unit root, irrespective of whether we include a drift term. Returns and first differences of positions, in contrast, are found to be stationary.

[7] Specifically, we subtract the mean and divide by the standard deviation. This facilitates a direct comparison of the effects across variables and markets. Moreover, it reduces the computational challenges of the minimization procedure as the parameters to be estimated are of similar order of magnitude. For the main specification, we have verified that the results are robust to using non-standardized data.

for instance in (G)ARCH models. Moreover, we formally test the assumption and cannot reject it.

Alternatively, identification is often achieved by imposing zero or sign restrictions. Zero restrictions would imply a delayed response of some endogenous variables to some structural shocks. This seems too restrictive, however, as futures prices and positions are likely to respond to shocks and each other contemporaneously at the weekly frequency. Sign restrictions, on the other hand, allow for an immediate impact among variables. Yet, they are not helpful in disentangling the shocks in our model as these shocks all imply the same sign pattern (compare Section 2.2), and it would thus take further strong assumptions, for instance on the relative magnitude of their impact, to disentangle them.

2.3.3 Data

To measure positions of the trader groups, we use publicly available data from the CFTC Supplemental Commitments of Traders (SCOT). In the reports, traders are classified into four categories: ‘commercial’ (producers, processors, and merchants), ‘non-commercial’ (commodity trading advisors (CTAs), commodity pool operators (CPOs), hedge funds, and other reportables), to which we for brevity mostly just refer as hedge funds, ‘non-reporting traders’, and ‘index investors’. Both index investors and non-commercial traders are financial investors without an interest in the physical delivery of the commodity. There are, however, some differences in their characteristic trading strategies (see Masters and White, 2008, Mou, 2011, Heumesser and Staritz, 2013). Traders in the non-commercial category actively gather and process commodity-specific information and base their trades thereon. CTAs and CPOs have an insightful knowledge of specific agricultural markets and hedge funds often take directional views by exploiting high-frequency cross-market information. These investors are typically active on both sides of the market. In contrast, index investors essentially use commodities to diversify portfolio risk, but have no particular interest in specific commodities. Their trades are based on rebalancing, rolling, or weighting considerations and occur at lower frequencies. They are typically only active on the long side of the market.

The SCOT reports cover all eleven agricultural commodities in the S&P Goldman Sachs Commodity Index (GSCI), one of the most widely used investible commodity indices, but exclude energy and metal futures markets. Observations start in July 2006, so that our sample runs from 04 July 2006 to 29 March 2016. The data frequency in the reports is weekly. To measure futures prices, we use corresponding nearby futures contracts available from Thompson Reuters Datastream. For the core analysis, we construct one aggregate index for each endogenous variable. The weights of the individual commodities in each index are based on the commodities' yearly varying weights in the S&P GSCI. In a robustness analysis, we also investigate the relations among the endogenous variables on the individual markets using market-specific price and position data, that is, we estimate one three-variable model for each of the eleven agricultural futures markets separately. Table 2.3.1 lists the commodities and their average weights in our sample.

Table 2.3.1: Average commodity weights used for construction of aggregate futures market indexes

Corn	*SRW Wheat*	*Live Cattle*	*Soybeans*	*Sugar*	*Lean Hogs*
20.9%	18.1%	12.9%	12.5%	9.1%	7.4%
Cotton	*HRW Wheat*	*Coffee*	*Feeder Cattle*	*Cocoa*	
6.1%	4.7%	4.1%	2.3%	1.4%	

The table lists the commodities used for construction of aggregate futures price, position, and spread indexes and their average weights in these indexes. The weights are updated yearly and based on the reported weights in the S&P GSCI. Differences to 100% are due to rounding errors.

While the SCOT reports have the advantage of being publicly available and distinguishing between index investors and other speculators, they also have notable drawbacks. The data might contain reporting errors due to potential missclassication of traders for several reasons. First, financial investors have incentives to try being classified as hedgers, since this entails them for preferential treatments like exemption from positions limits or posting lower margins to clearinghouses. These incentives might have even increased after the the Dodd-Frank Act in 2010, when additional regulatory measures started to get implemented. Second, in particular large financial entities might trade for different reasons, like setting up a trade for a customer, proprietary trading, or index trading. The reports, however, are based on aggregated total end-of-day positions of individual traders and not on the underlying motifs behind their specific trades. Third, the CTFC itself changes the classification of traders from time to time, for instance, if additional information on a trader is available or when its client base changes.

Overall, these potential reporting errors in the data could show up in the estimated VAR model. With speculators being partly classified as hedgers, our results might actually represent a lower bound for the impact of speculators' position changes on futures price formation and any detected significant impact should still be supportive for the hypotheses. We also explicitly control for the impact of the Dodd-Frank act on our results by splitting the sample at this point in the sensitivity analysis, and we ensure that the changes in volatility used to identify the model are not solely driven by specific re-classification of traders by assessing the robustness of the results to various definitions of volatility regimes.

We add several exogenous variables to the model. First, we control for physical supply in the US as most of the included commodities are to a large extent produced there. Specifically, we build an index of crop conditions following Bruno et al. (2017) for this purpose. Second, changes in uncertainty and risk aversion can have an impact on commodity futures prices and financial investors' risk bearing capacity. Cheng et al. (2015) show that speculators adjust positions to changes in the CBOE Volatility Index (VIX). We thus control for changes in the VIX. Third, changes in the price of oil can affect the price of agricultural products (see Baffes, 2007). One argument is that oil prices are part of production costs. Wang et al. (2014) find effects of oil shocks on agricultural commodity prices. As oil prices are highly correlated with the VIX including both of them jointly into the model would lead to problems of multicollinearity and therefore would make it difficult to interpret significance levels. We therefore use changes in the oil price orthogonal to changes in the VIX, computed as the residuals from a regression of oil returns on VIX changes. Fourth, we add the size of the Federal Reserve balance sheet as a measure of aggregate liquidity. Finally, the model contains monthly dummy variables to capture seasonal effects, following Kilian and Murphy (2014). All exogenous variables enter the model contemporaneously. As explained in Section 2.2, the estimated structural shocks in equation (2.1) thus explain the variation in the data that is left after controlling for the exogenous variables, and have to be interpreted accordingly. We provide a detailed description of the data in Appendix 2.A.

2.3.4 Estimation

Before the estimation, we need to determine the volatility regimes used to identify the model. Following Ehrmann et al. (2011), we apply a statistical approach. Specifically, we compute the rolling standard deviation for each reduced-form residual in u_t. We then calibrate a threshold for the rolling standard deviations above which the corresponding residual is classified into a high volatility regime. In particular, we use a window of 15 weeks to compute the rolling standard deviations and a threshold of one standard deviation. We define regime 1 as a low volatility regime, where the standard deviation of all three residuals is below one. Regimes 2 to 4 are characterized by high volatility of only one of the residuals, while the other two residuals display low volatility.

The approach of defining one high volatility regime for each residual is motivated by the identification idea that a relative volatility shift of the underlying structural shock helps to trace out the effects of that shock on the other variables. The choice of the window and the threshold is then largely dictated by the need to have sufficient observations in each regime and the objective of minimizing the number of observations which do not fit into any regime, for example, because two reduced-form errors are in the high volatility regime simultaneously. We drop these observations from the estimation of the regime-specific reduced-form covariance matrices. Finally, note that the approach generates four volatility regimes, while two regimes are in principle enough for identification. Hence, the model is overidentified and the overidentifying restrictions implied by a regime-invariant A can be tested.

Table 2.3.2 shows the estimated variances of the residuals and the number of observations per regime. It also contains the regime-specific estimated covariances between the residuals. The latter illustrate the idea underlying identification through heteroskedasticity. In regime 3, for example, where the reduced-form errors of the index investor equation display high volatility, the covariance between these residuals and those of the futures returns equation increases strongly relative to the regime 1, where both residuals show low volatility. Similarly, the covariance between the residuals of the hedge funds equation and of the futures return equation increases substantially in regime 4. These changes in the covariances provide the additional information needed for identification.

Table 2.3.2: Variance/covariance of the reduced-form shocks in the different regimes

Regime	(1) All low volatility	(2) Return high volatility	(3) Index inv. high volatility	(4) Hedge funds high volatility
$V(u_t^F)$	0.47	1.17	0.38	0.52
$V(u_t^I)$	0.42	0.57	1.68	0.48
$V(u_t^H)$	0.49	0.49	0.58	1.25
$C(u_t^F, u_t^I)$	0.13	0.11	0.26	0.14
$C(u_t^F, u_t^H)$	0.27	0.47	0.25	0.54
Observations	152	59	51	96

The table shows the estimated variances and covariances of the reduced-form errors in the different volatility regimes. The sample period is 04 July 2006 - 29 March 2016.

To see whether our regime definition is supported by the data, we test formally for the constancy of the reduced-form covariance matrix. Recall that for identification we not only require changes in the volatility of the reduced-form residuals, which we expect given our construction of regimes, but in particular significant changes in the covariances between residuals across regimes. Following Lanne and Lütkepohl (2008), we thus perform pairwise likelihood ratio tests on the null hypothesis that two regimes have the same covariance matrix. Moreover, we test the joint null hypothesis that all four covariance matrices are the same. Table 2.3.3 shows that all null hypotheses are strongly rejected by the data. It is known that such likelihood ratio tests do not have optimal small sample properties. The null might be rejected too often. However, our test statistics are large, so that we reject the equality of the matrices with confidence, and in particular the joint equality of all matrices. The data prefer a model with changes in volatility over the assumption of homoskedasticity.

With the volatility regimes in hand, we estimate the model as in Ehrmann et al. (2011) by minimizing the following matrix norm:

$$||g'g|| = \sqrt{tr[gg']} = \sqrt{vec(g)vec(g)'}, \quad with \quad g = \sum_{k=1}^{4}[A\Sigma_{u,k}A' - \Sigma_{\epsilon,k}] \qquad (2.3)$$

Table 2.3.3: Tests for constancy of reduced-form covariance matrix

H_0	$\Sigma_{u,1} = \Sigma_{u,2} = \Sigma_{u,3} = \Sigma_{u,4}$		$\Sigma_{u,1} = \Sigma_{u,2}$	$\Sigma_{u,1} = \Sigma_{u,3}$
LR statistic	746.82		31.86	46.46
p-value	(0.00)		(0.00)	(0.00)
H_0	$\Sigma_{u,1} = \Sigma_{u,4}$	$\Sigma_{u,2} = \Sigma_{u,3}$	$\Sigma_{u,2} = \Sigma_{u,4}$	$\Sigma_{u,3} = \Sigma_{u,4}$
LR statistic	30.68	43.82	48.22	38.93
p-value	(0.00)	(0.00)	(0.00)	(0.00)

The table shows results of likelihood ratio tests on the null hypothesis that all regimes have the same reduced-form covariance matrix and that pairwise regimes have the same reduced-form covariance matrix. p-values are in parentheses.

and $\Sigma_{u,k}$ the regime-specific covariance matrix of the reduced-form residuals. Statistical inference is based on bootstrapping. Specifically, we generate 200 draws of the data using the regime-specific covariance matrices and for each draw we estimate the coefficients by minimizing the norm in (2.3). We compute p-values as the share of estimates beyond zero.

2.3.5 Identification and parameter stability tests

As outlined above, we use changes in volatility for identification. To uniquely determine A with this method, that is, to achieve identification in a statistical sense, the estimated variance-ratios of the uncorrelated structural shocks have to be sufficiently distinct across regimes (see Lütkepohl and Netšunajev, 2014). To check whether this is the case, we first study the variance-ratios $\phi_k^{S,S'} = \sigma_k^S/\sigma_k^{S'}$ for each pair of shocks (S, S'), which are given in Table 2.B.1 in Appendix 2.B. The estimated ratios and standard errors strongly suggest that for each pair there is at least one regime where the ratio changes sufficiently relative to the other regimes, that is, where the one-standard error intervals do not overlap. While these changes are indicative of statistically significant changes in volatility ratios, we also test more formally for identification. For each shock pair, we use a linear Wald test on the joint null hypothesis that the variance-ratio is the same across regimes, that is, that $\phi_1^{S,S'} = \phi_2^{S,S'} = \phi_3^{S,S'} = \phi_4^{S,S'}$, which would invalidate the identification of A. Inference in these tests is based on 200 bootstrap replications. Table 2.3.4 contains the Wald test statistics and the associated p-values, which show that for each pair of shocks the null hypothesis of no changes in volatility is strongly rejected by the data. The model is statistically fully identified.[8]

Table 2.3.4: Identification tests

H_0	Wald statistic (bootstrapped)	p-value (bootstrapped)
$\phi_1^{I,F}=\phi_2^{I,F}=\phi_3^{I,F}=\phi_4^{I,F}$	22.96	0.00
$\phi_1^{H,F}=\phi_2^{H,F}=\phi_3^{H,F}=\phi_4^{H,F}$	27.19	0.00
$\phi_1^{H,I}=\phi_2^{H,I}=\phi_3^{H,I}=\phi_4^{H,I}$	28.92	0.00

The table shows the Wald statistics and associated p-values of linear Wald tests on the joint null hypothesis that the estimated variance ratios of two structural shocks, $\phi_k^{S,S'} = \sigma_k^S/\sigma_k^{S'}$, are the same across volatility regimes, for each pair of structural shocks. Here, σ_k^S is the estimated variance of shock $S = F, I, H$ in regime $k = 1, \ldots, 4$. The tests are based on 200 bootstrap replications.

Having established statistical identification, we can test the assumption of a time-invariant impact matrix A as it becomes overidentifying with more than two regimes. For this we perform the following Likelihood ratio test: $LR = 2(logL_T - logL_T^r)$, where L_T^r is the maximum of the likelihood under the H_0 of time-invariant A and L_T is the maximum likelihood under H_1, which corresponds to the maximum likelihood of the reduced-form model with changes in volatility (compare Herwartz and Lütkepohl, 2014). The LR-statistic is 5.32 and the corresponding p-value is 0.50, not rejecting the constancy of A at conventional significance levels.

Finally, we investigate whether there is a break in the relation between the exogenous and endogenous variables on 21 July 2010. This dates splits the sample into a crisis and post-crisis half. It is, first, motivated by Cheng et al. (2015) who show, based on different position data however, that the behavior of financial investors can change in crises. Hedge funds, for example, may be more sensitive to prices or may increasingly trade for reasons unrelated to agricultural commodities, such as losses in other markets. Second, the date corresponds to the day of the Dodd-Frank act. The following regulations of commodity derivative markets may have changed the

[8] Recently, the possibility has been discussed that Wald tests for identification might not have the usual asymptotic χ^2-distributions as the model has unidentified parameters under the null (see Kilian and Lütkepohl, 2017). Our test statistics, however, are large and importantly test results are strongly supported by the distribution of the (bootstrapped) variance-ratios of the structural shocks across regimes. Hence, a lack of statistical identification should not be an issue.

functioning of futures markets. However, joint Chow tests for the three parameters of interest, referring to oil prices, the VIX, and crop conditions, do not reject the hypothesis of constant parameters across subsamples in the three equations. Moreover, all Chow tests of individual coefficient are insignificant, except for the effect of the VIX on hedge funds positions, where the null hypothesis of no break can be rejected at the 10% level. To account for the latter observation, we report subsample estimates in the robustness analysis of Section 2.5, which confirm our main results. All in all, the tests in this subsection indicate that the data support the assumptions of changing volatility during the sample period and time-invariant slope coefficients.

2.4 Empirical results: demand shocks and commodity futures returns

2.4.1 Interpretation of structural shocks

While we have shown that the model is statistically identified, our agnostic identification strategy has a well-known drawback. The structural shocks are more difficult to interpret since they are not based on a priori (zero or sign) assumptions or disaggregated data. We address this issue in several ways, in particular with the model outlined in Section 2.2 in mind.

First, we explore the significance of the exogenous control variables, meant to capture common factors affecting both futures prices and positions, to obtain an impression of the variation that remains in the reduced-form errors which are decomposed into the structural shocks. Table 2.4.1 shows the estimated effects of the most significant exogenous variables on the endogenous variables, corresponding to the entries in Λ in the *reduced-form* model (2.2). Standard errors are robust to heteroskedasticity and statistical significance is denoted by a, b, c for the 1%, 5%, and 10% level, respectively. The index of crop conditions has the expected negative effect on prices and is highly statistically significant. Better weather conditions lead to lower returns. While index investor positions are insensitive to crop conditions, net long positions of hedge funds decrease in response to improved physical supply conditions. Moreover, all three endogenous variables respond strongly to changes in the VIX and oil prices. In line with Tang and Xiong (2010) and Cheng et al.

(2015), these responses suggest that changes in the risk bearing capacity of financial investors, as proxied by changes in the VIX, or re-balancing motives, induced by oil price changes, are important drivers of financial investors' positions that induce similar movements in agricultural commodity futures returns. This co-movement indicates that changes in these variables are to a large extent transmitted to futures returns through financial investors.

Table 2.4.1: Effects of selected exogenous variables on the endogenous variables

	Exogenous variable		
	Crop Conditions Index	VIX	Oil price (orthogonal to VIX)
Futures returns	-0.36^a	-0.30^a	0.31^a
	(0.00)	(0.00)	(0.00)
Index inv. positions	0.09	-0.16^a	0.26^a
	(0.40)	(0.00)	(0.00)
Hedge funds positions	-0.23^a	-0.13^a	0.15^a
	(0.01)	(0.00)	(0.00)

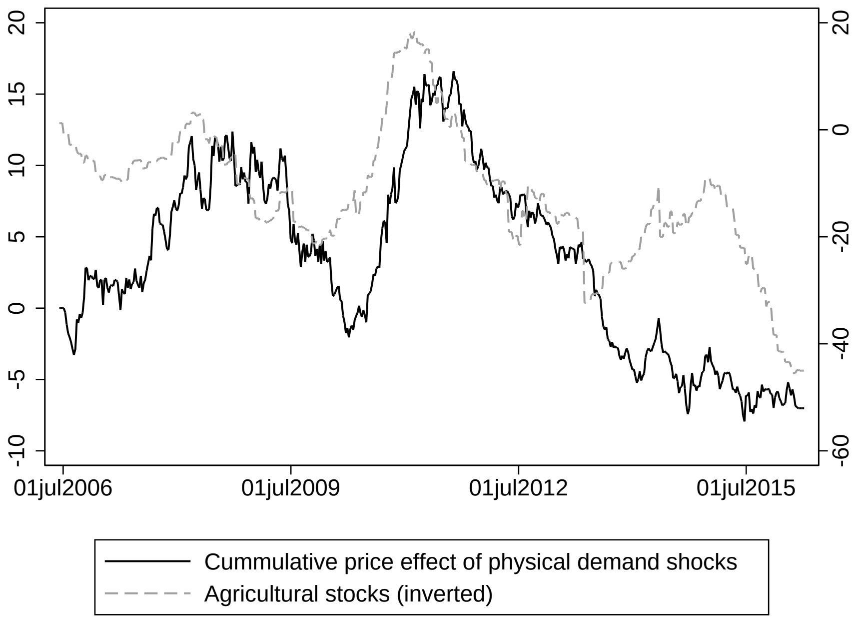
The table shows the effects of selected exogenous variables on the endogenous variables from the baseline VAR, obtained from estimating the rows of the reduced-form model (2.2). a, b, c denote significance at the 1%, 5%, 10% levels. Heteroskedasticity robust p-values below point estimates.

Given that the crop condition variable has a strong effect on returns and therefore appears to be a good measure of physical supply, we interpret the *structural shock* to the equation for futures returns, ϵ_t^F, that explains the largest part of the remaining variability in futures returns (see the forecast error variance decompositions in Section 2.4.3), as shifts in physical demand for agricultural commodities. In the notation of the theoretical model, the shock ϵ_t^F thus is interpreted as corresponding to the shock η. The interpretation of ϵ_t^F as a physical demand shock is also supported by model-external information. Figure 2.4.1 shows the cumulative effect of ϵ_t^F on futures prices, obtained from a historical decomposition, and the (inverted) level of agricultural stocks in the US.[9] The variables display a strong co-movement for most

[9] The stock variable is based on data from the US Department of Agriculture and described in Appendix 2.A. We do not include it into the model as it is only available at a monthly frequency. The stocks variable is not available for sugar, coffee, and cocoa that constitute close to 15 % of the aggregates. The relationship depicted in 2.4.1, however, is robust to excluding these commodities from the estimations.

of the sample, indicating that they are driven by similar underlying demand shifts, with changes in stocks leading the relation. As intuition would suggest, physical demand shifts tend to be buffered by inventories first and then show up in prices over time.

Figure 2.4.1: Agricultural stocks and cumulative effect of estimated agricultural specific demand shocks on the futures price



Note: The figure shows the level of agricultural stocks as reported by the US Department of Agriculture (grey line, right scale inverted) and the agricultural futures price implied by cumulated agricultural demand shocks obtained from a historical decomposition (black line).

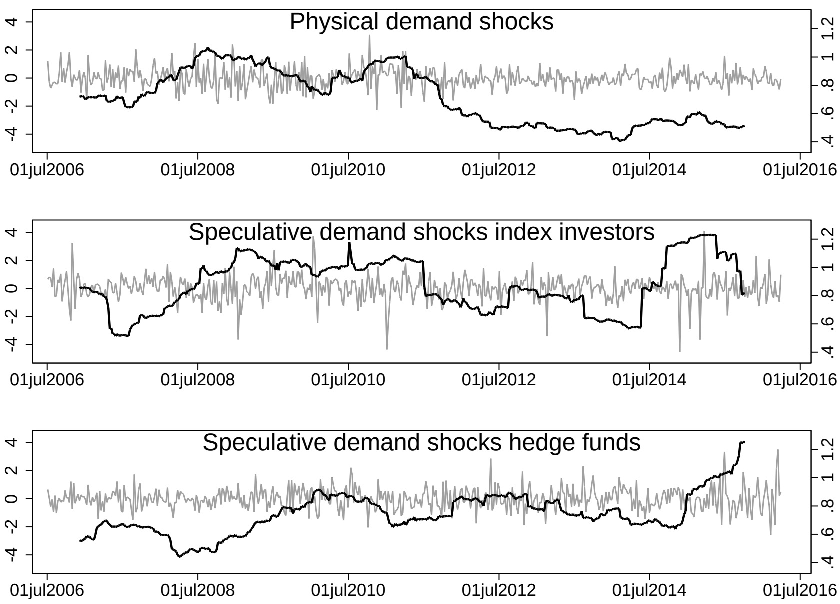
With physical demand as well as supply, financial market risk, and oil price changes accounted for, we interpret the remaining two structural shocks as investor-specific speculative demand shifts in line with the theoretical model. Specifically, structural shocks to the equation for index investor positions, ϵ_t^I, are interpreted as idiosyncratic shifts in their speculative demand. Analogously, we interpret structural shocks to the equation for hedge funds, ϵ_t^H, as demand shifts of hedge funds. Both ϵ_t^I and ϵ_t^H thus capture speculative demand shocks unrelated to changes in the risk bearing capacity or re-balancing motives as captured by VIX or oil price movements. This orthogonality allows complementing the analysis of Tang and Xiong (2010) and Cheng et al. (2015) by focusing on changes in speculative demand unrelated to these motives. Index investors, for example, may adjust positions in response to demand changes of their institutional or retail clients, and hedge funds might trade based on private information, say.

To further assess our labeling of these two structural shocks, we follow Herwartz and Lütkepohl (2014) and evaluate whether the structural shocks display distinct volatility patterns and higher volatility during those periods that we expect, given our shock interpretation. Figure 2.4.2 shows the structural shocks (grey line) and their centered 52-weeks rolling standard deviation (black line). The shocks ϵ_t^I display higher volatility during the first sample half and in 2014/15, whereas the volatility of ϵ_t^H-shocks increases sharply towards the very end of the sample. These changes in volatility correspond to the time-varying importance and activity of the two investor groups on commodity futures markets. While index investors were relatively more active in the first part of the sample, and in particular in 2007/08 where many institutional and retail investors sought exposure to commodities as a new asset class, a significant portion of these investors left the market afterwards when long-only strategies were no longer profitable as commodity prices experienced sharp boom-bust cycles. Their share in total long positions, for instance, declined from 32% in 2008 to 25% in 2014. In contrast, hedge funds employ trading strategies which allow them to earn positive returns in periods of both rising and declining prices (see Mayer, 2009). Their activity was relatively stable during most of the sample and only started to intensify when commodity prices began a steady decline from 2014 onwards.

2.4.2 Contemporaneous shock propagation

Having labeled the structural shocks, we now present their effects on the endogenous variables. Table 2.4.2 shows the estimated contemporaneous impact matrices A and A^{-1}. We do not show impulse response functions, as they do not provide additional insights given that there is virtually no persistence in the differenced data. To evaluate Hypothesis 1, we focus on the impact of speculative demand shocks on futures returns, that is, on parameters $\alpha_{1,2}$ and $\alpha_{1,3}$. According to the direct effects, demand shocks of both investor groups lead to a significant contemporaneous increase in commodity futures returns. The point estimates are both significant at the one percent level. A similar conclusion can be drawn from the overall effects

Figure 2.4.2: Volatility patterns of structural shocks



Note: The figure shows the estimated structural shocks (grey line, left axis) together with their (centered) 52 weeks rolling standard deviations (black line, right axis). The estimated shocks are based on a structural VAR identified through heteroscedasticity.

which take into account all contemporaneous feedback among the endogenous variables. The overall effects imply that an exogenous increase of index investors' net long demand by one standard deviation leads to an increase in commodity futures returns by 0.15 standard deviations within the same week. The effect for hedge funds is even stronger. Here, a demand shock increases returns by 0.39 standard deviations. Together, the estimates support Hypothesis 1. Speculative demand shocks that increase net long positions impact positively on futures returns. Returning to the motivating theoretical model, these results imply that the demand curve of hedgers (or the respective counter-party in a general setting) is not infinitely elastic with respect to changes in the futures price and that exogenous changes in speculative demand consume liquidity. In other words, hedgers require a compensation for meeting speculators net long demand.

Regarding Hypothesis 2, the point estimates for the effect of physical demand shocks on speculators' positions provide mixed evidence (parameters $\alpha_{2,1}$ and $\alpha_{3,1}$). Positions of index investors do not respond significantly to physical demand shocks,

Table 2.4.2: Contemporaneous effects among endogenous variables

	Impulse		
Response	Futures return	Index inv. pos.	Hedge funds pos.
	Direct effects (A)		
Futures Return	1.00	-0.16	-0.33
	.	.a	.a
Index inv. positions	0.03	1.00	-0.16
	.	.	.
Hedge funds positions	-0.29	0.08	1.00
	.a	.	.
	Overall effects (A^{-1})		
Futures Return	1.11	0.15	0.39
	.a	.b	.a
Index inv. positions	0.01	0.99	0.16
	.	.a	.
Hedge funds positions	0.32	-0.04	1.10
	.a	.	.a

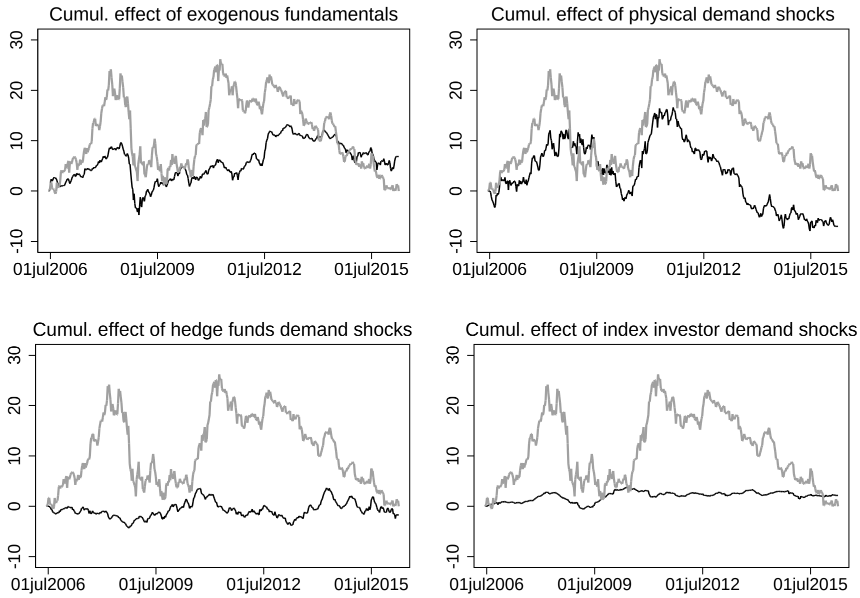
The table shows the estimated impact effects of structural shocks of one standard deviation on the endogenous variables, based on a structural VAR identified through heteroskedasticity. Impulse variables are in columns, response variables are in rows. The sample period is 04 July 2006 - 29 March 2016. The number of observations used for identification is 358. .a, .b, .c below point estimates denote significance at the 1%, 5%, 10% level, respectively.

whereas for hedge funds we find a significant positive effect. For this trader group, a physical demand shock of one standard deviation leads to an increase in net long positions by 0.32 standard deviations. Through the lens of the theoretical model, the significant response of hedge funds' net long positions to physical demand shocks suggests that - through their increased net long demand - they provide liquidity to producers who have higher hedging needs.

2.4.3 Drivers of commodity futures prices

We next assess the relevance of alternative explanations for the commodity price swings contained in our sample by means of a historical decomposition of the futures return series. The upper left panel of Figure 2.4.3 contains the cumulated changes in the futures price and the cumulative effect of the exogenous variables on the futures price. It shows that the exogenous variables are an important driver of futures prices and explain in particular the secular price movements well. There is a strong co-movement between both series. Local supply conditions as well as changes in the VIX and oil prices explain about half of the boom-bust cycle of commodity prices in 2008/09, and a relevant part of the steady price decline from 2012 onwards.

Figure 2.4.3: Cumulative effects of exogenous variables and structural shocks on agricultural futures prices.



Note: The figure shows the cumulative change of the agricultural futures price (grey line) and the cumulative effect of the exogenous variables and of the structural shocks (black lines) on agricultural futures prices over the sample 04 July 2006 - 29 March 2016.

The figure also shows the cumulative effects of physical and speculative demand shocks on futures prices. The cumulative effects are based on a historical decomposition of the futures return series which yields the weekly contribution of each structural shock to the futures return. The top right panel reveals that, next to the exogenous variables, physical demand shocks are the other main driver of futures prices. They explain approximately the other half of the boom-bust cycle in 2008/09, nearly the entire price surge in 2010/11, and account for a major fraction of the subsequent decline. Together, the top two panels suggest that the exogenous variables and physical demand shocks are the main drivers of prices and in particular explain the longer-term price movements.

Compared to these main forces, the effects of speculative demand shocks on futures prices appear more modest, but are not negligible. Especially speculative demand shifts of hedge funds explain the higher frequency (that is, short-term) movements in returns well, in particular in the second sample half. Speculative demand shocks of index investors, on the other hand, seem to play only a limited role. This preliminary conclusion does not imply, however, that index investors are unimportant for commodity price formation in general. First, our results indicate that they transmit changes in the VIX or price of oil to futures prices (see Section 2.4.1). Second, while their impact is apparently small on average, they might have relevant effects when their volatility increases. We investigate this issue next.

Specifically, we compute one week ahead forecast error variance decompositions to quantify the regime-specific and average economic significance of the different types of structural shocks. Since we have four different volatility regimes, we obtain four decompositions. They yield the percentage contribution of each shock to the variance of the endogenous variables in each regime. In addition, we compute a weighted average of the regime-specific decompositions to measure the average importance of the shocks over the full sample, using the number of observations per regime as weights. Table 2.4.3 shows that speculative demand shocks explain 15 percent of the variability in futures returns in regime 1, where all shocks display low volatility. Shocks of hedge funds are important, whereas index investor demand shifts play a more limited role. The positions of index investors, in turn, are almost entirely driven by own shocks, consistent with their trading strategies. Hedge funds positions, on the other hand, also respond to physical demand shocks.

This asymmetry between investors increases in regime 2, where physical demand shocks are more volatile relative to the other shocks. Physical demand shocks now explain 19 percent of the variation in hedge funds positions and still nothing of the changes in index investor positions. Reversely, the importance of speculative demand shocks in futures price determination increases and becomes important when positions are more volatile. In regime 3, demand shifts of index investors explain 9 percent of the variation in futures returns. In regime 4, hedge funds demand shocks account for nearly one third of the fluctuation. Finally, the weighted forecast error

Table 2.4.3: Forecast error variance decompositions

	Impulse		
	Futures return	Index inv. pos.	Hedge funds pos.
Regime 1			
Futures return	0.85	0.02	0.13
Index investor positions	0.00	0.98	0.02
Hedge funds positions	0.07	0.00	0.93
Regime 2			
Futures return	0.95	0.01	0.04
Index investor positions	0.00	0.98	0.02
Hedge funds positions	0.19	0.00	0.81
Regime 3			
Futures return	0.73	0.09	0.18
Index investor positions	0.00	0.99	0.01
Hedge funds positions	0.04	0.00	0.95
Regime 4			
Futures return	0.68	0.02	0.30
Index investor positions	0.00	0.95	0.05
Hedge funds positions	0.02	0.00	0.98
Weighted FEVD			
Futures return	0.80	0.03	0.17
Index investor positions	0.00	0.97	0.03
Hedge funds positions	0.07	0.00	0.93

The table shows the forecast error variance decompositions over the horizon of one week for volatility regimes 1-4 and a weighted average of these, using the number of observations per regime as weights, based on a structural VAR identified through heteroskedasticity. In regime 1 all structural shocks have low volatility. In regimes 2-4 the volatility of shocks to, respectively, physical demand, index investor demand, and hedge funds demand is high relative to the other structural shocks.

variance decomposition reveals that taking into account these changes in volatility increases the importance of speculative demand shifts relative to their importance in the tranquil period. Financial demand shocks on average account for almost one fifth of the variance of futures returns. All in all, however, the decompositions show that, next to exogenous fundamentals, shocks to physical demand are the main driver of commodity prices.

2.5 Sensitivity analysis

2.5.1 Alternative definitions of volatility regimes

As a final step in the analysis, we assess the sensitivity of our main results to various modifications of the model. First, we analyze the robustness of the results to changing the calibration and definition of the volatility regimes. In the baseline specification, we use 15-weeks rolling standard deviations of the residuals and a threshold of 1 standard deviation above which underlying observations are classified into volatility regimes. We investigate how the results change when we modify either the threshold (from 1 to 1.1 and 1.2 standard deviations, respectively) or the length of the window (from 15 weeks to 10, 12, and 18, respectively). Table 2.5.1 shows that the main results are robust to these alterations.

Also note that in the main specification we drop some observations from the computation of the regime-specific covariance matrices as they do not fit into any of the four volatility regimes. As a further robustness check, we adjust the regime definitions so that only a few observations are discarded. In contrast to the baseline definitions, now the second regime contains all observations where the residuals for the futures returns are volatile and the residuals for index investor positions tranquil, independent of the volatility of hedge fund positions, and vice versa for the third regime. Again, our main results are insensitive to these changes.[10]

2.5.2 Additional group of traders

Second, aside from commercials, non-commercials, and index investors, the SCOT reports contain the additional category of traders called ‘non-reportables’ (see above). To assess the sensitivity of our estimates for index investors and hedge funds to including another trader group, we add the net long positions of non-reportables as a fourth endogenous variable to the baseline model. Adding this fourth variable requires adjusting the volatility regime definition slightly. To account for the additional shock, we add a fifth regime where only the reduced-form residuals of the

[10] The same holds when, instead of the volatility in hedge fund positions, the volatility in futures returns or index investor positions is disregarded for the computation of the other two volatility regimes, respectively. With the reported combination, however, the lowest number of observations is discarded.

Table 2.5.1: Effects between investors' positions and futures prices with different regime definitions

Alt. regime definitions		*1*	*2*	*3*	*4*	*5*	*6*
Index Inv. -> Fut. Return	0.15	0.17	0.20	0.14	0.20	0.24	0.14
	.b	.b	.a	.b	.a	.a	.b
Hedge funds -> Fut. Return	0.39	0.45	0.44	0.41	0.39	0.49	0.39
	.a	.a	.a	.a	.a	.a	.a
Fut. Return -> Index Inv.	0.01	-0.08	-0.07	0.04	0.00	-0.09	-0.02
	.	.	.	.	.	.	.
Fut. Return -> Hedge funds	0.32	0.26	0.33	0.30	0.31	0.18	0.38
	.a	.a	.a	.a	.a	.b	.a
Window	15	15	15	10	12	18	15
Threshold	1	1.1	1.2	1	1	1	1
Regime definition	main	main	main	main	main	main	other

The table shows the estimated structural *overall* effects between futures prices and investors' positions, based on market-specific six-variable structural VAR models. The arrows indicate the relation between impulse and response variable. The sample period is 04 July 2006 - 29 March 2016. .a, .b, .c denote significance at the 1%, 5%, 10% levels.

new equation display volatility above the threshold. To obtain a sufficient number of observations per regime, we use 12-weeks instead of 15-weeks rolling standard deviations of the residuals.

Table 2.5.2 contains the results. It shows that the contemporaneous structural relations between futures returns, hedge fund positions, and index investor positions are robust to this model alteration. Speculative demand shocks still impact significantly positive on returns, with the size of the coefficients being comparable to the baseline case. On the other hand, net long positions of hedge funds, but not of index investors respond significantly to the physical demand shock.

2.5.3 Alternative sample periods

Third, while statistical tests do not reject the assumption of a constant impact matrix A, one Chow test indicates the possibility that the impact of the VIX on hedge funds positions might vary between crisis and tranquil times (see Section 2.3.5). Also, financial speculators might have had stronger incentives to be clas-

Table 2.5.2: Effects between investors' positions and futures prices in a model with non-reportables

	baseline	model with non-reportables
Index Inv. -> Fut. Return	0.15	0.25
	.b	.b
Hedge funds -> Fut. Return	0.39	0.32
	.a	.a
Fut. Return -> Index Inv.	0.01	-0.04
	.	.
Fut. Return -> Hedge funds	0.32	0.34
	.a	.a

The table shows the estimated structural effects between futures prices and investors' positions, based on the a structural VAR model where non-reportables net long positions are added as a fourth endogenous variable. The arrows indicate the relation between impulse and response variable. The sample period is 04 July 2006 - 29 March 2016. .a, .b, .c denote significance at the 1%, 5%, 10% levels.

sified as hedgers after the introduction of the Dodd-Frank act, possibly affecting the categorization of traders (see Section 2.3.3). Therefore, we split our sample in a pre-crisis/crisis sample and a post crisis sample and carry out individual estimations for the two samples. Specifically, we try two different break dates. The first is the implementation of the Dodd-Frank act on 07 July 2010. The second is 07 June 2011 following Cheng et al. (2015), who show that the behavior of financial investors can change between crisis and tranquil times.

Table 2.5.3 contains the parameters of interest for the different sub-sample estimations. It shows that the relation between structural shocks, hedge funds positions, and futures returns is basically the same across the two samples. The impact of index investors speculative demand shocks, on the other hand, is significant for the post-crisis samples, but less or not at all significant in the crisis. Whether this is indeed due to a less significant impact of shocks to index investor positions during the crisis, or due to more difficulties in identifying the shock in the particular sample, cannot be distinguished.

Table 2.5.3: Effects between investors' positions and futures prices in subsamples

	crisis		post-crisis	
	1	*2*	*3*	*4*
Index Inv. -> Fut. Return	0.36	0.23	0.18	0.20
	.c	.	.a	.b
Hedge funds -> Fut. Return	0.33	0.35	0.39	0.50
	.b	.a	.a	.a
Fut. Return -> Index Inv.	-0.11	-0.13	-0.12	-0.10
	.	.	.	.
Fut. Return -> Hedge funds	0.27	0.31	0.37	0.26
	.c	.b	.a	.c
Sample start	04jul2006	04jul2006	10aug2010	21jun2011
Sample end	27jul2010	07jun2011	29mar2016	29mar2016

The table shows the estimated structural effects between futures prices and investors' positions, based on the baseline structural VAR model, for different start and end dates of the sample. The arrows indicate the relation between impulse and response variable. .a, .b, .c denote significance at the 1%, 5%, 10% levels.

2.5.4 Single markets

Fourth, we assess whether our main results based on the aggregated indexes reflect a general pattern on agricultural futures markets or whether they are driven by a few (dominant) markets. To this end, we estimate the structural model (2.1) for each individual market, that is, we replace the indexes in y_t by market-specific variables and, regarding the exogenous variables, use market-specific crop conditions.[11] For the market-specific models, we calibrate the thresholds and windows for the volatility regimes individually. This ensures that on each market there are enough observations per volatility regime and that statistical identification is achieved every time.

Table 2.5.4 shows the results which are ordered by market size. They closely mirror the findings for the aggregate level. The effect of demand shifts of hedge funds on futures returns is positive and highly statistically significant in all eleven markets. Quantitatively, the impact varies between 0.26 and 0.50, spanning the corresponding point estimate for the aggregate level of 0.39. Similarly, demand shocks of index

[11] For the meat markets, we use crop conditions for corn. For sugar, coffee, and cocoa, no measure of crop conditions is available.

investors impact significantly on returns in nine of eleven markets. Regarding the physical demand shocks, in ten of the eleven markets hedge funds systematically increase their long exposure in response to the shock. Index investors, in contrast, show significant reactions to the shock only in three markets.

Table 2.5.4: Effects between investors' positions and futures prices on individual markets

	Corn	*SRW Wheat*	*Live Cattle*	*Soybeans*	*Sugar*	*Lean Hogs*
Index Inv. -> Fut. Return	0.11	0.20	0.14	0.07	-0.00	0.20
	.c	.a	.c	.	.	.b
Hedge funds -> Fut. Return	0.47	0.37	0.26	0.46	0.42	0.30
	.a	.a	.a	.a	.a	.a
Fut. Return -> Index Inv.	0.11	-0.07	0.07	0.14	-0.07	-0.12
	.	.	.	.c	.	.c
Fut. Return -> Hedge funds	0.23	0.23	0.18	0.32	0.27	0.13
	.a	.a	.a	.a	.a	.c
	Cotton	*HRW Wheat*	*Coffee*	*Feeder Cattle*	*Cocoa*	
Index Inv. -> Fut. Return	0.12	0.25	0.11	0.10	0.11	
	.c	.b	.b	.c	.c	
Hedge funds -> Fut. Return	0.33	0.36	0.50	0.34	0.42	
	.a	.a	.a	.a	.a	
Fut. Return -> Index Inv.	-0.00	0.03	0.01	0.05	0.23	
	.	.	.	.	.a	
Fut. Return -> Hedge funds	0.18	0.17	0.28	0.13	0.16	
	.a	.b	.a	.	.a	

The table shows the estimated structural effects between futures prices and investors' positions, based on market-specific structural VAR models. The arrows indicate the relation between impulse and response variable. The sample period is 04 July 2006 - 29 March 2016. .a, .b, .c denote significance at the 1%, 5%, 10% levels.

2.5.5 Narrative approach to identify volatility regimes

Fifth, in the baseline specification we have used a statistical approach to determine the volatility regimes. An alternative used in the literature is a narrative approach (see, Rigobon, 2003), which specifies volatility clusters based on a time line of major economic and political events. As a further sensitivity analysis, we thus apply such a narrative approach and then re-estimate our baseline model using the narratively-determined regimes. Specifically, we divide the sample into four volatility regimes. The first regime runs from the beginning of the sample until the bankruptcy of Lehman Brothers in September 2008. Regime 2 and 3 are then separated by the implementation of the Dodd-Frank act in July 2010, while the last Regime begins with the Federal Reserve ending its asset purchases in October 2014, corresponding to the end of quantitative easing. Table 2.5.5 displays matrix A^{-1} as estimated using the narrative volatility regimes. It shows that results of this exercise are similar to the baseline specification. Shocks to net long positions of both investor groups impact positively on futures returns, while only positions of hedge funds respond significantly to the physical demand shock.[12]

Table 2.5.5: Contemporaneous effects among endogenous variables - narrative regime definition

	Impulse		
Response	Futures return	Index inv. pos.	Hedge funds pos.
Futures return	1.11	0.22	0.43
	.a	.b	.a
Index inv. positons	-0.04	0.99	0.02
	.	.a	.
Hedge funds positions	0.32	0.06	1.12
	.a	.	.a

The table shows the estimated impact effects of structural shocks of one standard deviation on the endogenous variables, based on a structural VAR identified through heteroskedasticity. Impulse variables are in columns, response variables are in rows. The sample period is 04 July 2006 - 29 March 2016. .a, .b, .c below point estimates denote significance at the 1%, 5%, 10% level, respectively.

[12] Results of the narrative approach are robust to the exact start and end date of the regimes. To assess this, we replaced one-by-one the bankruptcy of Lehman Brothers with the failure of Bear Stearns in March 2008, the implementation of Dodd-Frank with the second breakpoint used in Section 2.5.3, and the end of asset purchases with announcements of Federal Reserve official to taper quantitative easing.

2.6 Conclusion

This study provides new evidence on the impact of financial investment on price formation in commodity futures markets. We use publicly available data on net long positions of hedge funds and index investors on agricultural commodity futures markets from SCOT reports of the Commodity Futures Trading Commission, and include them in a vector autoregression along with the corresponding futures price. Controlling exogenously for physical supply and financial market risk and using the heteroskedasticity in the data, we identify idiosyncratic shocks to speculative demand of both investor groups.

Our results suggests that speculative demand shocks of both index investors and hedge funds impact significantly and positively on commodity futures returns. The shocks appear also economically relevant as they account for about one fifth of futures return fluctuations on average, and for up to one third of the variability in returns during periods of high speculative demand volatility. Overall, physical demand shocks and exogenous variables explain most of the secular movement in commodity futures prices.

The findings complement recent studies that detect significant effects of financial investments on commodity futures prices based on highly disaggregated and partly private data (see Cheng et al., 2015, and Henderson et al., 2015). While these data potentially allow for a more precise estimation of the effects of speculative trading on commodity futures returns, and simpler and more transparent econometric approaches, one advantage of the statistical approach employed in this study is that it allows using publicly available data such that the analysis can be replicated and readily updated in the future. Moreover, the SVAR model allows quantifying not only the statistical significance of the effects of speculative demand shocks, but also their economic importance both on average in the sample and during historical episodes.

Appendix

2.A Data and sources

Table 2.A.1: Data construction and sources

Variable	Construction and source
Futures prices for agricultural commodities	Nearby (next-to-maturity) futures prices of eleven agricultural commodities. As position data measure open positions on each Tuesday, we use Tuesday futures prices. If Tuesday is not a trading day, we use the closing price of the trading day before. Individual returns are computed as log differences and aggregated into one variable by multiplying them with their respective weights in the S&P GSCI. Our results are robust to first computing an aggregate futures price index with rescaled individual series and then taking the return of this index. Source: Datastream.
Investors' positions	Aggregated data on open positions of different trader groups on eleven agricultural markets. The underlying reports divide traders into four categories: index investors, non-commercial traders, non-reporting traders, commercial traders. The index investors category includes positions of managed funds, pension funds, and other investors that are generally seeking exposure to a broad index of commodity prices as an asset class in an unleveraged and passively-managed manner, as well as positions of entities whose trading predominantly reflects hedging of over-the-counter transactions involving commodity indices - for example, a swap dealer holding long futures positions to hedge a short commodity index exposure opposite institutional traders. Traders are classified as commercials if the trader uses futures contracts in that particular commodity for hedging as defined in CFTC Regulation 1.3, 17 CFR 1.3(z). Examples are entities that predominantly engages in the production, processing, packing or handling of a physical commodity and use the futures markets to manage or hedge risks associated with those activities. The non-commercial category contains speculative traders like hedge funds, registered commodity trading advisors (CTAs), registered commodity pool operators (CPOs), or unregistered funds identified by CFTC.

	The non-reporting category contains traders that hold positions below specific reporting levels set by CFTC regulations. To construct aggregate position indexes, we combine net long positions in individual markets in two steps. First, we divide by average open interest in each market in 2010 to dispose of the underlying units (bushels, pounds, et cetera). Then, we average over markets with the respective weights. Source: CFTC SCOT reports.
Index of crop conditions	Weather conditions are measured following Bruno et al. (2017). We use weekly crop conditions reports of the US Department of Agriculture (USDA) which survey the condition of cotton, corn, soybeans, and wheat plants in major producing US states. On a given week, a percentage of crops is assessed to be in an 'excellent', 'good', 'fair', 'poor', or 'very poor' condition. We weight the assessments using a linear scheme to construct a measure of individual crop conditions. The resulting series are set to zero when no information is available, that is when there is nothing yet in the ground. We construct a weather conditions index based on the eight US based commodities using the (adjusted) S&P GSCI weights. Thereby, additional weight, namely that for live cattle, lean hogs, and feeder cattle, is given to corn as it is the most import source of feed for cattle and pigs. As a robustness check, we also exclude the weights for meat commodities, as they are not directly affected by the weather. This yields very similar results. For the included commodities not produced in the US (sugar, coffee, and cocoa) no such weather index is available. As they constitute less than 15 % of our aggregate, it is not surprising that the supply measure is, nevertheless, highly significant in the baseline estimations. Our main results are robust to dropping the non US commodities, the meat commodities, or even both. For the individual market estimates, we use the weather index for corn for live cattle, lean hogs, and, feeder cattle.
Oil Price	WTI oil price. Source: St. Louis Fed FRED database.
VIX	CBOE Volatility Index: VIX, Source: St. Louis Fed FRED database.
Fed balance sheet	Total assets of the Fed. Source: St. Louis Fed FRED database.

Agricultural stocks	Actual agricultural stocks for eight US based commodities, constructed as in Bruno et al. (2017). Meat stocks: USDA total storage figures for beef and pork (excluding frozen ham). For grain and cotton stocks: monthly stock forecasts reported in the current USDA forecasts of US supply-use balances of major grains.

2.B Variance-ratios of the structural shocks

Table 2.B.1: Variance-ratios of the structural shocks per volatility regime

Regime	(1)	(2)	(3)	(4)
σ_k^I/σ_k^F	1.33	0.58	7.66	1.74
	(0.24)	(0.16)	(2.39)	(0.39)
σ_k^H/σ_k^F	1.19	0.37	2.05	3.65
	(0.24)	(0.11)	(0.61)	(0.88)
σ_k^I/σ_k^F	0.91	0.66	0.28	2.14
	(0.18)	(0.21)	(0.08)	(0.52)
Observations	152	59	51	96

The table shows the estimated volatility-ratios of the structural shocks of the endogenous variables in the different regimes. The shocks are named after the equation they appear in. Bootstrapped standard deviations of the ratios are reported in parentheses.

CHAPTER 3

The Anchoring of Inflation Expectations: Evidence from SVARs identified with Macroeconomic News Announcements[1]

3.1 Introduction

Inflation expectations are an important determinant of future inflation and crucial for the conduct of monetary policy. Both financial markets participants and central bankers therefore closely monitor expectations, in particular those that are timely available from inflation-indexed bonds or inflation swaps. Monetary policy makers increasingly explain their decisions with the need to keep expectations well anchored (see Bernanke, 2007, Yellen, 2015, or Draghi, 2015b). Yet, it is not obvious how to define and measure the anchoring of expectations empirically.

In the literature, it is usually assumed that long-term inflation expectations are well-anchored if they do not respond to news about short-run macroeconomic developments. This reasoning builds on the notion that the news should have no impact on rational long-term expectations as long as the central bank is fully credible and transparent about an explicit or implicit inflation target. A significant reaction of expectations to macroeconomic news, in turn, implies a lack of credibility and de-anchored expectations. Following Gürkaynak et al. (2010b), this criterion is often assessed in news-regression frameworks by regressing a measure of long-term expec-

[1] I thank Benjamin Beckers, Dieter Nautz, Till Strohsal, Mathias Trabandt, Lars Winkelmann, and participants of a seminar at Freie Universität Berlin for helpful comments and suggestions.

tations on macroeconomic news announcements (MNAs). MNAs are observables that capture the surprise component of macroeconomic data releases and are thus used to proxy macro news (see, among others, Gürkaynak et al., 2010b, Beechey et al., 2011, Autrup and Grothe, 2014, Ehrmann, 2015, Bauer, 2015, or Speck, 2016). A significant impact response to some MNAs on their release day is then interpreted as a de-anchoring.

Recent empirical studies take a more dynamic perspective. They assume that shocks can drive long-term inflation expectations away from a central bank target. If expectations return to the target eventually, they are still regarded as anchored in the long run. Most contributions with dynamic models, however, are based on univariate reduced-form equations (Mehrotra and Yetman, 2014, Strohsal and Winkelmann, 2015, Strohsal et al., 2016). This makes it difficult to economically interpret the shocks that drive expectations away from a target. Moreover, univariate models do not take into account that shock transmission to long-term expectations might run through short-term expectations, albeit spillovers from short-term expectations are found to be a potential source of de-anchoring (see Jochmann et al., 2010).

Nautz et al. (2016), in contrast, use a bi-variate structural vector autoregressive (SVAR) model containing both short-term and long-term US inflation expectations to study the degree of anchoring with respect to macroeconomic news. Specifically, they identify an unobservable macro news shock, which is supposed to capture all new information about short-run macroeconomic developments. Identification is based on the restriction that the shock has no permanent impact on long-term expectations, or, put differently, it is based on the assumption that expectations are anchored in the long run. The restriction and, thus, the long-run anchoring, is tested using a statistical identification approach based on the heteroskedasticity in the data, and not rejected. With the identified shock, the authors then also evaluate the anchoring of inflation expectations in the short run by measuring the transitory impact of the shock on expectations and its contribution to their variance and historical development.

However, it is not obvious that the shocks Nautz et al. (2016) identify with a statistical approach indeed capture news about the macro economy. If these shocks do not (solely) relate to macro news, then imposing a long-run restriction for identification will also impose a long-run anchoring of inflation expectations without properly assessing it. In fact, when structural shocks are identified using statistical properties of the data, usually additional economic information has to be used to label them (compare, for instance, Herwartz and Lütkepohl, 2014, or Chapter 2). Netsunajev and Winkelmann (2016), for instance, use a heterosdesaticity based approach to analyze spillovers between inflation expectations in different countries in a SVAR. To attach an economic meaning to the statistically identified shocks, they regress them on MNAs from different geographical areas after the estimation.

In this paper, I provide new evidence on the anchoring of US inflation expectations in a structural VAR framework, where the identified macro news shocks correlate with observable information on macroeconomic surprises, the MNAs, by construction. Specifically, I use the MNAs as proxy variables to identify the unobservable macro news shock in the SVAR, following the identification through proxy variables approach developed by Stock and Watson (2012) and Mertens and Ravn (2013). In this way, no statistical approach or long-run restriction has to be imposed, while the anchoring of expectations can be assessed based on the estimated impulse response functions of the model.[2]

A necessary condition for the identification approach to produce meaningful results is a sufficiently high correlation between the proxy variables and the reduced-form VAR residuals. To fulfill the condition, I employ daily financial market based measures of inflation expectations. The high number of daily observations also allows me to focus the investigation on the period after the financial crisis, for which a particular de-anchoring of expectations is found in news-regressions (see Galati et al., 2011, Autrup and Grothe, 2014, or Nautz and Strohsal, 2015). Nautz et al. (2016), in contrast, use monthly survey data and have to assume a constant degree of expectations anchoring over a period from 1991 to 2015 for identification.

My main results are as follows. I find that long-term inflation expectations respond significantly to the macro news shock on impact. The estimated impulse response functions show that the shock phases out after about 50 trading days, indicating a de-anchoring of expectations in the short run, but not in the long run.

[2] A seminal contribution by Gertler and Karadi (2015) uses the proxy variables approach to study the impact of monetary policy shocks. They use the immediate financial market reaction to a monetary policy decision, i.e. the unexpected component of the decision, to identify the shock. My strategy is thus also based on their work as I employ a surprise component to identify a shock. However, in line with news-regressions, I take a broad set of MNAs and study their joint impact on inflation expectations.

Moreover, the macro news shock is estimated to be a non-negligible source of distortion as it explains more than 10 % of the forecast error variance of long-term inflation expectations. These results are in line with the findings of Nautz et al. (2016) and imply that the latter are robust to aligning the identification approach more closely to the news-regression literature and observable information on macroeconomic surprises. A sensitivity analysis nevertheless indicates, that solely imposing a long-run restriction for identification would imply a lower impact of the shock on expectations in the daily financial market dataset. Moreover, in line with results from news-regressions, I find a stronger impact of macro news shocks on long-term expectations after the financial crisis compared to the pre-crisis period.

I also carry out a counterfactual exercise based on the estimated model to analyze whether macro news shocks are an important driver of the historical development of inflation expectation over the sample period. I find that macro news shocks are not responsible for the very low levels of financial market based long-term expectations observed between mid-2014 and mid-2016, where expectations have fallen to levels as low as 1.5 %. Given a central bank target of 2 %, this result indicates that future assessments of expectations anchoring should not solely focus on macroeconomic news as a potential source of de-anchoring.

The remainder of the paper is structured as follows. In Section 3.2, I review the approaches used to assess the anchoring of expectations in more detail and explain how my modeling approach relates to them. The structural VAR model along with the identification strategy is then described in Section 3.3. Section 3.4 contains a description of the expectations data and the MNAs that are used. Empirical results are given in Section 3.5, while robustness and sensitivity of the baseline results are evaluated in Section 3.6. Section 3.7 concludes.

3.2 How to measure the anchoring of inflation expectations

3.2.1 News regressions

The most widely used approach in the literature to assess the anchoring of inflation expectations are news-regressions (see, among others, Gürkaynak et al., 2010b, Beechey et al., 2011, Autrup and Grothe, 2014, Ehrmann, 2015, Bauer, 2015, or

Speck, 2016). Following Gürkaynak et al. (2010b), a measure of long-term inflation expectations is regressed on a set of macroeconomic news announcements (MNAs).[3] MNAs capture surprise realizations of macroeconomic variables and are used to proxy macro news. They are calculated as the difference between the realized and the expected value of a variable on its announcement day, whereas the expected value stems from a survey of professional or consumer forecasts. Specifically, news-regressions take the following form:

$$\Delta\pi^e_{l,t} = c + \beta' X_t + u_t, \tag{3.1}$$

where $\Delta\pi^e_{l,t}$ is the first difference of a measure of long-term expectations and X_t is a vector of MNAs. A significant response of expectations to one or more of the MNAs on their announcement day ($\beta \neq 0$) is then interpreted as a de-anchoring.

News-regressions assess the anchoring of inflation expectations with respect to a structural source of de-anchoring, macroeconomic news, by using an observable proxy for it. The approach, however, relies on crucial assumptions. Failure to reject the null hypothesis of $\beta = 0$ might just stem from relevant MNAs being omitted from the regression. Similarly, while macroeconomic news arrive on the market every day, news-regressions capture only news on release days of variables for which MNAs are available. Also, MNAs might not be completely free from measurement error as, for instance, expectations from a particular survey might differ from those of the whole market or those of all consumers at some points in time.

Most importantly, rejecting the null hypothesis of $\beta = 0$ is only indicative of a de-anchoring of inflation expectations on announcement days of MNAs and thus in the very short run. Ignoring the dynamic reaction of expectations to macro news over time, this potentially leads to misleading conclusions about the degree of de-anchoring. If the response of expectations fades out quickly, the problem of de-anchoring does not appear to be severe. On the other hand, if expectations return to their pre-news level only slowly or not at all, this indicates a strong degree of de-anchoring.

[3] In an earlier study, Gürkaynak et al. (2005b) assess the sensitivity of long-term interest rates to macroeconomic news and find a significant impact of data surprises on long-term forward rates. Accordingly, they conclude that agents adjust their inflation expectations in response to the news.

3.2.2 Dynamic models

Only few studies analyze the anchoring of expectations from a more dynamic perspective. These studies are based on the notion that shocks can push long-term inflation expectations away from an explicit or implicit central bank target. However, if expectations return to the target eventually, they can still be regarded as anchored in the long run. The degree of anchoring in the short(er) run is then assessed by studying how long it takes for expectations to return to the target: the more persistent the reaction to shocks, the more severe the de-anchoring.

Mehrotra and Yetman (2014), Strohsal and Winkelmann (2015), and Strohsal et al. (2016) take such a dynamic perspective. Their contributions, however, are based on reduced-form equations. This makes it difficult to interpret the estimated responses of inflation expectations to shocks economically, as the shocks are not stemming from a structural econometric model. Compared to news-regressions, the de-anchoring of expectations cannot be directly linked to an underlying structural source. Moreover, the studies use univariate models that only partially or not at all account for the joint movement of short-term and long-term inflation expectations. Yet, there is evidence that shock transmission to long-term expectations running through short-term expectations is a potential source of de-anchoring (see, for instance, Jochmann et al., 2010).

Nautz et al. (2016), in contrast, use a bi-variate structural vector autoregressive (SVAR) model containing both short ($\pi^e_{s,t}$) and long-term inflation expectations ($\pi^e_{l,t}$) to study the degree of anchoring for the US. Connecting their work to news-regressions, they identify a so-called macro news shock which is supposed to capture all new information about short-run macroeconomic developments.

Their SVAR model is specified as follows:

$$\begin{pmatrix} \pi^e_{s,t} \\ \Delta\pi^e_{l,t} \end{pmatrix} = c + \Pi(L) \begin{pmatrix} \pi^e_{s,t-1} \\ \Delta\pi^e_{l,t-1} \end{pmatrix} + B \begin{pmatrix} \epsilon^{news}_t \\ \epsilon^{target}_t \end{pmatrix}, \tag{3.2}$$

where c, B, and $\Pi(L)$ are coefficient matrices and the stationary series of short-term expectations enters the VAR in its level, whereas the long-term expectations series, which is found to be integrated of order one, is first differenced. ϵ^{news}_t is the macro news shock, while the second shocks ϵ^{target}_t is supposed to capture (perceived) changes in the monetary policy strategy, including the central bank's inflation target. A significant response of long-term expectations to the second shock, indicating that consumers or market participants adjust their perception of the target, is interpreted as still anchored expectations, albeit they are then anchored at a different level.

Identification of the shocks is achieved through a long-run restriction of the Blanchard and Quah (1989) type: it is assumed that macro news shocks have no persistent impact on the level of long-term inflation expectations. As this is equivalent to assuming that long-term expectations are anchored in the long run, the authors test the validity of the restriction with a statistical identification approach. In particular, they use the heteroskedasticity in the data to identify structural shocks. Given the statistically identified shocks, the long-run restriction cannot be rejected. Then, the authors assess the anchoring of expectation in the short run by measuring the relative importance of the macro news shocks for the variance and historical development of long-term expectations.

It is not obvious, however, that the statistically identified shocks indeed capture news about macroeconomic developments. In particular, the approach based on heteroskedasticity relies on strong assumptions: the variance of the structural shocks is assumed to significantly change over time while their impact on inflation expectations, on the other hand, is assumed to stay constant. Should the shocks not (only) capture macro news, then a long-run anchoring would be imposed without properly testing for it.

3.2.3 Using MNAs as proxy variables in a VAR

This paper builds on the bi-variate VAR specification of Nautz et al. (2016). Instead of imposing a long-run restriction and testing it with a statistical approach, however, I propose a different identification strategy for the macro news shock. I use a set of MNAs as proxy variables for the unobservable shock and apply the approach developed by Stock and Watson (2012), Olea et al. (2013), and Mertens and Ravn (2013), that allows to identify structural shocks based on available proxies.[4] In doing so, I make use of the information content of MNAs and assure that the identified

[4]In the literature, the proxy variables are also often referred to as external instruments (compare also Chapter 4). Henceforth, both expressions are used interchangeably.

shocks will be correlated with observable macroeconomic surprises by construction, aligning the SVAR approach more closely to the news-regression literature.

In principle, a dynamic response of expectations to MNAs can also be obtained by including them as exogenous variables in a VAR model. However, using the surprises as proxy variables to identify a shock has two important advantages that also address two shortcomings of news-regressions. First, the approach yields an estimated series of macro news shocks that has entries also on days on which no data releases are available through MNAs, accounting for the fact that news constantly arrive on the market. Second, the approach is robust to measurement error in the proxy variable (see Mertens and Ravn, 2013), alleviating concerns about potential measurement errors in MNAs.

Different to Nautz et al. (2016), who work with monthly survey data of inflation expectations, I use daily financial market based data. This is for two reasons. First, for the identification approach to work, the correlation between the reduced-form residuals from the VAR model and the proxy variables has to be sufficiently high. This requirement holds for daily financial market data, where a significant reaction to some MNAs on announcement days is observable (see Section 3.4.3), but is violated for monthly survey data. Second, for a sufficient number of observations with monthly data the sample of Nautz et al. (2016) covers a period from 1991 to 2015, for which they have to assume constant impact of macro news shocks. News-regressions, however, find that the degree of anchoring in the US has changed after the financial crisis, with a particular de-anchoring afterwards (see Galati et al., 2011, Autrup and Grothe, 2014, or Nautz and Strohsal, 2015). Using daily data allows me to focus on this particular pre-crisis period and avoids imposing a constant impact of macro news shocks over a prolonged period.

3.3 A SVAR model to identify macro news shocks with MNAs

3.3.1 Reduced-form model

The reduced-form VAR model that I employ can be written as:

$$\begin{pmatrix} \pi^e_{s,t} \\ \pi^e_{l,t} \end{pmatrix} = c + \Pi(L) \begin{pmatrix} \pi^e_{s,t-1} \\ \pi^e_{l,t-1} \end{pmatrix} + \begin{pmatrix} u_{s,t} \\ u_{l,t} \end{pmatrix}, \tag{3.3}$$

where the 2×1 vector c includes constant terms, the matrix $\Pi(L)$ in lag polynomials captures the autoregressive part of the model, and $u_{s,t}$ and $u_{l,t}$ are serially uncorrelated innovations, or reduced-form shocks, with covariance matrix Σ_u. In contrast to Nautz et al. (2016), I include both measures of expectations, $\pi^e_{s,t}$ and $\pi^e_{l,t}$, in levels. Whereas their long-run restriction à la Blanchard and Quah (1989) is based on short-term expectations being stationary and long-term expectations being integrated of order one, and included in first differences, the proxy variables approach does not require such a setup. Unit root tests on the daily inflation expectations series that I employ (see description below) mostly point towards long-term expectations being integrated of order one, while results are mixed for short-term expectations. Thus, both underlying structural shocks in my model could, in principle, have a permanent impact on at least long-term expectations.[5]

3.3.2 Structural model and shock identification

The reduced-form VAR innovations and, thus, inflation expectations are assumed to be driven by two uncorrelated structural shocks. The first shock, which I aim

[5] Note that, if short-term expectations are indeed integrated of order one, also a co-integration relationship between the two series might shape the results. While I do not explicitly test for co-integration, the VAR specification in levels potentially allows for it. On the other hand, if short-term expectations are stationary and long-term expectations integrated of order one, then a specification with first differenced long-term expectations as in Nautz et al. (2016) could also be used. Therefore, I have additionally carried out an estimation in a setup with first differenced long-term expectations (eq. 3.2) and identification of macro news shocks based on the proxy variables. It shows that the permanent impact of the macro news shock on the level of long-term expectations is statistically not differentiable from zero, as indicated by (comparably wide) confidence bands. In this setup, the macro news shock explains about 15 % of the forecast error variance in the change of long-term expectations, comparable to results by Nautz et al. (2016).

to identify, is the macro news shock, ϵ_t^{news}. Its definition is based on Nautz et al. (2016): the shock should contain all new information on short-run macroeconomic developments. Therefore, it will be constructed such that it correlates with observable macroeconomic surprises, namely the MNAs. Given that the shock captures short-run developments, it should have a significant impact on short-term inflation expectations. On the other hand, its impact on long-term expectation and, thus, their (de-)anchoring, is precisely what will be assessed in this paper.

The second shock, ϵ_t^*, can be thought of as capturing shocks that refer to the long-run monetary policy strategy, like the shock to agents' perception of the inflation target identified by Nautz et al. (2016). However, since no proxy variable for such a shock is available, I leave ϵ_t^* unidentified in the SVAR specification. Since I evaluate the anchoring of inflation expectations with respect to macroeconomic news, as in the literature, leaving ϵ_t^* unidentified will not impact the anchoring assessment.

The reduced-form innovations u_t are linearly related to ϵ_t^{news} and ϵ_t^* through

$$u_t = b^{news}\epsilon_t^{news} + b^*\epsilon_t^*. \tag{3.4}$$

The 2×1 vector b^{news} captures the impact impulse vector to a macro news shock of size 1. It is assumed that the structural shocks have a unit variance ($\Sigma_\epsilon = I$), as commonly done in structural VAR models. From this it follows that $\Sigma_u = BB'$ where B is a matrix with $B = (b^{news}, b^*)$. The identification strategy for the macro news shock is based on the proxy variables approach developed by Stock and Watson (2012) and Mertens and Ravn (2013), and its adaption to monetary policy by Gertler and Karadi (2015). With proxy variables m_t available such that

$$E(m_t\epsilon_t^{news}) = \phi \neq 0, \tag{3.5a}$$

$$E(m_t\epsilon_t^*) = 0, \tag{3.5b}$$

where ϕ is a vector of the size of m_t with some non-zero entries, Mertens and Ravn (2013) show how to consistently estimate an impulse vector $\tilde{b}^{news}$, which differs from b^{news} only up to a scalar μ, by exploiting the correlation between m_t and the estimated reduced-form residuals. Estimating $\tilde{b}^{news}$ is already sufficient to compute impulse responses to a macro news shock that affects short-term inflation expectations by a pre-scaled amount on impact.[6] With the additional assumption $\Sigma_\epsilon = I$, a consistent estimate for b^{news} can be computed. This allows studying the dynamic propagation of a one standard deviation macro news shock and obtaining forecast error variance and historical decompositions with respect to the shock.[7]

As proxy variables m_t for the unobservable macro news shock I use MNAs. With m_t at hand, there are different options for how to implement the identification of the VAR model. In this paper, a two stage least squares approach as in Gertler and Karadi (2015) or Cesa-Bianchi et al. (2015) is employed. I use this particular approach as m_t consists of more than one series, as surprise realizations of different macroeconomic variables prove to be a good proxy for the macro news shock. The exact choice of MNAs included in m_t is described in Section 3.4.3.

The approach starts with obtaining an estimate of the reduced-form residuals for short-term and long-term expectations: u_t^{short} and u_t^{long}. Then, in the first stage, u_t^{short} is regressed on the proxy variables m_t:

$$u_t^{short} = \tau m_t + \eta_{1t}, \tag{3.6}$$

to form the fitted value $\hat{u}_t^{short}$, where τ is a vector of the same length as the number of proxies. Intuitively, in this first stage regression the variation in u_t^{short} is captured that is due to the proxy variables and, thus, due to the macro news shocks. The second stage regression is then carried out as follows:

$$u_t^{long} = \gamma \hat{u}_t^{short} + \eta_{2t} \tag{3.7}$$

where $\hat{u}_t^{short}$ is orthogonal to the error term η_{2t} given assumption (3.5b). The estimated coefficient $\hat{\gamma}$ is a consistent estimate of $b_{long}^{news}/b_{short}^{news}$ (compare de Wind, 2014, Gertler and Karadi, 2015, or Cesa-Bianchi et al., 2015).

[6] Note that estimating $\tilde{b}^{news}$ is sufficient to compute the *relative* responses of the variables in the system, b_j^{news}/b_i^{news}, as $\tilde{b}_j^{news}/\tilde{b}_i^{news} = b_j^{news}/b_i^{news}$ since μ is constant. For the bi-variate model of expectations this corresponds to $b_{long}^{news}/b_{short}^{news}$ being computable. Setting the response of short-term expectations to the macro news shock, b_{short}^{news}, to a pre-scaled amount, the corresponding response for long-term expectations, b_{long}^{news}, can be recovered.

[7] de Wind (2014) explains in detail how to exploit the covariance restrictions along with 3.4, 3.5a, and 3.5b to obtain a consistent estimate of b^{news}, and also how to derive the corresponding row of the inverse impact matrix B^{-1} without any additional identification assumptions. Along with the estimated reduced-form residuals, this information is sufficient to obtain impulse response functions to a one standard deviations macro news shock as well as forecast error variance and historical decompositions with respect to the shock.

3.3.3 Anchoring criteria assessed with the SVAR

Estimating the SVAR model and identifying the shock allows evaluating the anchoring of long-term inflation expectations with respect to macroeconomic news. In particular, based on the literature outlined in Sections 3.1 and 3.2, I will assess the following anchoring criteria with the estimated model. First, connecting to news-regressions, long-term expectations are defined to be instantaneously de-anchored if they significantly respond to the macro news shock on impact. This can be evaluated with the estimated impact impulse vector b^{news}. Second, as in other studies with dynamic models, long-term inflation expectations are defined to be anchored in the long run if the macro news shock has no permanent impact on them. Given the approach, this can be assessed by checking whether the estimated response of long-term expectations to the identified shock fades out eventually.

Third, if long-term expectations are anchored in the long run but not on impact, the time it takes for the effects of macro news shocks to fade out is an indication for the degree of de-anchoring in the short(er) run. Fourth, similar to Nautz et al. (2016) the overall degree of de-anchoring will be further evaluated by the share of forecast error variance of long-term expectations explained by the shock. Fifth, the de-anchoring of expectations in specific time period will be analyzed with a counterfactual exercise that studies how expectations would have developed according to the estimated model if no macro news shocks had occurred.

3.4 Data: financial market based expectations and MNAs

3.4.1 Inflation expectations

To measure inflation expectations, I use daily financial market data. This is necessary for the identification approach to work and allows me to focus on the period after the financial crisis, in which a particular de-anchoring of expectations is found in other studies. Specifically, Nautz and Strohsal (2015) perform multiple endogenous breakpoint tests to news-regressions and detect a breakpoint in July 2009, with long-term inflation expectations more significantly responding to MNAs ever since. My baseline sample thus begins in July 2009 and runs until August 2016. By choosing this starting date I also avoid that market turbulences during the financial crisis severely distort the measures of expectation and, thus, bias my results (see below).

For the baseline specification, I use data from the growing market of inflation-linked treasury securities, which has become an important source of information about future inflation. Expectations are measured by break-even inflation rates (BEIR / BEI rates), the difference between yields of nominal and inflation-indexed treasury bonds with the same maturity.[8] As a measure of short-term expectations I employ the three year spot BEI rate. The long-term inflation expectation is the five year five year forward BEI rate, the inflation expectations over five years in five years. Later, I also evaluate the sensitivity of the main results to using the two year or four year spot BEI rate as a short-term and the one year in nine years forward BEI rate as a long-term measure. The two series included in the baseline specification are displayed in Figure 3.4.1. It shows that in particular long-term expectations display a pronounced downward trend after mid 2014 that has brought them to historically low levels. On the other hand, the downward trend is far less pronounced in short-term expectations.

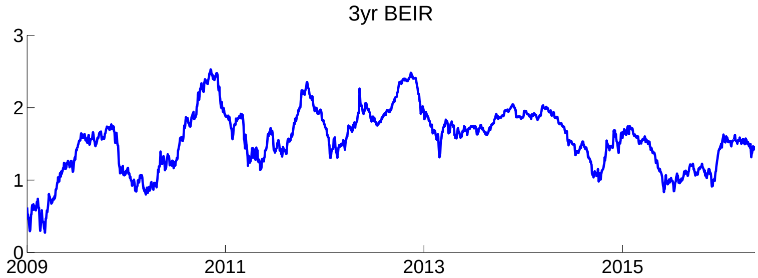
Liquidity premium and inflation risk premium

BEI rates, however, do not only capture expectations about future inflation but potentially also include inflation risk and liquidity premia. While the former premium contains information that should not be disregarded for the anchoring of expectations, the latter might reduce the information content of BEI rates. For instance, liquidity conditions, such as trading volume, might affect the (relative) pricing of nominal and inflation-indexed bonds and thus have an impact on BEI rates. Such liquidity effects were particularly pronounced during the height of the financial crisis in 2008, as flight-to-safety effects in the nominal treasury market strongly compressed the spread between nominal and inflation-indexed bond yields (see D'Amico et al., 2014). As the sample begins in July 2009, my results are not biased by this specific episode. Nevertheless, I also perform different robustness exercises to ensure that results are not solely driven by reactions in the liquidity premium. On the one hand, I carry out additional estimations using data on in-

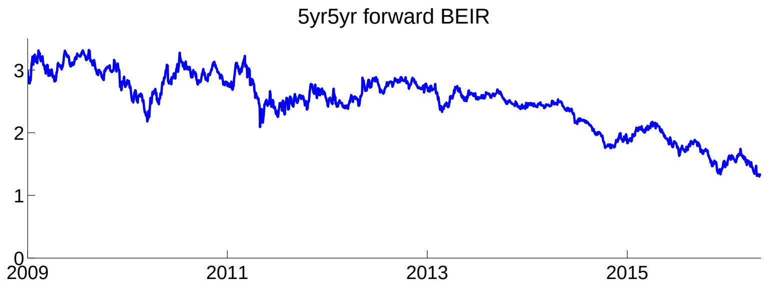
[8] $BEIR_t = yield_t^{nominal} - yield_t^{indexed}$. BEI rates are taken from the database provided by the Federal Reserve Board staff (see also Gürkaynak et al., 2010a). Details on all the data used is given in Table 3.A in Appendix 3.A.

Figure 3.4.1: Inflation expectations for baseline model

a) short-term expectations



b) long-term expectations



Note: The figure shows the inflation expectations measures used in the baseline estimation. The short-term measure is the three year spot break-even inflation rate (3yr BEIR), the long-term measure is the break-even inflation rate over five years in five years (5yr5yr forward BEIR). The sample is 07/01/2009 - 08/26/2016.

flation swaps instead of BEI rates. While inflation swaps potentially also include a liquidity premium, it is not necessarily linked to the one entailed in BEI rates. On the other hand, I employ two different regression based approaches used in the literature to pre-filter the BEI rates for liquidity premia. Results with different pre-filtered data serve as a useful benchmark to compare the baseline results.

3.4.2 MNAs

As proxy variables m_t for the unobservable macro news shock I use MNAs as utilized in news-regression frameworks. Specifically, I collect 33 MNAs from Bloomberg, covering the lion's share of available MNAs and including all important ones.[9] The

[9]Using MNAs from this source is standard in the literature as Bloomberg is used by many financial market participants to get an idea of what the market expects a macroeconomic release to be. Importantly, the survey expectations from Bloomberg are unlikely to be biased or stale at most points in time, as forecasts can be adjusted until the very last moment before the release.

complete list can be found in Table 3.A.2 in Appendix 3.A. For each announcement of each included variable, I obtain both the actual data release and the expected realization stemming from a survey of economists. The respective MNA is the difference between the realized number and its expected value. Most of the included series have announced realizations once a month, with a few exceptions that are realized quarterly, like GDP growth, or weekly, like initial jobless claims.

Following the news-regression literature, the raw MNAs are standardized before including them in the analysis (see, for instance, Balduzzi et al., 2001). This is done because the different series are released in different units, and standardizing them allows for comparability. Also, the proxy variables then have the same standard deviation as assumed for the macro news shock. For a specific economic indicator i, the standardized MNA series at time t, MNA_{it}, is computed as follows:

$$MNA_{it} = \frac{R_{it} - E_t(R_{it})}{\hat{\sigma}_i}, \tag{3.8}$$

where R_{it} is the actual realization of the economic indicator, $E_t(R_{it})$ is the expected realization based on the survey of economists, and $\hat{\sigma}_i$ is the sample standard deviation of $R_{it} - E_t(R_{it})$.

In addition to the MNAs available from Bloomberg, I follow Nautz and Strohsal (2015) and construct an additional series capturing (conventional) monetary policy surprises. The calculation of the monetary policy surprise series is done as in their study: I take percentage point changes in the 30-day T-bill rate on monetary policy announcement days.

3.4.3 Choice of MNAs to identify the VAR model

All in all, my dataset of potential proxies thus contains 34 MNAs. Recall that conditions (3.5a) and (3.5b) require proxy variables to be correlated with the shock of interest and uncorrelated with the other structural shocks. While these conditions are not directly testable as the structural shocks are unobserved, one can test whether there is a sufficiently strong correlation between the proxies and the reduced-form VAR innovations. This is a necessary condition for them to be considered a useful tool for analyzing the underlying drivers of the variables. In particular, if the proxies do not have sufficient explanatory power in the first stage, results from the second stage regression will not be informative. Interpreting the proxies as external instruments, as often done in the literature, this would correspond to a weak instrument problem due to non-relevant instruments. One can test the relevance of the proxy variables by adding a constant to equation (3.6) and performing an F-test on their joint significance.

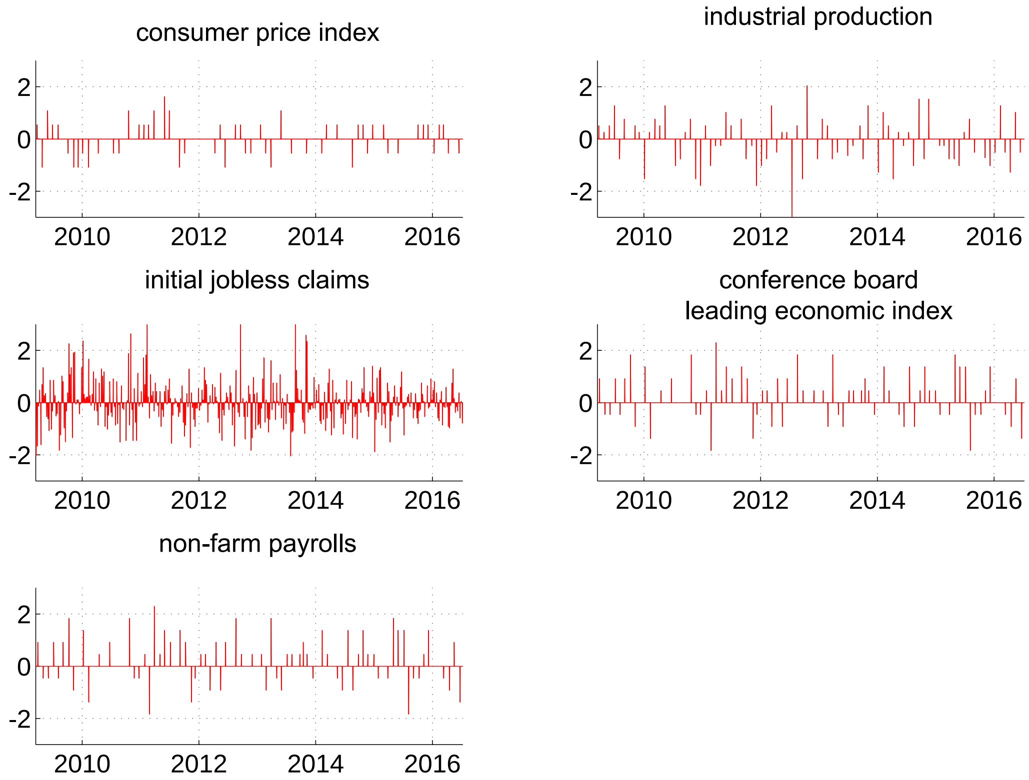
To avoid a weak instrument problem, I aim at maximizing the correlation between the set of proxy variables and ${\hat{u}_t}^{short}$ in the first stage and, thus, the corresponding F-statistic. I start by estimating the reduced-form VAR from (3.3) to obtain ${\hat{u}_t}^{short}$ and ${\hat{u}_t}^{long}$. Lag length selection in the VAR model is based on standard selection criteria, resulting in a lag length of four for the baseline specification.[10] Then, I employ a general to specific approach based on the total of 34 MNAs. Of the 34 series, I only keep the ones as part of the proxy variables set that are statistically significant in the first stage regression on the 5 % percent level. In other words, I only include the MNAs that have a significant impact on short-term inflation expectations. In the baseline specification, the set of proxies contains surprise realizations of the consumer price index, industrial production, initial jobless claims, the conference board leading economic index, and non-farm payrolls. The included standardized MNAs are displayed in Figure 3.4.2. While MNAs for initial jobless claims are available weekly, the other four variables are available on a monthly frequency. The resulting F-statistic in the first stage is 11.98. This is well above the value of ten usually required for relevant instruments in the identification approach (compare also Stock et al., 2002) and indicates that a weak instrument problem is not present.[11]

[10]In fact, the AIC criterion and the final prediction error suggest a lag length of four, while the BIC suggest a lag length of two. According to correlograms and Portmanteau tests, however, a lag length of four is necessary to obtain residuals free from autocorrelation.

[11]For the monthly dataset of Nautz et al. (2016) based on survey data, in contrast, I do not find that the correlation between MNAs and reduced-form residuals is sufficiently high for the identification approach to work, presumably due to the lower data frequency and survey participants adjusting their expectation to more than just observable MNAs within a month. This reflects both strengths and weaknesses of the proxy variable identification approach. While it guarantees a high correlation between the MNAs and the estimated structural macroeconomic news shocks, it can only be applied if the proxy variables have explanatory power.

I find a similar lack of explanatory power when trying to use the approach with daily financial market based expectations for the euro area to compare US results to. While no break-even inflation rates based on bonds for the euro area as a whole are available, inflation swap rates as used in the robustness exercise can be obtained from Bloomberg. Using the swap rates and MNAs for 130 European variables, both euro area aggregates and individual series for the four largest member states, I run a similar SVAR exercise as for the US. However, it shows that the

Figure 3.4.2: Set of MNAs used as proxy variables in the baseline model



Note: The figure shows the set of proxy variables for the baseline specification. It contains MNAs for the consumer price index, industrial production, initial jobless claims, the conference board leading economic index, and non-farm payrolls, all from Bloomberg. The series have been standardized prior to the estimation to make them comparable.

Results from the first stage identification regression are given in Table 3.4.1. It shows that in particular MNAs for consumer prices, initial jobless claims, and non-farm payrolls have a highly significant impact on short-term expectations. The

correlation between the reduced-form residuals and the MNAs is considerably lower than in the US model. In fact, given a similar set of intuitive MNAs as for the US, the F-statistic for the first stage regression is still indicative of a weak instrument problem. Different explanations are conceivable for the low correlation. On the one hand, the inflation swap data for the euro area seem to contain more noise than their US counterparts. On the other hand, this could also be interpreted as a better anchoring of expectations with respect to macroeconomic news. Financial market based inflation expectation in the euro area, however, feature an even more pronounced decline than the one displayed for the US in Figure 3.4.1. Hence, it is difficult to conclude that inflation expectations in the euro area are better anchored in general, instead the proxy variable identification approach does not appear to be suitable for the particular dataset.

estimated coefficients have intuitive signs. For instance, a one standard deviation positive surprise in inflation, i.e. a higher than expected realization, increases short-term inflation expectations on impact by about 4 basis points. Similarly, a positive surprise in non-farm payrolls is associated with an increase in short-term expectations, whereas a higher than expected number of initial jobless claims is correlated with lower expectations. MNAs for industrial production and the conference board leading economic index, on the other hand, are found to be less significant, but still slightly above the 5 % significance level cutoff for the baseline specification, and display counterintuitive signs.[12] The R-squared of 3 % is about the same as usually found in news-regressions. It increases to more than 7 % if only observations are taken into account for which MNAs are available for the five included variables. Having identified the VAR with the proxy variables, I next turn to the empirical results and evaluate the anchoring criterion outlined in Section 3.3.3.

Table 3.4.1: Results from first stage identification regression

First stage equation: $u_t^{short} = \tau_i MNA_{it} + \eta_{1t}$					
			τ_i		
	cons. price index	industrial production	init. jobless claims	conf. board lead. index	non-farm payrolls
u_t^{short}	0.038***	-0.010**	-0.005***	-0.012**	0.017***
	(5.61)	(-2.13)	(-2.77)	(-2.14)	(3.54)
F-stat: 11.98	Observations: 1833	R^2: 0.03			

Note: The table shows results from the first stage identification regression, where the reduced-form residuals of short-term inflation expectations are regressed on the set of proxy variables. ***, **, * denote significance at the 1%, 5%, 10% levels. T-statistics are given below estimates.

[12] The main empirical results are robust to dropping the two MNAs from the set of proxy variables.

3.5 Empirical results

3.5.1 Response of inflation expectations to macro news shocks

Figure 3.5.1 displays the estimated impulse response functions to a one standard deviation macro news shock from the baseline SVAR model, along with bootstrapped 90 % confidence bands.[13] The responses show that the shock has a significant effect on short-term inflation expectations, as already indicated by the first stage identification regression. On impact, they increase by about four basis points and are significantly raised for about 200 trading days. To assess the outlined anchoring criteria, the response of long-term expectations is important. Following the macro news shocks, long-term expectations significantly increase on impact by about 1.5 basis points. The response stays significantly positive for about 50 trading days. In the long run, however, the response fades out and expectations return to their pre-shock level eventually. Taken together, the response thus indicates a de-anchoring of expectations in the short run, but not in the long run. This is in line with the findings by Nautz et al. (2016). A similar persistence in the reaction of long-term inflation expectations is also found by Strohsal and Winkelmann (2015) with respect to reduced-form shocks in a univariate ESTAR model and by Netsunajev and Winkelmann (2016) in a cross-country SVAR model for spillovers.

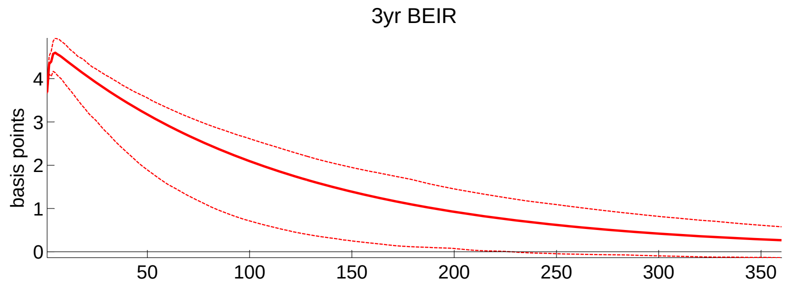
3.5.2 Forecast error variance decomposition

Given that inflation expectations react significantly to the macro news shock in the short run, I continue by investigating the relative importance of the shock for the variance of expectations using a forecast error variance decomposition. The percentage contribution of the macro news shock to the forecast error variance of short-term and long-term expectations is displayed in Table 3.5.1. It shows that macro news shocks explain almost all the variance in short-term expectations. This is different for long-term expectations. The numbers, nevertheless, indicate that the influence of macro news shocks on the forecast error variance of long-term inflation

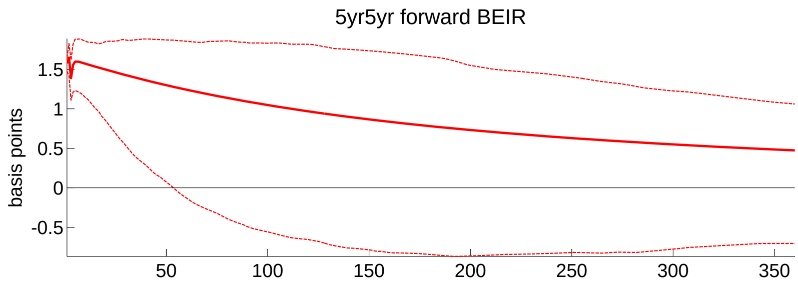
[13] As in Mertens and Ravn (2013) and Gertler and Karadi (2015), for statistical interference a fixed-design wild bootstrap procedure is applied that accounts for estimation uncertainty in both stages of the structural VAR estimation: reduced-form estimation and identification regressions. This is achieved by re-sampling both VAR data and proxy variables.

Figure 3.5.1: Responses of inflation expectations to the macro news shock

a) short-term expectations



b) long-term expectations



Note: The figure shows the estimated impulse responses, along with their 90 percent confidence bands, obtained using 500 bootstrap replications, of short-term and long-term inflation expectations to a one standard deviation macro news shock. The sample is 07/01/2009 - 08/26/2016. Short-term inflation expectation: three year break-even rate, long-term inflation expectation: five year five year forward break-even rate.

expectations is non-negligible. In the short run, the fraction of variance explained by the shock is 16 %, while over time this declines to about 10 %. In line with the anchoring criteria, this indicates that even though expectations are anchored in the long run, they are still considerably distorted by the shocks.[14]

Also note that the percentage contributions to long-term expectations appear to be higher than what Nautz et al. (2016) on average find in their study with survey data, in particular compared to the (dominating) regime where the variance of macro news shocks is estimated to be low. However, the results are not directly comparable

[14]Note that the forecast error variance decomposition remains valid for a limited finite horizon despite at least long-term inflation expectations presumably being integrated of order one. However, in integrated systems the underlying prediction mean squared error diverges for increasing forecast horizons (compare Kilian and Lütkepohl, 2017). Therefore, I do not report longer horizons than shown above or even the limiting variance decomposition.

Table 3.5.1: Percentage contribution of macro news shocks to variance of inflation expectations

	Horizon in days							
variable	1	5	10	20	30	90	150	360
π_s^e	99	99	99	99	99	99	99	99
π_l^e	16	16	16	16	16	14	13	11

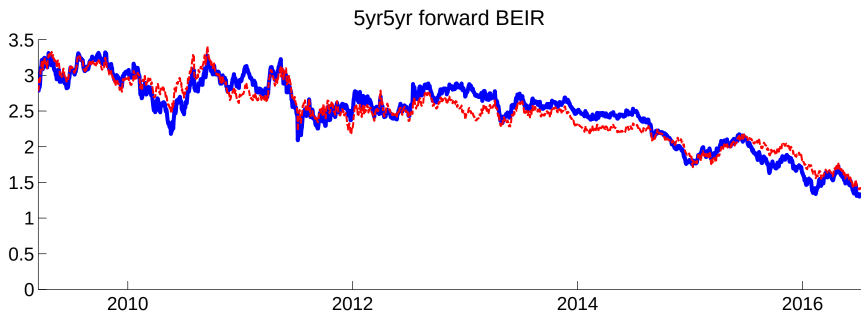
Note: The table shows forecast error variance decompositions of inflation expectations from the baseline SVAR model with respect to the macroeconomic news shock. π_s^e: short-term expectations, π_l^e: long-term expectations.

as their decomposition refers to the change in long-term expectations, and not the level, as they use a different setup of the model (see discussion in Section 3.3). In the sensitivity analysis, I further investigate how the choice of the identification approach and of the sample shapes my results regarding the share of variance explained. I find that using a long-run restriction setup, as Nautz et al. (2016) do, instead of the proxy variables approach, leads to macro news shocks explaining about 5 % of the forecast error variance of changes in long-term expectations in the financial market data (see Table 3.C.1 in Appendix 3.C). On the other hand, I also repeat the SVAR estimation with the proxy variables based on only a pre-crisis sample and find that the contribution of macro news shocks to the forecast error variance of long-term expectations then considerably declines: it is about half as large as in the post-crisis baseline sample (see Table 3.C.2 in Appendix 3.C).

3.5.3 Counterfactual analysis: can macro news shocks explain the secular movement in inflation expectations?

While the forecast error variance decomposition in Section 3.5.2 measures the average degree to which macro news shocks influence long-term inflation expectations, it is also of interest how they affected expectations at particular time periods. For instance, to what extent did the shocks contribute to the decline of long-term expectations that started in mid-2014? Or, to put it differently: can the decline be explained by a series of detrimental macro news shocks? To answer this question, I carry out a counterfactual exercise based on the estimated SVAR model in which all macro news shocks are artificially set to zero from the beginning of the sample.[15] The resulting series of long-term expectations, i.e. the estimated level of expectations without macro news shocks, is displayed in Figure 3.5.2. It shows that without the impact of the shocks, long-term inflation expectations would have been somewhat lower at several periods between 2012 and mid-2014, while they would have been a bit higher in the second half of 2015. Overall, however, the observed decline in long-term expectations cannot be attributed to macroeconomic news.

Figure 3.5.2: Actual long-term inflation expectations data (blue thick line) and counterfactual series with macro news shocks set to zero from beginning of the sample (red thin line)



Note: The counterfactural long-term inflation expectations series is based on a counterfactual exercise with the estimated model, in which all macro news shocks are set to zero from the beginning of the sample onwards.

In turn, this implies that the other shock in the SVAR model is responsible for inflation expectations to fall to about 1.5 % in mid-2016. For future assessments of expectations anchoring, this result might have implications. Following the literature, the anchoring criteria evaluated in this study all relate to the impact of macroeconomic news. However, the counterfactual analysis shows that, if the only potential source of de-anchoring were macro news, then the - according to the predominant definition - anchored level of expectations would have been at 1.5 % in

[15] This approach does not correspond to a standard historical decomposition, as these are designed for stationary data. Instead, results are based on an exercise in which a conditional counterfactual series for long-term expectations is computed for which all realizations of macro news shocks from the beginning of the sample are replaced by zero. Note that such an approach has the caveats, that results are very sensitive to the exact starting point of the counterfactual exercise and that the cumulative effect of the turned-off shock prior to the starting point is ignored (compare Kilian and Lütkepohl, 2017).

mid-2016. This is difficult to align with an implicit or explicit central bank target of about 2 %. To illustrate this further, consider a case where the other shock reflects agents' expectations about the inflation target. Then the observed decline would be driven by agents strongly adjusting their expectations downwards and indicate a severe de-anchoring, without macro news contributing to it.

3.6 Sensitivity and robustness analysis

3.6.1 Alternative BEI rates

Finally, I evaluate the sensitivity of the main results to various modifications in the specification.[16] First, I assess whether results are robust to the specific definition of short-term expectations. For this, the model is re-estimated replacing the three year spot BEI rate by the two and four year spot BEI rate, respectively. The response of long-term expectations to the macro news shock and, thus, results regarding the anchoring of expectations are basically unchanged by this alteration (see Figure 3.B.3 in Appendix 3.B).

Then, I also re-estimate the model using a different measure of long-term inflation expectations. Specifically, I replace the five year five year forward BEI rate from the baseline specification by the one year nine year forward BEI rate, the inflation expectations over one year in nine years. While the five year five year forward is often used in the anchoring literature and central banks closely monitor it, the rate might still entail some business cycle component that potentially influences its reaction to the macro news shock. Therefore, as a sensitivity check, I use a rate that solely covers a period even further into the future. Impulse responses, forecast error variance decompositions, and a counterfactual exercise for the corresponding model can be found in Appendix 3.B and 3.C (see Figure 3.B.4 panel a), Figure 3.B.5 panel a), and Table 3.C.3). The effects of macro news shocks on long-term expectations are found to be very similar to the baseline specification. The response of expectations to the shock is significant on impact and for about 50 trading days, but fades out eventually. Moreover, macro news shocks account for between 10 %

[16] For the individual robustness and sensitivity exercises, the lag length of the model and the exact choice of MNAs has to be slightly adjusted in some specifications, for instance for swap rates, to obtain residuals free from autocorrelation and sufficient explanatory power of the proxy variables.

and 15 % of the forecast error variance, but cannot explain the secular decline after mid-2014 also observed in the one year nine year forward rate.

3.6.2 Inflation swap rates

Next, I assess the sensitivity of the results to generally using break-even inflation rates from treasury bonds as financial market based measures of expectations. As outlined above, break-even rates also include a potentially distorting liquidity premium. Therefore, I re-estimate the bivariate model replacing the BEI rates by expectation measures based on inflation swaps following Bauer (2015).[17] While swap rates also include a liquidity premium, the premium is not necessarily connected to the one in BEI rates. Therefore, using them is a useful exercise to evaluate the sensitivity of the main results. Specifically, I use swap rates with the same maturities as in the baseline model. The results from the model with swap data can be found in Appendix 3.B and 3.C in Figure 3.B.4 panel b), Figure 3.B.5 panel b), and Table 3.C.3. The impulse responses show a reaction of long-term inflation expectations to the macro news shock that is very similar to the one in the baseline specification, but with responses fading out slightly later. Accordingly, the share of forecast error variance explained by the shock is a bit higher (16 % after some trading days). As its counterpart based on treasury bonds, the five year five year future rate based on swaps features a strong decline after mid 2014 that cannot be explained by macro news shocks.

3.6.3 Liquidity adjustment for BEI rates

I also evaluate the sensitivity of the estimates by following an approach that is often used in the literature to deal with the issue of liquidity premia: pre-filtering BEI rates by regressing them on measures of liquidity risk. For this purpose, different measures of liquidity risks are brought forward. I use two approaches based on the relative trading volume on inflation-indexed treasury bond markets (see Gürkaynak

[17]Inflation swaps are financial arrangements in which one party pays the CPI inflation rate on an underlying notional amount while the other party pays a fixed interest rate on the same notional. Specifically, I use data on end-of-day rates of US zero-coupon inflation swaps from Bloomberg. Zero-coupon means that no payments are received until the settlement date of the arrangement. The data is available for maturities from two to ten years, so that I can use the three year swap rate and construct a five year five year forward swap rate, as in the baseline specification.

et al., 2010a, Bauer, 2015, Nautz and Strohsal, 2015) and the implied stock market volatility (VIX) (see Galati et al., 2011, Christensen and Gillan, 2012, Netsunajev and Winkelmann, 2016, Nautz et al., 2017), respectively.[18] The results using the pre-filtered data can be found in Appendix 3.B and 3.C in Figure 3.B.4 panel c) and d), Figure 3.B.5 panel c) and d), and Table 3.C.3. The impulse response of VIX-pre-filtered long-term expectations to the macro news shock is very similar to the one with the unfiltered baseline data. Filtering with market volume, on the other hand, yields a more persistent response that nevertheless also fades out eventually. Forecast error variance decompositions show that the shock explains at most 16 % (volume) and 22 % (VIX) of the variance in long-term expectations, respectively. This is slightly more than in the baseline case, but still very comparable, even though the time profile of the explained share shifts a bit. Historical decompositions confirm the results from the counterfactual exercise for the non-filtered BEI rates: the decline in long-term expectations after mid-2014 is not due to a series of detrimental macro news shocks.

3.6.4 Alternative identification strategy: long-run restriction

Then I assess how the main results differ if the macro news shock is identified with a long-run restriction as in Nautz et al. (2016) instead of the proxy variables approach. For this purpose, I re-estimate the model using the setup of Nautz et al. (2016) from eq. (3.2), i.e. I include the long-term inflation expectations in first differences while the short-term expectations enter in levels. For identification I impose that macro news shocks have no persistent impact on the level of long-term expectations.

The resulting impulse responses of inflation expectations, along with their counterparts from the baseline model, are displayed in Figure 3.B.1 in Appendix 3.B. Note that the response of long-term expectations is cumulated from the response of

[18]Details on the pre-filtering regressions are given in Appendix 3.D. I do not use the pre-filtered BEI rates in the baseline specification for several reasons. First, the filtered rates from the two approaches differ considerably and it is not a priori clear which one to prefer. Second, the pre-filtering regressions are very sensitive to the specification of the sample. Therefore, results from a pre-crisis estimation as carried out in Section 3.6.5 would have been difficult to compare to the baseline model. Similarly, also a robustness check with inflation swap data would not have a straightforward interpretation. Third, pre-filtering the data with a regression based approach potentially induces a generated regressor problem that is difficult to correct for in the context of the SVAR model.

the first-differenced variable in the model. The macro news shock identified through the long-run restriction is found to have a somewhat smaller effect on long-term expectation. On impact, it raises them by less than one basis point, about half the effect that is found in the baseline specification, and the response stays significantly below its proxy variables SVAR counterpart for some trading days. This difference appears to be also reflected in the corresponding forecast error variance decompositions. Identified through the long-run restriction, macro news shocks explain about 5 % of the variance in the change of long-term expectations (see Table 3.C.1 in Appendix 3.C), whereas they explain more than 10 % for the level of expectations in the proxy variables SVAR. However, it has to be kept in mind that due to the different specifications, the numbers are not directly comparable.

3.6.5 Inflation anchoring before and after the financial crisis

Lastly, I investigate the role of sample selection in shaping the main results. Recall that the baseline sample comprises only the post-crisis period as news-regressions find a different degree of de-anchoring with respect to MNAs after the financial crisis compared to the period before. I carry out an additional estimation of the baseline model with only pre-crisis data to assess how this affects my main results with the proxy variables SVAR approach. In particular, I use a sample that starts in 2004, once the inflation-indexed bond markets were out of infancy (see, for example, Bauer, 2015), and ends with the bankruptcy of Lehman brothers in September 2008.[19] The corresponding impulse responses to the macro news shock are shown in Figure 3.B.2 in Appendix 3.B. Short-term expectations display a very similar reaction to the shock as in the baseline sample. For long-term expectations, however, the impact effect of the macro news shock is only about half as large and fades out in about half the time. The difference in inflation anchoring between sample periods is also reflected in the forecast error variance decompositions. The corresponding decomposition for the pre-crisis sample (see Table 3.C.2 in Appendix 3.C) shows that macro news shocks contribute only half as much to the variance of long-term expectations as

[19] The immediate aftermath of the Lehman brothers bankruptcy is thus excluded from both baseline and pre-crisis sample. This is necessary as the associated financial market turmoil severely distorted the information content of market based inflation expectations, compare Section 3.4.1.

found for the baseline. Thus, also the dynamic SVAR approach indicates a stronger de-anchoring of inflation expectations with respect to macro news after the crisis.

3.7 Conclusion

Well-anchored long-term inflation expectations should not respond to news about short-run macroeconomic developments. Empirically, this is often assessed in news-regressions by measuring the impact response of expectations to macroeconomic news announcements (MNAs) on announcement days. Recent contributions take a more dynamic perspective and study the reaction of expectations to macroeconomic news over time.

In this paper, I provide new evidence on the anchoring of US inflation expectations in a dynamic framework. Specifically, I build on Nautz et al. (2016) and analyze the effects of macro news shocks in a bivariate SVAR model containing short-term and long-term expectations. In contrast to their study, however, I do not impose the restriction that macro news shocks have no permanent impact on long-term expectations to identify the model. Instead, I connect the SVAR approach more closely to news-regressions and use MNAs as proxy variables for identification, ensuring that the structural shocks will be correlated with observable information on surprise realization of macroeconomic variables by construction. To measure inflation expectations, I use daily financial market data based on inflation-indexed treasury bonds, as this implies a sufficiently high correlation between the proxy variables and the reduced-form VAR residuals required for the identification approach to work.

I find long-term inflation expectations to be de-anchored in the short run as they react significantly to macro news shocks for about 50 trading days. Eventually, the response to the shock fades out, indicating an anchoring of expectations in the long run. A forecast error variance decomposition implies that macro news shocks are a non-negligible source of distortion as they account for about 10 % of the variance in long-term expectation. These results are in line with the findings by Nautz et al. (2016), showing that the latter are robust to applying an identification strategy based on observable macroeconomic surprises, namely the MNAs. A sensitivity analysis, on the other hand, suggests that solely imposing a long-run restriction would indicate a lower degree of de-anchoring in the daily financial market dataset that I use. Also, I show that the dynamic response of expectations to macro news shocks is stronger after the financial crisis compared to the period before, in line with similar results from the news-regressions literature.

My results also cast doubt on whether future research on inflation expectations anchoring should continue being based on the notion that expectation are fully anchored as long as they are insensitive to macroeconomic news. A counterfactual exercise shows that the decline of long-term inflation expectations in financial market data observed after 2014 cannot be attributed to macro news shocks. Yet, a - by definition - anchored level of long-term inflation expectations of less than 1.5 %, as observed in 2016, is difficult to align with a high degree of confidence in the ability of the central bank to keep inflation close to a 2 % target. Possible extensions of the SVAR approach with proxy variables might thus aim at identifying a broader range of structural shocks that are potentially important for the anchoring of expectations. Potential candidates are shocks to the perceived central bank target (see Nautz et al., 2016, Arias et al., 2016), shocks to the inflation risk premium and inflation uncertainty (see Hördahl and Tristani, 2012, 2014, Nagel, 2016) or shocks to expectations about commodity prices (see Coibion and Gorodnichenko, 2015).

Appendix

3.A Data and sources

Table 3.A.1: Data construction and sources

Variable	Construction and source
Break-even inflation rates	Source: Federal Reserve Board. Initially constructed by Gürkaynak et al. (2010a). Link to WP with latest data update: https://www.federalreserve.gov/econresdata/feds/2008/index.htm.
MNAs	Source: Bloomberg. 33 series, see Table 3.A.2
Monetary policy surprise	Percentage point changes in the 30-day T-bill rate at days of Federal Reserve Board Meetings. Source: Bloomberg, Nautz and Strohsal (2015), own calculations.
Inflation swap rates	End-of-day rates on US zero-coupon inflation swaps. Source: Bloomberg.
Relative trading volume of inflation-indexed bonds	Trading volume of inflation-indexed treasury bonds divided by trading volume of nominal treasury bonds. Source: Federal Reserve Bank of New York. The trading volume is published about once a week. I set all days with no observation equal to the last observed relative volume.
VIX	Option-implied volatility index. Source: Chicago Board Options Exchange, retrieved from St. Louis FRED.
Corporate bond spread	Spread between AAA-rated corporate bond yields and nominal government bond yields. Source: Bloomberg, own calculations.

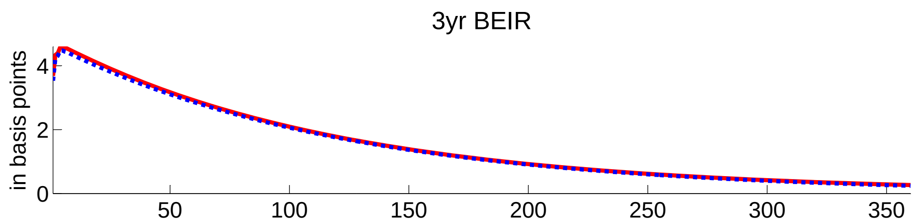
Table 3.A.2: List of included economic data releases (MNAs)

Bloomberg
CPI Urban Consumers YoY NSA
CPI Urban Consumers Less Food&Energy YoY NSA
Personal Consumption Expenditure CPI YoY SA
CPI Urban Consumers MoM SA
University of Michigan Consumer Confidence Indicator
US Government Budget Balance (FED)
American Consumer Spending Growth Rates MoM SA
Markit Manufacturing PMI SA
Markit Composite PMI SA
Markit Services PMI Business Activity SA
Industrial Production MoM 2007=100 SA
Trade Balance of Goods and Services SA
Core Producer Price Index
Producer Price Index - Finished Goods
Initial Jobless Claims SA
Housing Starts/Permits
Difference Between Exports and Imports
GDP Chained 2009 Dollars QoQ SAAR
PPI Finished Goods SA MoM%
PPI Final Demand MoM SA
Capacity Utilization % of Total Capacity
Business Inventories MoM SA
Avg. Hourly Earnings YoY% SA
Avg. Hourly Earnings MoM% SA
Construction Spending Total SA
Construction Spending Total MoM SA
Durable Goods New Orders Industries MoM SA
Conference Board Leading Indicators MoM
Productivity Outpout Per Hour Nonfarm Business Sector QoQ SA
Unit Labor Costs Nonfarm Business Sector QoQ % SAAR
Retail Sales (Less Auto and Gas Stations) SA MoM % Change
Personal Income MoM SA
Nonfarm Payrolls Total MoM SA
Self-contructed (see above)
30-day Treasury-bill rate

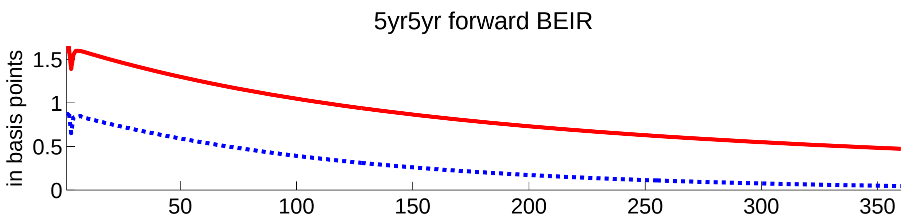
3.B Additional Figures

Figure 3.B.1: Responses of inflation expectations to the macro news shock from baseline model (red line, dashed confidence bands) and from model identified through long-run restriction (blue dotted line)

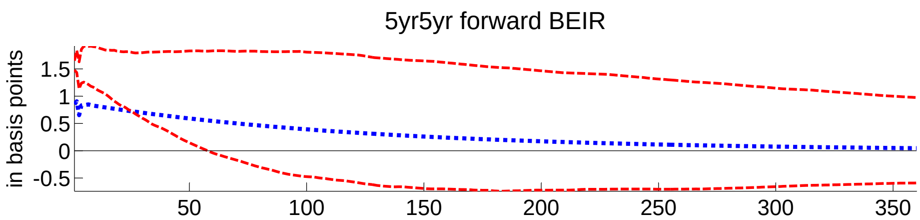
a) short-term expectations



b) long-term expectations



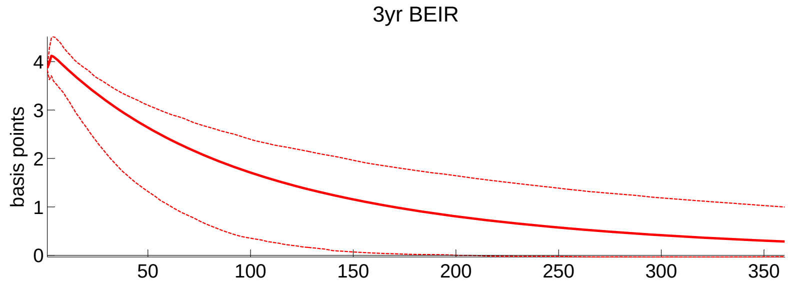
c) long-term expectations - confidence bands



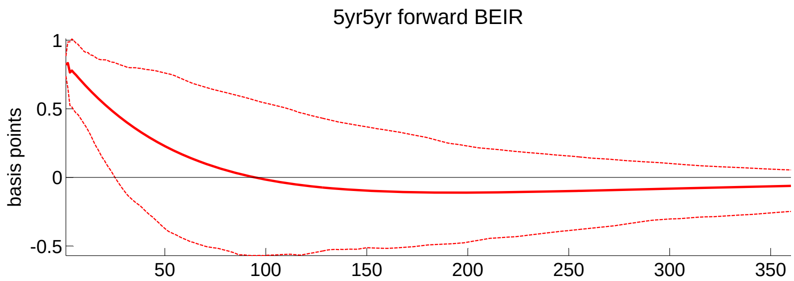
Note: The figure shows the estimated impulse responses of inflation expectations to a one standard deviation macro news shock from the baseline model (red line) and from a model where the shock is identified using a long-run restriction (blue dotted line). The last panel contains the confidence bands, obtained using 500 bootstrap replications, for the baseline response of long-term expectations (red dashed line) along with the mean response from the model identified using a long-run restriction. The sample is 07/01/2009 - 08/26/2016.

Figure 3.B.2: Responses of inflation expectations to the macro news shock - pre-crisis period

a) short-term expectations

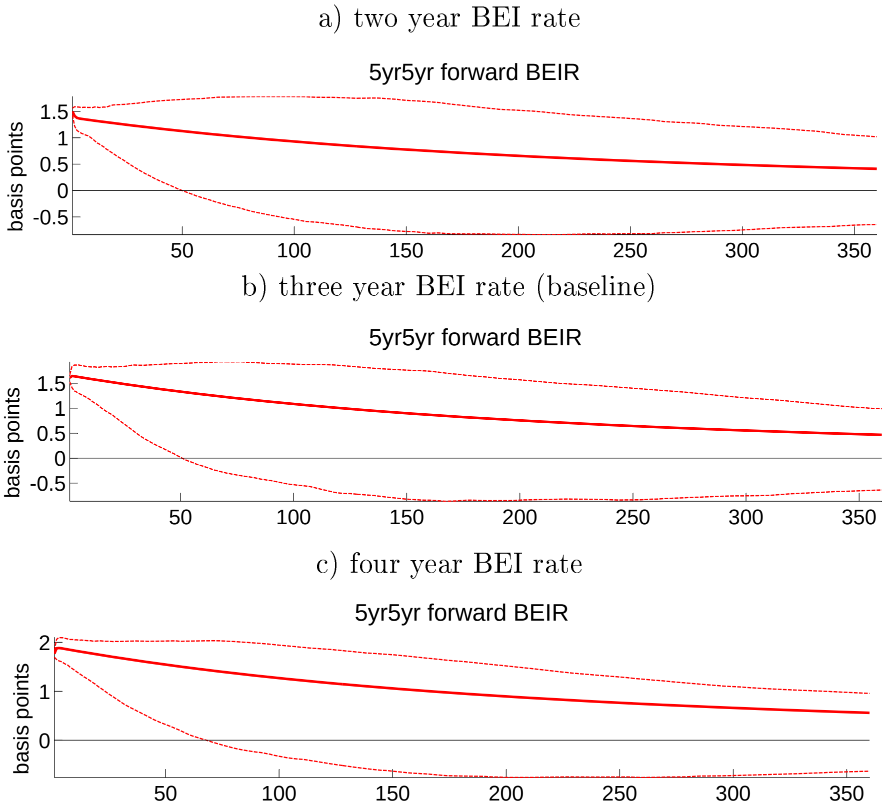


b) long-term expectations



Note: The figure shows the estimated impulse responses, along with their 90 percent confidence bands, obtained using 500 bootstrap replications, of short-term and long-term inflation expectations to a one standard deviation macro news shock. The sample is 01/01/2004 - 08/31/2008. Short-term inflation expectation: three year break-even inflation rate, long-term inflation expectation: five year five year forward break-even rate.

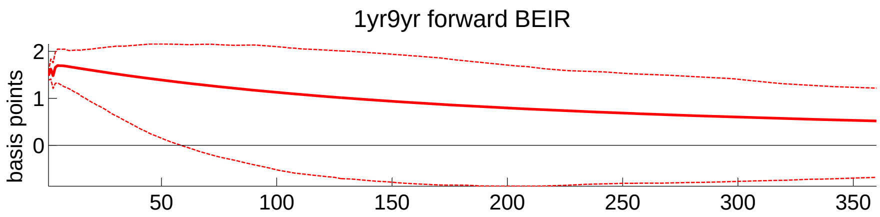
Figure 3.B.3: Responses of long-term expectations from models with different short-term BEI rates



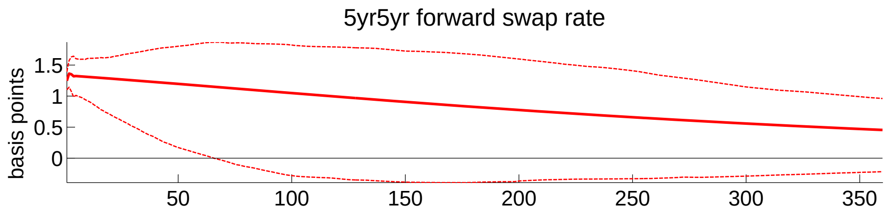
Note: The figure shows the estimated impulse responses, along with their 90 percent confidence bands, obtained using 500 bootstrap replications, of long-term inflation expectations to a one standard deviation macro news shock. The sample is 07/01/2009 - 08/26/2016. The different responses come from models with different measures of short-term inflation expectation. Top: two year break-even inflation rate, middle: baseline, bottom: four year break-even inflation rate.

Figure 3.B.4: Responses of long-term expectations from models with different expectations measures

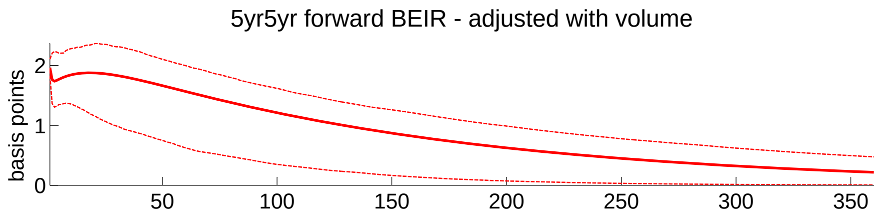
a) different forward rate for long-term expectations



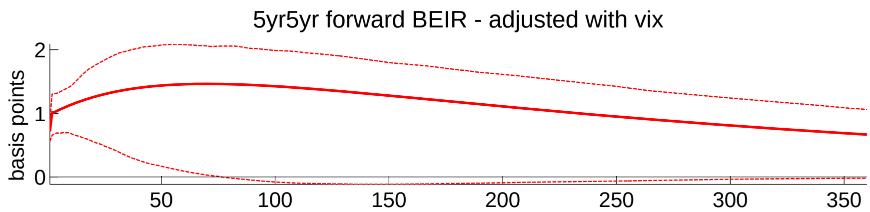
b) inflation swap rates



c) BEI rates adjusted with trading volume



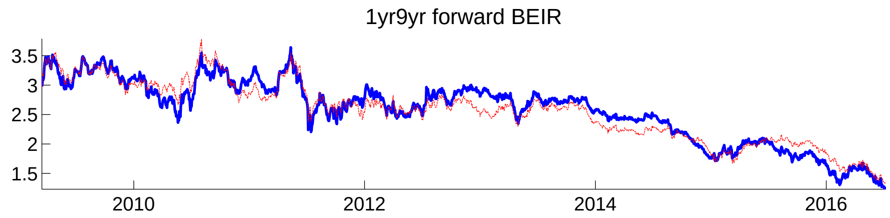
d) BEI rates adjusted with implied volatility



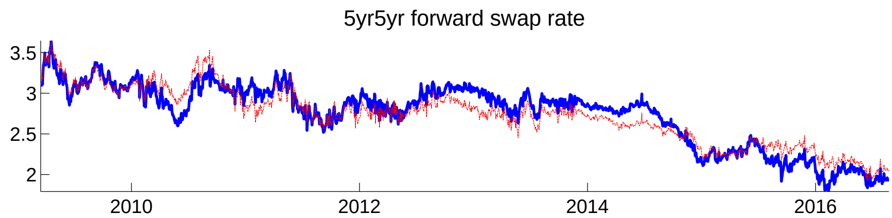
Note: The figure shows the estimated impulse responses, along with their 90 percent confidence bands, obtained using 500 bootstrap replications, of long-term inflation expectations from different model specification to a one standard deviation macro news shock. The sample is 07/01/2009 - 08/26/2016. Top: one year nine year BEI rate as long-term measures, top middle: inflation swaps instead of BEI rates, bottom middle: BEI rates pre-filtered with trading volume, bottom: BEI rates pre-filtered with implied stock market volatility.

Figure 3.B.5: Counterfactual series of long-term expectations from models with different expectations measures

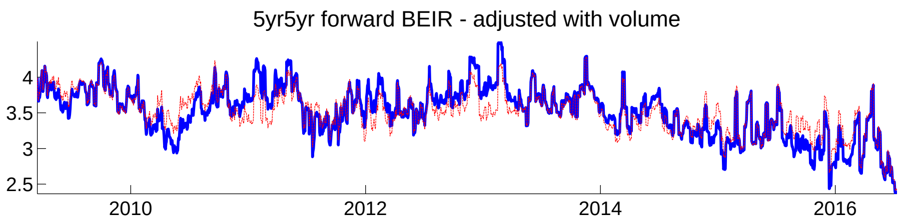
a) different forward rate for long-term expectations



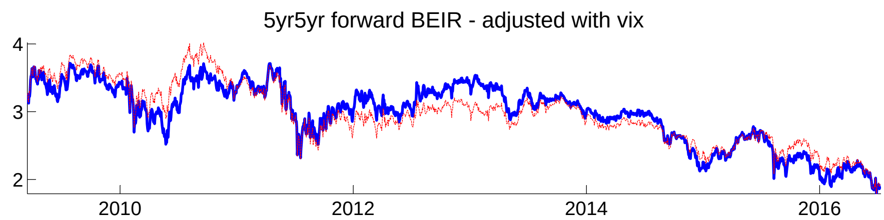
b) inflation swap rates



c) BEI rates adjusted with trading volume



d) BEI rates adjusted with implied volatility



Note: The counterfactual long-term inflation expectations series are based on counterfactual exercises with the estimated models, in which all macro news shocks are set to zero from the beginning of the sample onwards. Actual long-term inflation expectations data (blue thick line) and counterfactual series with macro news shocks set to zero from beginning of the sample onwards (red thin line). Top: one year nine year BEI rate as long-term measures, top middle: inflation swaps instead of BEI rates, bottom middle: BEI rates pre-filtered with trading volume, bottom: BEI rates pre-filtered with implied stock market volatility.

3.C Additional Tables

Table 3.C.1: Percentage contribution of macro news shock to variance of inflation expectations from model identified with long-run restriction

	Horizon in days							
Variable	1	5	10	20	30	90	150	360
π_s^e	91	95	96	96	96	97	97	97
$\Delta\pi_l^e$	5	5	5	5	5	5	5	5

Note: The table shows forecast error variance decompositions of inflation expectations from an SVAR model specification where the macro news shock is identified using a long-run restriction as in Nautz et al. (2016). π_s^e: short-term expectations, $\Delta\pi_l^e$: first difference of long-term expectations.

Table 3.C.2: Percentage contribution of macro news shock to variance of inflation expectations in pre-crisis sample

	Horizon in days							
Variable	1	5	10	20	30	90	150	360
π_s^e	99	99	99	99	99	99	99	99
π_l^e	9	8	8	7	7	5	5	5

Note: The table shows forecast error variance decompositions of inflation expectations from the baseline Proxy SVAR model specification estimated over a pre-crisis sample period (01/01/2004 - 08/31/2008). π_s^e: short-term expectations, π_l^e: long-term expectations.

Table 3.C.3: Percentage contribution of macro news shock to variance of long-tern inflation expectations from models with different expectations measures

		Horizon in days							
Variable	Model	1	5	10	20	30	90	150	360
π_l^e	1yr9yr forward BEIR	12	14	14	15	15	13	12	10
π_l^e	5yr5yr forward swap rate	7	12	13	13	14	15	15	16
π_l^e	5yr5yr forward BEIR (adjusted with volume)	1	1	1	3	5	16	22	22
π_l^e	5yr5yr forward BEIR (adjusted with VIX)	2	3	4	5	6	14	16	16

Note: The table shows forecast error variance decompositions of long-term inflation expectations from different SVAR model specification. Top: one year nine year BEI rate as long-term measures, top middle: inflation swaps instead of BEI rates, bottom middle: BEI rates pre-filtered with trading volume, bottom: BEI rates pre-filtered with implied stock market volatility. π_l^e: long-term expectations.

3.D Pre-filtering the data for the sensitivity analysis

For the sensitivity analysis, I pre-filter the BEI rates from the main specification by regressing them on several measures of liquidity risk. The measures aim at capturing the movements in the liquidity premium entailed in the rates. Specifically, I follow two approaches from the literature to carry out the regressions. In the first approach, I regress the BEI rates on the spread between AAA-rated corporate bond yields and nominal government bond yields and on the relative trading volume on inflation-indexed treasury bond markets as in Gürkaynak et al. (2010a), Bauer (2015), or Nautz and Strohsal (2015).[20] In the second approach, I regress the BEI rates on the spread between AAA-rated corporate bond yields and nominal government bond yields and on the implied stock market volatility (VIX) as suggested by Galati et al. (2011), Christensen and Gillan (2012), Nautz et al. (2017), or Netsunajev and Winkelmann (2016).

Then, I take the residuals from the two regressions to obtain two different measures of pre-filtered, liquidity adjusted BEI rates. The results from the regressions are reported in Table 3.D.1. In both cases, results are broadly in line with the studies mentioned above. However, the two approaches yield estimates of inflation expectations that differ considerably.

[20] The trading volume is not available on daily basis, but published about once a week. Therefore, I set all days with no observation equal to the last observed relative volume.

Table 3.D.1: Pre-filtering regressions for BEI rates

	3yr BEI rate	5yr5yr forward BEI rate
AAA-spread	-0.22	-0.15
	(0.06)	(0.05)
Trading Volume	0.08	-0.37
	(0.01)	(0.01)

	3yr BEI rate	5yr5yr forward BEI rate
AAA-spread	0.13	-0.43
	(0.05)	(0.07)
VIX	-0.04	0.02
	(0.01)	(0.01)

Note: The table shows results of least squares estimations of the three year BEI rate and the five year five year forward BEI rates from the baseline specification on a) the AAA-spread and the Trading volume, and b) the AAA-spread and VIX. AAA-spread: spread between AAA-rated corporate bond yields and nominal government bond yields, Trading volume: relative trading volume on inflation-indexed treasury bond markets, VIX: index of implied stock market volatility. The sample is 07/01/2009 - 08/26/2016.

CHAPTER 4

Unconventional Monetary Policy, Fiscal Side Effects and Euro Area (Im)balances[1]

4.1 Introduction

Following the outbreak of the global financial crisis in 2007, nearly all major central banks engaged in unconventional monetary policy, in the form of credit easing, forward guidance, or asset purchases. The new tools spurred an intense public and academic debate about their benefits and costs. While more and more rounds of easing have been implemented, evidence on the effectiveness of these policies and on how they pass-through to the real economy is still scarce. Even less is known about potential side effects. In this paper, we provide new evidence on the macroeconomic effects of unconventional monetary policy in the euro area.

The public discussion about the benefits and costs of unconventional monetary policy is particularly intense in the euro area. On the one hand, proponents claim that the adopted measures are effective in fulfilling the central bank's mandate of price stability (see Draghi, 2015a, 2016, Constancio, 2016). On the other hand,

[1] This chapter is based on an article that is joint work with Michele Piffer and Malte Rieth. We thank Christiane Baumeister, Refet Gürkaynak, Philipp König, Michele Lenza, Dieter Nautz, Morten Ravn and participants of the Workshop on Empirical Monetary Economics 2016 in Paris, 6th IWH INFER Workshop on (Ending) Unconventional Monetary Policy 2016 in Halle, FMM Annual Conference 2016 in Berlin, Verein für Socialpolitik Annual Meeting 2016 in Augsburg, RCEA Macro-Money-Finance Workshop 2016 in Rimini, the IAAE Annual Conference 2016 in Milan, and an internal seminar for helpful comments and suggestions. We also thank Korbinian Breitrainer for excellent research assistance.

opponents argue that the monetary interventions induce laxer fiscal policy and a widening of euro area imbalances (see Schmidt et al., 2015, Liikanen, 2015, Weidmann and Knot, 2015). They fear that windfall gains from unexpectedly lower interest payments are used to increase primary public expenditures. Moreover, crisis-hit countries could lose competitiveness if prices respond to the common monetary stimulus more strongly there than in other member countries. The public debate develops in parallel to the re-emergence of classic academic questions regarding the effectiveness and fiscal consequences of monetary policy in a currency union (see Orphanides, 2016). How does fiscal react to monetary policy? Are there differences in the responses of the member state economies to a common monetary shock? Does monetary policy accentuate or attenuate internal imbalances within the union?

In this paper, we aim to provide both a quantitative assessment of the arguments made in the debate and some answers to the underlying academic questions. Specifically, we use vector autoregressions to study the effectiveness and side effects of unconventional monetary interventions. Our results show that claims of both parties in the public debate appear to be supported by the data. On the one hand, unconventional monetary policy shocks are effective at lowering public and private interest rates and increasing economic activity, consumer prices, and inflation expectations. On the other hand, the shocks lead to a rise in primary public expenditures, a divergence of consumer prices within the union, and a widening of internal trade balances. Especially the latter findings contribute to the academic debate on unconventional monetary policy, which largely focuses on the effectiveness and transmission of this policy (see below) and less on the side effects.

The identification of causal effects associated with unconventional monetary policy raises new challenges, because identification cannot fully rely on strategies developed for conventional interest rate policies (see Wright, 2012). In this paper, we achieve identification by exploiting daily data on government bond spreads computed against Germany of various euro area countries at different maturities. We extract the common component of changes in spreads around announcements of unconventional monetary policy measures by the European Central Bank (ECB). Building on the 'event study' literature (see Kuttner, 2001, or Gürkaynak et al., 2005a), we view the estimated common yield variations in a narrow window around the policy announcements as a noisy measure of the exogenous components of policy decisions. We then use this measure as an instrument for unobserved unconventional monetary policy shocks in several proxy vector autoregressions (VAR) in order to estimate the average effect of the policy surprises on the macro-economy. In doing so, we follow the methodology developed by Stock and Watson (2012) and Mertens and Ravn (2013) and used by Gertler and Karadi (2015) and by Rogers et al. (2016).

Our results are as follows. Exogenous monetary expansions that lower the average two-year rate on euro area government bonds (excluding Germany) lead to a significant rise in consumer prices and output, and a significant decline of the unemployment rate in the euro area as a whole. Various measures of inflation expectations at different horizons also increase significantly. The monetary policy shocks seem to be transmitted through private and public interest rates, financial market uncertainty and risk aversion, asset prices, as well as credit conditions. All in all, the dynamics of output and consumer prices implied by our model are more similar to the behavior of these variables in empirical models for conventional interest rate policy (see Christiano et al., 1999) than in models for unconventional monetary policy identifying policy innovations as exogenous changes in central banks' balance sheets (see Gambacorta et al., 2014 or Chapter 1). Specifically, relative to balance sheet shocks, the response of output and prices in our model is slower, while the peak effects take place later and are stronger, with output leading prices.

At the same time, our estimates reveal several side effects of the monetary interventions. The fall in sovereign bond yields and public interest payments after the expansion leads to a rise in primary public expenditures. This holds on average for the euro area as a whole as well as for most, but not all, of the largest member states. When looking at individual countries and expenditure categories, the rise in primary expenditures seems to be mainly driven by increases in public consumption. Moreover, as the economies of the member countries are affected differently by the common monetary surprise and since national fiscal authorities respond differently, intra-euro area trade balances widen. In particular crisis-hit countries lose price competitiveness relative to Germany. Their bilateral real exchange rates appreciate, as the increase in domestic demand and prices is more pronounced in these countries.

This paper contributes to a literature on the effects of unconventional monetary policy, which has evolved around two principal approaches. The first approach uses high frequency identification and mainly assesses the contemporaneous effects of these policies on variables available at high frequency, typically financial variables. Among others, Krishnamurthy and Vissing-Jorgensen (2011), Gagnon et al. (2011), Wright (2012), Rogers et al. (2014), and Fratzscher et al. (2016a) find that unconventional policies lower interest rates and term premia and increase asset prices.[2] The second approach uses structural VARs and quantifies the dynamic effects on macroeconomic variables, either on a monthly or on a quarterly frequency using (combinations of) zero and sign restrictions. Ciccarelli et al. (2013), Baumeister and Benati (2013), and Kapetanios et al. (2012) identify monetary policy shocks as exogenous variations in interest rates or spreads. On the other hand, Peersman (2011), Gambacorta et al. (2014), Weale and Wieladek (2016), and Boeckx et al. (2017) isolate unexpected changes in central banks' balance sheets.

In this paper, we use high frequency data for the identification of a VAR model for unconventional monetary policy in the euro area. In doing so, we combine the two approaches discussed above, and complement the analysis of existing VAR studies that are mainly for the US and the UK. We build on Gertler and Karadi (2015), who show how the identification through external instruments allows embedding high frequency financial market data on monetary policy surprises into a structural VAR model for the US economy.[3] In particular, we follow the modification of this approach by Rogers et al. (2016), who use high frequency data by combining estimates of the relative response of variables based on data at different frequencies. The authors mainly analyze the effects of US unconventional monetary policy on exchange rates, dedicating less attention to the euro area. We focus on the euro area in detail and investigate fiscal effects and country heterogeneity.

Our work also builds on Altavilla et al. (2016), who analyze the effects of selected ECB policies by studying the reaction of sovereign yields on days of policy announcements. We follow their approach of measuring the surprise component of monetary policy to construct our external instrument, but extend their framework by proposing a panel setup that extracts the unexpected common variation in spreads of different countries and maturities. We view this extension as important,

[2] Christensen and Rudebusch (2012), Hamilton and Wu (2012), and Wu and Xia (2016) use term structure models to evaluate the impact of unconventional monetary policy on yields and the macro-economy.

[3] Their approach is also used by Cesa-Bianchi et al. (2016) to study unconventional monetary policy for the UK.

given the partially segmented nature of financial markets in the euro area after the financial crisis and the temporary inversion of yield curves. In addition, we use the high frequency estimates for the identification of the VAR models rather than identifying the latter with a recursive structure, as in their paper. Compared to contributions that employ central banks' balance sheets, the identification of the model through information contained in yields has the advantage of capturing the effects of monetary interventions without restricting them to their implementation. This is important because the *announcement* of monetary interventions is a main source of the effectiveness of monetary policy in general (see Blinder et al., 2008) and, in particular, in recent years, when central bank communication in form of forward guidance has become a main policy tool (see den Haan, 2013, Ed.).

Our work also connects to a literature on monetary and fiscal policy interactions (see Fragetta and Kirsanova, 2010, Davig and Leeper, 2011, Traum and Yang, 2011). These papers estimate DSGE models featuring shocks to both monetary and fiscal policy, with the aim of determining policy leadership regimes in the sense of Leeper (1991). More closely related to us methodologically is Rossi and Zubairy (2011). The authors include monetary and fiscal variables jointly in a VAR and recursively identify both fiscal and conventional monetary shocks in the US. Lastly, the paper connects to a discussion on the causes and consequences of euro area current account imbalances (see Blanchard and Giavazzi, 2002, Chen et al., 2013, or Kang and Shambaugh, 2016). It particularly relates to Wyplosz (2013), Comunale and Hessel (2014), and Unger (2017), who stress the role of domestic demand in explaining current account surpluses and deficits.

The paper is structured as follows. In Section 4.2 we discuss the VAR model and the identification strategy. Section 4.3 contains the main results for the euro area as a whole, and then for individual countries. The last section concludes.

4.2 The VAR model

4.2.1 Reduced-form model

The VAR model used can be written as

$$y_t = c + \Pi(L)y_{t-1} + u_t, \tag{4.1}$$

and refers to variables at a monthly frequency. The $k \times 1$ vector c includes constant terms, the matrix $\Pi(L)$ in lag polynomials captures the autoregressive part of the model, and the vector u_t contains k serially uncorrelated innovations, or reduced-form shocks, with $V(u_t) = \Sigma$ and $u_t \sim N(0, \Sigma)$. We use the common lag length selection criteria to choose the number of lags, resulting in a lag length of two. The reduced-form model is estimated on monthly data from 1999M1 to 2015M6. We start the sample with the introduction of the euro to capture the relations between variables in the monetary union. Identification of the unconventional monetary policy shock, however, will rely only on data in the period starting from which unconventional measures were carried out, that is, from 2007M8 to 2015M6.

We employ different specifications for the endogenous variables in y_t. In the baseline specification, y_t includes the six variables discussed below, which refer to euro area aggregates. In the remaining specifications, y_t includes the baseline variables plus one additional variable, which changes across specifications, ranging from measures of inflation expectations to financial variables, fiscal variables and others, both at a euro area level and for single countries. In adding one additional variable at a time, we follow Gertler and Karadi (2015) who use this approach of modifying the marginal variable in a baseline VAR. We follow their approach, which is particularly flexible and does not require a Bayesian perspective, a Panel VAR, or Factor structure to deal with the curse of dimensionality.

The variables included in the baseline specification are

$$y_t = \begin{pmatrix} \text{Two-year rate on euro area government bonds} \\ \text{Stock market volatility} \\ \log(\text{Credit to non-financial firms}) \\ \log(\text{Harmonized index of consumer prices}) \\ \log(\text{Interpolated GDP}) \\ \text{Unemployment rate} \end{pmatrix}.$$

These variables capture financial and interest rate conditions, prices, as well as measures of real economic activity. As a variable reflecting the stance of monetary policy, i.e. a ‘policy indicator’, we use a weighted average of the two-year rates on government bonds for nine euro area countries excluding Germany.[4] In using a (medium-term) government bond rate as policy indicator in a VAR identified with an external instrument, we follow Gertler and Karadi (2015) who employ the one-year US treasury rate. Compared to the short-term interest rates typically used in VAR studies on conventional monetary policy, the variable used has the advantage of taking into account non-standard policy innovations, which are aimed at influencing expectations and yields at longer horizons. Moreover, short-term interest rates like the Eonia or the Euribor would be less suitable because they are constrained by the zero lower bound in our sample. In contrast, the two-year bond rates used is less constrained by the zero lower bound, which it crosses only at the end of the sample.

In addition to the consumer price index (CPI) and the real output variable commonly included in monetary VAR models, we also add a measure of equity market volatility. Specifically, we include the VStoxx volatility index, which is based on option prices of stocks in the EuroStoxx 50. We do so to capture the relation between sovereign bond rates, financial stress, and monetary policy during the identification period, as several non-standard ECB measures were triggered by financial market developments or were specifically aimed at reducing financial risk and uncertainty in the euro area (see Boeckx et al., 2017). Moreover, we add credit to non-financial firms to the model as, for instance, long-term refinancing operations of the ECB, which constitute an important share of the unconventional measures in our sample, have the purpose of stimulating bank lending. Lastly, as a measure of labor market slack and inflation pressure, we include the unemployment rate, since the ECB's large-scale asset purchase programs are specifically targeted at lifting prices and inflation expectations. Appendix 4.A contains details on all variables used in the different specifications and their construction.

The VAR innovations are assumed to be linearly driven by a non-standard monetary policy shock ϵ_t^m, which we aim to identify, and other structural shocks ϵ_t^*, which are of no interest for the purpose of this paper. The VAR innovations u_t are related

[4]The countries are Austria, Belgium, Finland, France, Ireland, Italy, Netherlands, Portugal, and Spain. We exclude German bonds since they played a particular role as a safe haven asset during the euro crisis, whereas for bonds of other euro area countries it is less clear whether they are considered as safe haven assets (see, for example, Altavilla et al., 2016 or Fratzscher et al., 2016a).

to structural shocks ϵ_t^m and ϵ_t^* through

$$u_t = b^m \epsilon_t^m + B^* \epsilon_t^*. \tag{4.2}$$

The $k \times 1$ vector b^m captures the impulse vector to a monetary shock of size 1 and is required to generate impulse responses.

Our identification strategy follows the variant of Rogers et al. (2016) of the identification with external instruments developed by Stock and Watson (2012) and Mertens and Ravn (2013) and its adaption to monetary policy by Gertler and Karadi (2015). Under the condition of a variable m_t being available such that

$$E(m_t \epsilon_t^m) \neq 0, \tag{4.3a}$$

$$E(m_t \epsilon_t^*) = 0, \tag{4.3b}$$

Stock and Watson (2012) and Mertens and Ravn (2013) show how to consistently estimate an impulse vector $\tilde{b}^m$, which differs from b^m only up to a scalar μ, by exploiting the correlation between m_t and the estimated VAR residuals. Estimating $\tilde{b}^m$ is sufficient to compute the *relative* responses of the variables in the system ($\tilde{b}_j^m / \tilde{b}_i^m = b_j^m / b_i^m$) and can be used to obtain impulse responses to a pre-scaled shock to the policy indicator. Building on this methodology, Rogers et al. (2016) use the event-study approach and employ high frequency data in order to refine the estimation of the relative response of the endogenous variables. We first discuss how we compute a measure, m_t (henceforth referred to as instrument or proxy), correlated with the unconventional monetary policy shock for the euro area. We then discuss our identification approach, given m_t, in detail.

4.2.2 A proxy for monetary policy shocks

To construct a proxy m_t for unconventional monetary shocks, we build on Kuttner (2001) and the subsequent literature that uses high frequency data in an event study manner. In general, this approach focuses on one or more selected financial indicators, directly or indirectly associated with the policy rate. It postulates that the price of the indicator closely before a monetary announcement already incorporates the (expected) endogenous response of monetary policy to the state of the economy.

Accordingly, any variation in this price from before to after the announcement reflects an exogenous and unexpected component of monetary policy revealed by the announcement and, consequently, is interpreted as exogenous with respect to the economy (see Gürkaynak et al., 2005a, for a discussion).

The proxy m_t is not required to be a correct measure of monetary shocks, as several forms of measurement error can be accounted for (see Mertens and Ravn, 2013). To construct a measure *correlated* with monetary policy shocks, we build on Altavilla et al. (2016) and use daily data on euro area government bond yields. In particular, we extract the common variation in sovereign spreads to Germany for different maturities of several crisis-hit countries around relevant monetary policy announcements by the ECB. Thereby, we extend the analysis of Altavilla et al. (2016) to a panel dimension across countries and maturities. Moreover, we use spreads instead of yields, following Rogers et al. (2016).

Specifically, we employ the regression

$$x_{ijt} = \alpha_i + \beta x_{ijt-1} + \sum_{a=1}^{A} \gamma_a D_{at} + \sum_{n=1}^{N} \delta_n z_{nt} + \eta_{ijt}, \tag{4.4}$$

on a daily frequency. In (4.4), x_{ijt} represents the sovereign bond spread versus Germany of country i on maturity j at time t, α_i are country-specific constants, D_{at} represents a dummy variable taking value 1 if the unconventional monetary policy announcement $a = 1, .., A$ took place at day t, otherwise zero, and z_{nt} controls for the release of macroeconomic news on variable $n = 1, .., N$. We include 139 macroeconomic news variables in z_{nt}, computed as the surprise component in economic data releases for the euro area, the UK, and the US, to attenuate the risk that the one day windows cover realizations of structural shocks that differ from the shocks of interest.[5]

The key coefficients in (4.4) are the estimated γ_a's. They capture the common variation in spreads in response to ECB announcement a. The vector $(\gamma_1, .., \gamma_A)'$ is transformed into one daily series m_t^D, taking value zero for non-announcement days and value γ_a on the day of announcement a. We then turn m_t^D into a monthly series

[5] For each variable, we construct a daily time series as the difference between the first-released data and the expected values, the latter corresponding to the median estimate of a panel of experts surveyed by Bloomberg, compare Chapter 3.

m_t^M by summing within months. Both m_t^D and m_t^M will be used for identification, after winsorizing them at 80% to control for outliers.[6]

We use $A = 32$ announcement days. They correspond to the days in which the ECB made explicit or implicit reference (either during regular meetings or other relevant speeches and communication) to at least one of the following three non-standard policy measures: forward guidance, credit easing, or quantitative easing. The choice of events closely follows Wright (2012) and Rogers et al. (2014), but is extended to include events through summer 2015. Since at the time of writing, in 2016, the sample of ECB unconventional monetary policy announcements is still relatively short, we do not distinguish among the precise types of monetary interventions, but aim at estimating the average effect of the measures. The first relevant event for our analysis occurred on August 22, 2007, the last one on January 22, 2015. The events comprise, for instance, announcements of long-term refinancing operations (LTROs), the securities market program (SMP), and outright monetary transactions (OMT). Appendix 4.A contains the full list of events.

To estimate (4.4), we use spreads of four countries and three maturities. Specifically, we use spreads of Ireland, Italy, Portugal, and Spain. We chose these countries because most of the non-standard measures in our sample were especially aimed at affecting the yields of these member states rather than those of Germany or other countries that were hit less by the crisis. For example, all four countries were covered by the SMP.[7] Regarding maturities, we use spreads for two, five, and ten years for two reasons. First, compared to bonds with longer maturity, these segments are typically more liquid, especially for the two smaller countries in the panel. Second, compared to bonds with shorter maturity, these segments are less constrained by the zero lower bound and, thus, provide more variation. We also select these countries and maturities because bonds and respective data are available throughout the full sample. We consider yields for different maturities rather than only for one maturity because the yield curve of all four countries was inverted at some point during the euro area debt crisis, when several important non-standard measures were announced. The inversion of the curve makes it *a priori* difficult to determine

[6]In the sensitivity analysis, we show that our results are qualitatively and also quantitatively relatively similar if we use the non-winsorized shock series.

[7]Greece was also contained in the SMP, but we exclude it from the estimation because its sovereign bonds were restructured in 2010 and because of a lack of data on two- and five-year yields.

which maturity best reflects the announced interventions; hence our use of several maturities. Finally, we use spreads instead of levels mainly to eliminate the effect of policy rate changes on the level of yields.[8]

4.2.3 Identification of the structural VAR

Having constructed a daily and a monthly measure correlated with unconventional monetary policy shocks, we now discuss how we use them for the identification of the structural VAR. Stock and Watson (2012) and Mertens and Ravn (2013) propose to identify the VAR using the regression

$$\hat{u}_{jt} = \alpha + \beta_i m_t + \eta_{jt}, \qquad i = 1, ..., k, \tag{4.5}$$

where $\hat{u}_{jt}$ is the estimated VAR reduced-form residual corresponding to equation j of model (4.1), and m_t is the instrument for ϵ_t^m at the same frequency as $\hat{u}_{jt}$. From these regression, the *relative* contemporaneous response of the variables in the VAR, b_j^m/b_i^m, can be obtained with i denoting the equation in which the policy indicator enters as dependent variable.[9] In other words, given equations (4.2) and (4.3), m_t allows for a consistent estimation of the *relative* contemporaneous response of the variables in the VAR to an unconventional monetary policy shock that changes the policy indicator by a scaled amount.

Furthermore, we follow Rogers et al. (2016) and use the event-study approach on high frequency data to refine the estimation of the relative contemporaneous response of the variables in the system. In a regression of the type

$$\Delta v_t = \gamma_1 + \gamma_2 \Delta r_t + \nu_t, \tag{4.6}$$

with Δv_t the first difference of a variable of interest and Δr_t the first difference of the policy indicator, the estimation of γ_2 is usually inconsistent since Δr_t is endogenous. The event-study approach exploits the fact that a consistent estimate

[8]In the impulse response analysis, we check that interest rates closely tied to the main refinancing rate of the ECB, such as the Euribor or Eurepo, do not react on impact to the identified policy innovations, supporting our interpretation that the latter reflect non-standard measures.

[9]The approach exploits the fact that, under equations (4.2) and (4.3), $E(u_t m_t) = b^m E(\epsilon_t^m m_t)$, hence $\hat{\beta}_j \xrightarrow{p} b_j^m \mu$ with $\mu = E(\epsilon_t^m m_t)/E(m_t)$ constant across j, and thus $\hat{\beta}_j/\hat{\beta}_i \xrightarrow{p} b_j^m/b_i^m$.

can be obtained if the periods in which Δr_t is exogenous can be isolated and that only those sub-periods are used in the regression (for an application, see Gürkaynak et al., 2005a, and Ehrmann and Fratzscher, 2005). In our application, these periods are the days of monetary policy announcements, so that we obtain a measure of exogenous variations in Δr_t on these days - our instrument m_t^D. The first best approach would then be to have data on the same daily frequency for all the variables included in the VAR, as this would allow estimating the relative contemporaneous effect of a structural shock of interest on all variables using (4.6).

However, daily data are only available for a subset of variables. We hence estimate β_j in model (4.5) with two separate types of regressions, in order to make use of high frequency data, whenever available. For variables y_{jt} available at a daily frequency, we estimate the corresponding element β_j by using all monetary policy announcement days and replacing the dependent variable in (4.6) with the daily first difference of the variable of interest, and replacing the regressor with the daily series m_t^D. For variables y_{jt} not available at a daily frequency, we approximate the unobserved high frequency first difference of the variable with the VAR innovation $\hat{u}_{jt}$ at a monthly frequency and use the monthly series m_t^M as regressor, as outlined above. In the baseline specification, for instance, we compute the variation on the daily frequency for the first two variables, namely the two-year rate and the VStoxx.

Note that for using the approach of Rogers et al. (2016), we have to make two assumptions. First, as standard in proxy SVAR models, we assume that the proxy is correlated with the structural shock of interest and uncorrelated with the other shocks, i.e. we assume that equations (4.3a and b) hold. Second, we assume that any shock to y_{jt} that occurs away from the time of the monetary policy announcements cannot be correlated with the jump that is associated with the monetary policy surprise. Under these two assumptions, the identification regressions that we carry out allow to identify the column of the structural impact matrix, $\tilde{b}^m$, that relates to the non-standard monetary policy shock (see Rogers et al., 2016).[10]

[10] The assumptions and the approach are in detail outlined in their paper. Also note that Rogers et al. (2016) implement their identification approach slightly differently than we do. They do not carry out identification regression (4.6) on a daily frequency, but rather sum the dependent daily variable and the instrument on the right hand side on announcement days within months, and then conduct the regression on the monthly frequency. Our results are both qualitatively and quantitatively basically unchanged if we implement the approach exactly as they do.

Equation (4.5) and its equivalent at the daily frequency also allow for an assessment of the strength of our instrument. We find that m_t is a strong instrument for our policy indicator. The F-statistic equals 40.39 and the β_j is positive. The high F-statistic suggests that a weak instrument problem is unlikely.[11]

4.3 Results

We discuss the effects of unconventional monetary policy using estimated impulse responses to a monetary policy innovation. The responses are reported along with their 90 percent confidence bands based on bootstrapping.[12] In all models, the shock is scaled to lower the average two-year rate on euro area bonds by 25 basis points. We first discuss the effectiveness and transmission of the monetary shock, then turn to the fiscal side effects, before finally evaluating the effects on country-specific variables, relative prices, and trade balances.

4.3.1 Effectiveness

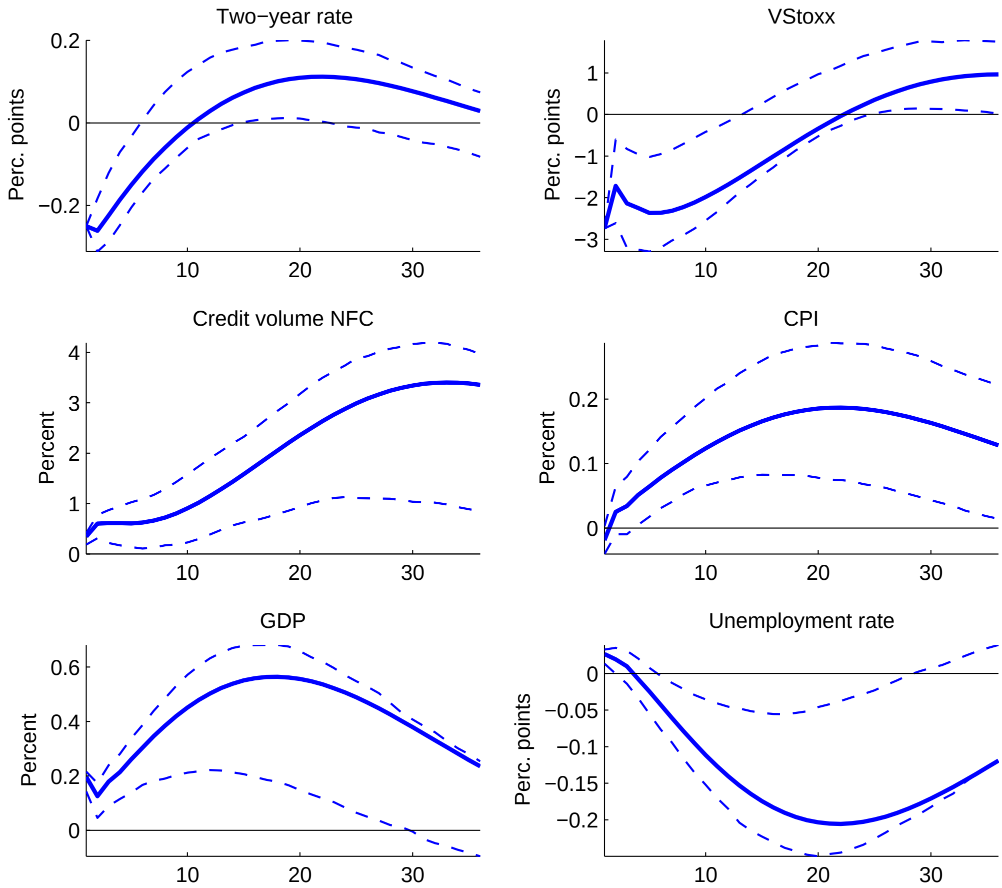
Figure 4.3.1 reports the impulse responses for the baseline VAR. The two-year rate drops on impact, as imposed, then remains significantly below trend for several months, before overshooting slightly after about one year. This surprise expansion leads to a significant and long-lasting reduction in risk aversion and uncertainty, as measured by the VStoxx. The volume of credit to non-financial corporations increases and reaches a peak after three years. These responses are associated with a gradual increase in consumer prices and GDP, with output peaking after about 18 months, slightly earlier than prices. The price dynamics are consistent with the overshooting in the two-year rate as an endogenous reaction of monetary policy. The responses of prices and output are also mirrored in the dynamics of the unem-

[11] The alternative monetary policy indicators considered in the sensitivity analysis are the five- and ten-year yield on euro area sovereign bonds excluding Germany. The F-statistics for these two indicators are 40.83 and 34.74, respectively.

[12] We apply a fixed-design wild bootstrap, as in Mertens and Ravn (2013) and Gertler and Karadi (2015). In principle, this procedure accounts for estimation errors in both stages of the structural VAR estimation (equations 4.1 and 4.5). For the variables identified on a daily frequency, however, the bootstrap procedure does not apply in the identification stage. Therefore, no confidence bands regarding the immediate impact are reported for these variables.

ployment rate, which bottoms after approximately two years, before returning to the level where it would have been without the monetary innovation.

Figure 4.3.1: Responses of baseline variables for the euro area



Note: The figure shows the estimated impulse responses, along with their 90 percent confidence bands, obtained using 500 bootstrap replications, of selected euro area variables to a monetary policy shock that lowers the average two-year rate by 25 basis points. The sample is 1999M1 through 2015M6.

Overall, the results are qualitatively in line with existing evidence on the effects of unconventional monetary policy shocks, but reveal several noteworthy quantitative differences. In Gambacorta et al. (2014), Weale and Wieladek (2016), and Boeckx et al. (2017), who identify policy surprises as shocks to central banks' balance sheets, output and prices respond faster, peaking approximately six to twelve months earlier, and reach their maximum simultaneously. Instead, we find a more sluggish response of both variables, peaking only after roughly two years, and with output leading prices. Interestingly, the dynamics implied by our estimates are more similar to responses to conventional monetary policy shocks (see Christiano et al., 1999, or Gertler and Karadi, 2015).

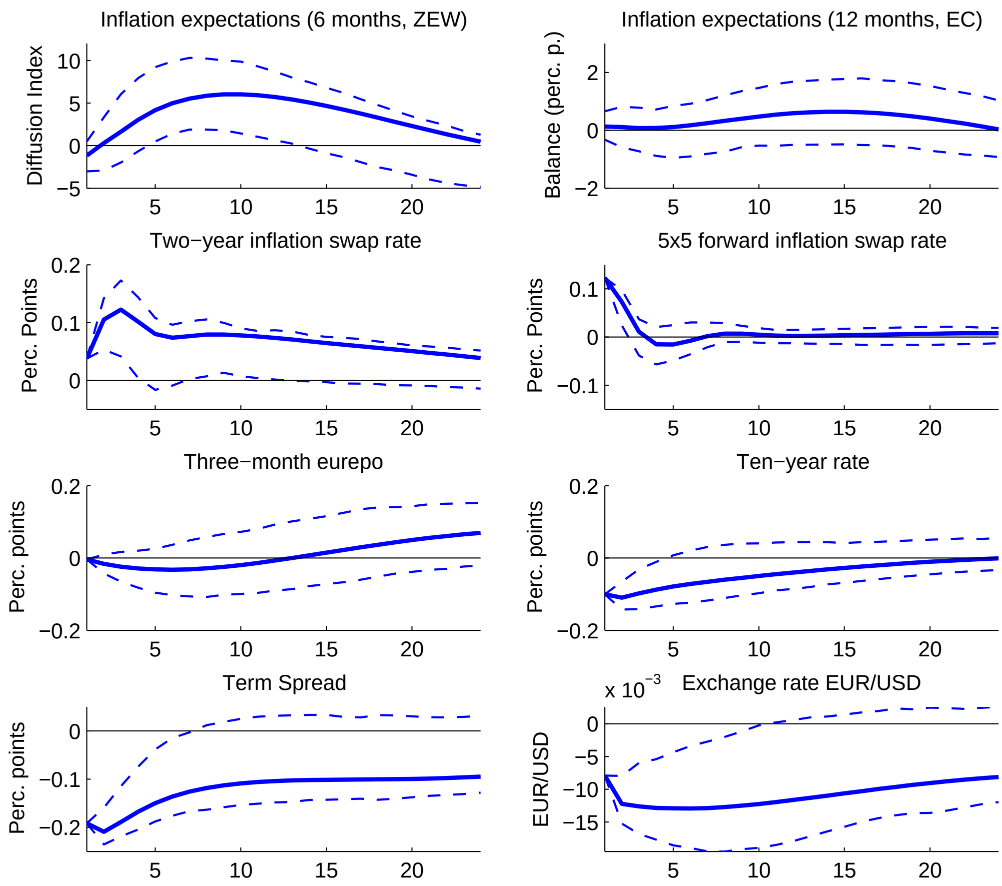
Regarding the effectiveness of monetary policy, our results are likewise more similar to the effects of conventional monetary policy. Gertler and Karadi (2015) find that in the US a shock to the one-year government bond rate of 20 basis points induces a maximum decline of output and prices of approximately 0.5 and 0.1 percent, respectively. Even though the effect on prices is only marginally significant in their case, while it is highly significant in our case, these point estimates are close to ours if we consider a contractionary shock. In contrast, the effects of comparably sized balance sheet shocks tend to be smaller. According to the estimates of Gambacorta et al. (2014), for example, a shock to central bank assets that lowers the VIX by one percentage point on impact has a peak effect on output and prices that is less than half of what we find if we rescale our shock to the two-year rate such that it lowers the VStoxx by one percentage point on impact.

We next evaluate the effects of the monetary surprise on several measures of inflation expectations, selected interest rates, and the Euro-Dollar exchange rate. Inflation expectations are a key determinant of actual inflation and interest and exchange rates represent important variables in the monetary transmission mechanism. Figure 4.3.2 contains the results. As outlined above, for this and all the following analyses, we augment the benchmark six-variable VAR with one additional variable at a time, and combine the responses of the marginal variables into one graph.[13] The figure shows the responses of two monthly survey-based measures of inflation expectations. The first is a survey of financial market experts, who are asked for their inflation expectations for the euro area over the next six months. The differential between the share of analysts who expect to see a rising inflation rate and the percentage who anticipate a falling inflation rate widens significantly five months after the shock, by about five percentage points. The survey is conducted by Centre for European Economic Research (ZEW). The second measure is a survey of con-

[13]Note that the sample may change depending on the marginal variable included. In particular, data on inflation swap rates is available only from 2008M9 onward, which considerably reduces the sample. The corresponding results for these variables should, thus, be treated cautiously. The figure notes contain details on the sample lengths for all estimations.

sumers, which assesses inflation expectations over a horizon of twelve months. The second survey is conducted by the European Commission. Inflation expectations increase as well according to this measure, but the rise is statistically insignificant.

Figure 4.3.2: Responses of inflation expectations, interest rates, and Euro-Dollar exchange rate



Note: The figure shows the estimated impulse responses, along with their 90 percent confidence bands, obtained using 500 bootstrap replications, of selected euro area variables to a monetary policy shock that lowers the average two-year rate by 25 basis points. The sample is 2008M9 through 2015M6 for the swap rates and 1999M1 through 2015M6 for the other variables.

The next two panels show the responses of two financial market-based measures of inflation expectations. The two-year swap rate increases significantly and rapidly in response to the shock. The behavior of the one-, five-, and ten-year swap rates are deferred to Figure 4.B.1 in Appendix 4.B. While all responses are qualitatively similar, swap rates for shorter maturities react more strongly and the effects last longer. From the impulse response of the five- and ten-year swap rate, we compute the five-year, five-year forward inflation swap rate, which has been one of the ECB's preferred measures of inflation expectations in recent years. Figure 4.3.2 shows that this indicator also increases significantly on impact, by about five basis points, and stays above trend for two months.

In the remaining panels, we analyze selected variables through which unconventional monetary policy surprises are potentially transmitted to the economy. The average ten-year rate on euro area government bonds (excluding Germany) and the term spread, defined as the difference between the response of the ten-year rate and the three-month Eurepo, both decline significantly on impact and stay below trend for roughly half a year. The exchange rate, on the other hand, appreciates. The latter finding is in line with Rogers et al. (2016) and can be rationalized by a reduction in break-up premia. The effect is relatively small, however. Finally, the insignificant response of the three-month Eurepo upon impact supports our identification strategy of using yield spreads in the first-stage regression instead of levels. It indicates that the identified monetary innovations reflect unexpected unconventional policy actions orthogonal to conventional policy rate changes.

Figures 4.B.2-4.B.4 in Appendix 4.B report the responses of further variables that provide additional insights and support our main results. Figure 4.B.2 shows that the real expansion is not limited to GDP but extends to alternative measure of real activity. Figure 4.B.3 shows that interest rates and asset prices in many other financial market segments are also affected by the shock, with stronger effects on rates of shorter maturities and riskier assets. Figure 4.B.4 shows that the monetary surprise expansion is associated with increases in credit volume and declines in credit rates for both households and non-financial corporations.

As a final step in this subsection, we evaluate the sensitivity of the baseline model to several alterations. Appendix 4.B contains the corresponding figures 4.B.5-4.B.7. First, we show that the responses of the six baseline variables change only slightly when adding the different marginal variables considered so far.[14] Second, we compute impulse responses without winsorizing the instrument before identification.

[14]The figure shows the response of the baseline variables when adding the marginal variables contained in Figures 4.3.2, 4.3.3, 4.B.2, 4.B.3, and 4.B.4, except for inflation swaps rates. For the latter, in fact, the data start only in 2008.

The results are qualitatively and quantitatively similar to the baseline results. Then, we include Germany into the computation of the average euro area two-year rate. The responses are more pronounced, given that the scale of the shock is the same and that German yields are less sensitive to the shock. Last, we use the five- or ten-year rate as policy indicator, instead of the two-year rate. The reaction of the baseline variables is relatively similar across specifications. Intuitively, the effects are stronger when yields for longer maturities unexpectedly drop by 25 basis points. All in all, the baseline results suggest that non-standard monetary surprise interventions by the ECB are effective in lifting economic activity, consumer prices, and inflation expectations.

4.3.2 Fiscal side effects

Next, we assess whether the identified monetary surprises have fiscal consequences. Such potential side effects are a primary concern of policymakers in many member countries, in the European Commission, as well as in the ECB itself (see Schmidt et al., 2015, Weidmann and Knot, 2015, Liikanen, 2015, European Commission, 2015, ECB, 2015). In particular, possible windfall gains, that is, savings on lower than expected public interest payments, are viewed as potentially generating skewed incentives and reducing governments' consolidation efforts. According to their 2013 country stability programs, for example, all of the four largest member states and Portugal planned reductions in the ratio of primary expenditure to GDP, ranging between 0.5 percentage points in Germany and 6.4 in Spain (see Table 4.3.1). However, according to the European Commission's assessment of the stability programs in 2015, all countries missed their target. At the lower end, Germany missed it by 0.7 percentage points and, at the higher end, Portugal by 2.6 percentage points. While there are several possible explanations for these misses, we assess whether there is evidence that unexpectedly lower government yields lead to lower public interest expenditures and higher public primary expenditures.

We start with an analysis for the euro area as a whole, using GDP weighted averages. The first four panels in Figure 4.3.3 show the behavior of the overall budget, the debt ratio, revenues, and expenditures. Consistent with standard theory, the average government balance in the euro area improves following the monetary surprise stimulus that lowers sovereign yields and raises output and prices. The maximum

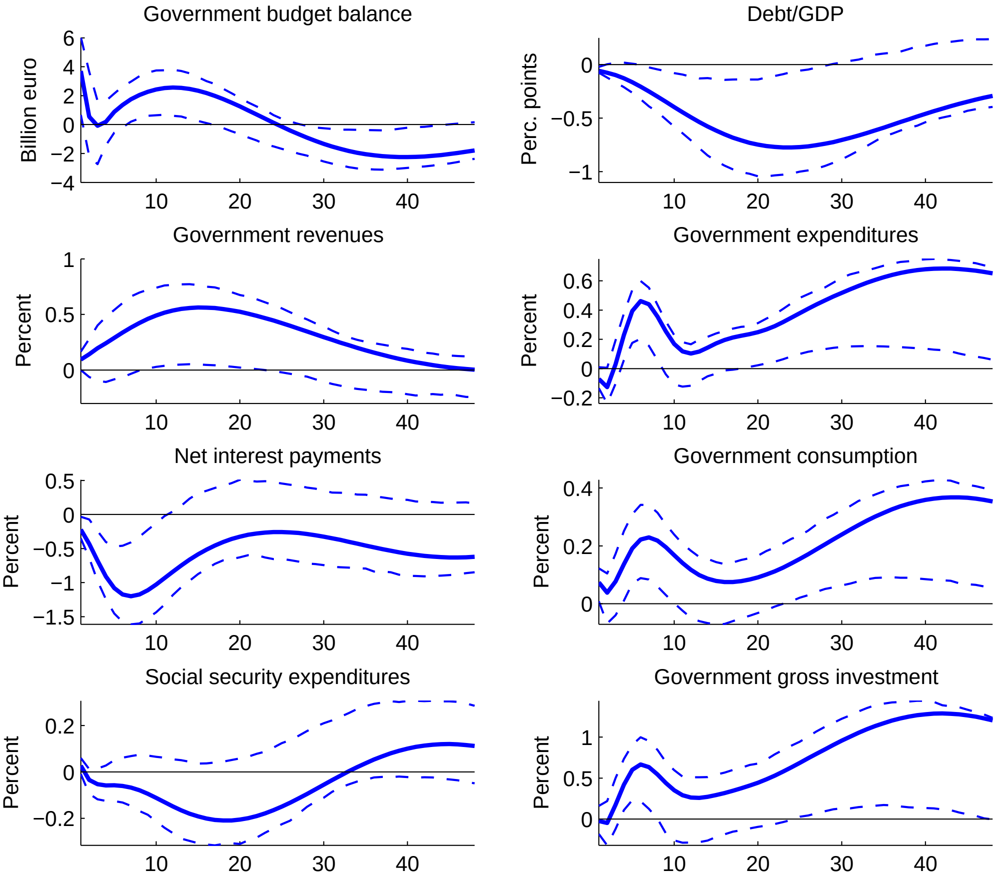
Table 4.3.1: Planned versus actual reduction in primary public expenditures 2012-2015 (change in percentage points of GDP)

Country	Stability program planned in 2013	Reduction assessed in stability programs 2015	Difference
Germany	-0.5	0.2	0.7
Spain	-6.4	-5.1	1.3
France	-1.2	0.6	1.8
Italy	-1.6	0.7	2.3
Portugal	-3.1	-0.5	2.6

response occurs after about one year and equals 2.5 billion euro. After about another year, however, the balance undershoots significantly. Due to the increase in output, the debt to GDP ratio nevertheless improves significantly. It declines by nearly one percentage point around two years after the shock. Decomposing the dynamics of the overall budget into changes in revenues and expenditures shows that revenues increase significantly as output exceeds trend. In line with conventional theory of automatic stabilizers (see Van den Noord, 2000), they thereby contribute to the improvement in the overall budget seen over the first two years after the shock. A quantitative comparison of the responses of revenues and GDP (see Figure 4.3.1) shows that there is nearly a one-to-one relationship between the two variables. This number is consistent with official estimates of the elasticity of revenues with respect to the output gap of unity in OECD countries.

The response of expenditures, on the other hand, is difficult to reconcile with the theory of automatic stabilizers. The official estimate of the elasticity of expenditures to the output gap in the euro area is –0.1 (see Girouard and André, 2006). This value would predict a small decline in expenditures when output increases. Moreover, in the special case of an interest rate shock that raises output, spending is expected to decline somewhat more strongly as public interest payments are likely to fall. In sharp contrast, the response of expenditures to the shock shows a strong, persistent, and mostly significant increase over a horizon of roughly four years. This finding rationalizes the undershooting of the overall balance and suggests that, on average across countries, fiscal policy is actively responding to non-standard monetary policy innovations in a procyclical manner.

Figure 4.3.3: Responses of government budget balance, debt, and expenditure by category



Note: The figure shows the estimated impulse responses, along with their 90 percent confidence bands, obtained using 500 bootstrap replications, of selected euro area variables to a monetary policy shock that lowers the average two-year rate by 25 basis points. The sample is 2003M1 through 2015M6 for the budget balance, 2000M1 through 2015M6 for debt/GDP, 2000M3 through 2015M6 for revenues and expenditures, and 2002M3 through 2015M6 for the detailed expenditure categories.

The bottom four panels decompose the dynamics of total expenditures into those of its components. As expected, the expansionary monetary shock leads to a significant reduction of net interest payments, which fall for about one year. The unanticipated drop in interest payments, in turn, is associated with a persistent increase in government consumption. Together, the two responses suggest that windfall profits from unexpectedly lower interest expenditure are partly used to increase intermediate good consumption and compensation of public employees. There is also some evidence of an increase in public investment. The latter is consistent with the decline in public interest rates, which renders public investments more profitable. Social security contributions, on the other hand, tend to fall as output rises and as the unemployment rate declines. Although it is not statistically significant, the decline in social expenditures is in line with the notion of automatic stabilizers on the spending side working mostly through unemployment benefits and age- and health-related outlays (see Darby and Melitz, 2008).

To quantify the average economic relevance of monetary policy shocks for the evolution of the different expenditure components we compute forecast error variance decompositions.[15] Specifically, Table 4.3.2 shows the percentage contribution of the monetary shock to the variance of the four spending categories. As fiscal policy responds only relatively slowly to the monetary shocks, the latter explain only a small fraction of the variability of the expenditure categories at shorter horizons. For longer horizons, however, they are a relevant driver of public expenditures. They account for between one fourth and one third of the forecast error variance at the 24-month horizon. Intuitively, they are particularly important for net interest payments and investment. However, they also explain 24 percent of the variability in government consumption.

Table 4.3.2: Percentage contribution of monetary policy shock to forecast errror variance of public expenditures (monthly horizon)

Horizon	Net interest payments	Government consumption	Social security expenditures	Government investment
1	2	3	2	2
6	2	2	2	4
12	13	12	11	19
24	34	24	22	36

Since, in the euro area, revenues and spending are largely determined at the member state level, we next study commonalities and differences in the response of fiscal policy to the common monetary surprise across member states. To focus

[15]Note that to be able to compute forecast error variance decompositions, we now additionally assume that the variance of the structural shocks is one, compare Chapter 3.

the discussion, we concentrate on three countries that were heavily and persistently affected by the sovereign debt crisis, Italy, Portugal, and Spain, as well as the two largest euro area economies, France and Germany. Combined, these countries provide a comprehensive picture of fiscal dynamics in the euro area as they account for more than 80% of the union's GDP. Moreover, we concentrate on the effects on government consumption as this is the most controversially discussed expenditure category in the public debate, other than public investment, and more directly controlled by the national fiscal authorities, unlike net interest payments or social security outlays.[16]

Figure 4.3.4 shows the peak effects of government consumption for the five countries following the expansionary monetary policy shock. They are all statistically significant. They are also economically relevant. At the maximum, public consumption increases by about one percent above trend in Spain and Portugal. In Italy and France it rises by roughly one half of a percent. In contrast, in Germany government consumption declines by approximately one half of a percent. Overall, these reactions are in line with the country-specific responses of sovereign yields to the common monetary shock (see next section). Sovereign yields decrease for Italy, Portugal, and Spain and tend to slightly increase in Germany and France.

4.3.3 Country heterogeneity and internal (im)balances

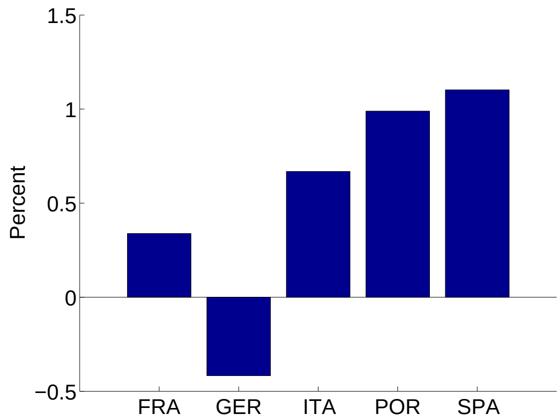
As a final step in the analysis, we investigate whether there is further evidence of heterogeneity in the reaction of the five economies to the common shock and whether the heterogeneous responses translate into relative price changes and movements in intra-union trade balances. Figure 4.3.5 contains the estimated peak effects of the monetary policy shock on country-specific two-year rates, local stock market indexes, GDP, and CPI.[17] All peak effects are statistically different from zero.

The figure shows a contrast between the maximum responses of the sovereign yields in France and Germany on the one hand, and Italy, Portugal and Spain on

[16]Figure 4.B.9 in Appendix 4.B shows the responses of all four expenditure categories in the different countries for completeness.

[17]The stock prices can also be understood as mirroring the development of uncertainty and risk aversion on a country level, with an inverted sign, as country-specific volatility indexes are not available for all countries. Figure 4.B.10 in Appendix 4.B shows the full responses of all four variables for all countries.

Figure 4.3.4: Peak responses of government consumption

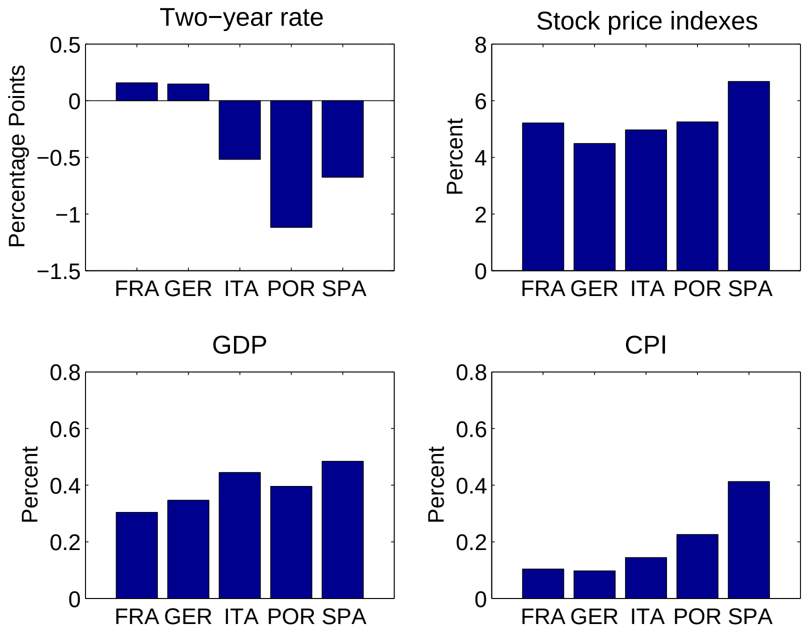


Note: The figure shows the estimated peak impulse responses of government consumption in selected countries to a monetary policy shock that lowers the average two-year rate on euro area government bonds by 25 basis points. The sample is 2002M3 through 2015M6 for Germany and 1999M1 through 2015M6 for the other countries.

the other hand. While yields increase in the former two countries, they decrease in the latter three. In the two largest economies the two-year rate rises by about 10 basis points. This positive–rather than negative–reaction of yields can be explained by at least two factors. First, government bonds of both countries were seen as a safe haven in euro-denominated securities markets, in particular during the height of the European debt crisis. As the non-standard policy interventions reduced uncertainty and increased risk appetite, the demand for safe-haven assets declined. Second, several of the policy measures contained in our sample most likely also affected the perceived risk of a break-up of the euro area. They thereby reduced revaluation risks contained in these bond prices.

In stark contrast, yields sharply decline by about 40, 60, and 100 basis points for Italy, Spain, and Portugal, respectively. This strong negative reaction to the common shock seems to be one relevant factor underlying the larger increase in primary expenditures in these three countries. Moreover, the huge drop in yields is associated with strong increases in equity prices, output and consumer prices. The responses of these variables are more pronounced than in France or Germany. In Spain, for example, the peak response of output and prices is about one half of a percent. Nevertheless, the peak responses of equity prices, GDP and consumer prices in France and Germany show that these two countries also profit from the expan-

Figure 4.3.5: Peak responses of further country-specific variables

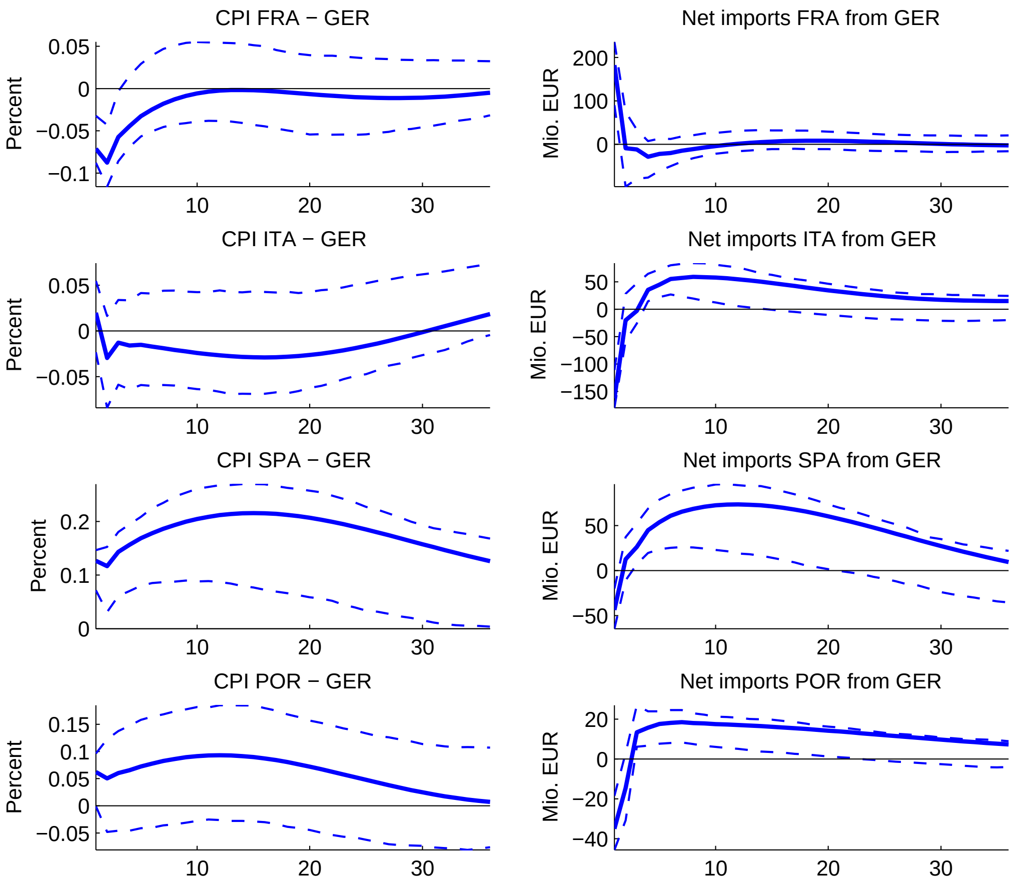


Note: The figure shows the estimated peak impulse responses of government consumption in selected countries to a monetary policy shock that lowers the average two-year rate on euro area government bonds by 25 basis points. The sample is 2000M1 through 2015M for the stock indices and 1999M1 through 2015M6 for the other variables.

sionary monetary shock, despite the increase in sovereign yields. This observation suggests that other forces might also be at play.

Specifically, relative price and demand developments within the monetary union might affect intra-euro area trade balances and thereby GDP. They may increase in particular surpluses in Germany, which is a large net exporting country, and deficits in other countries. We investigate this issue next. Figure 4.3.6 shows the responses of the CPI-difference of France, Italy, Portugal, and Spain relative to Germany; that is, the CPI based bilateral real exchange rates, together with the dynamics of the respective bilateral trade balances. While bilateral exchange rates seem largely unresponsive in France and Italy, they increase in Portugal and Spain. Relative prices in Spain rise significantly on impact and for more than two years. The maximum response is 0.2 percent after about 18 months. For Portugal, relative prices to Germany also increase, but the effect is less pronounced, barely missing statistical significance.

Figure 4.3.6: Responses of bilateral real exchange rates and trade balances



Note: The figure shows the estimated impulse responses, along with their 90 percent confidence bands, obtained using 500 bootstrap replications, of selected country-specific variables to a monetary policy shock that lowers the average two-year rate by 25 basis points. The sample is 2001M1 through 2015M6 for the net imports and 1999M1 through 2015M6 for the CPI differences.

These real exchange rate movements, together with changes in relative demand, are largely matched by the bilateral trade balances. While the response of net imports of France is not distinguishable from zero, Spanish net imports from Germany increase significantly roughly six months after the shock and reach a peak of more than 50 million. To put this number into perspective, a cumulative increase in net exports of approximately 600 million over a horizon of one year (50 million per month) is equivalent to a 1.5 percent increase in the *total* yearly trade deficit of Spain of 43 billion. Net imports of Italy and Portugal also increase significantly in reaction to the shock. Finally, the impact responses of the trade balances are consistent with the idea that price effects dominate their dynamics in the very short run. While Spain experiences a real appreciation and a drop in nominal net imports, in France a real depreciation is matched with a jump in net imports. In sum, the results indicate that there is a heterogeneous reaction across countries to common monetary policy shocks that entails some adverse side effects on relative prices and internal trade balances.

4.4 Conclusion

In this paper, we estimate the macroeconomic effects of unconventional monetary policy in the euro area using structural VARs, identified with an external instrument. We find that monetary interventions are effective in stabilizing the real economy and in countering risks to financial and price stability. An expansionary shock leads to an increase in consumer prices, output, and inflation expectations.

The analysis contributes to the literature on the macroeconomic effectiveness of monetary policy. Our results are qualitatively similar to existing contributions, which find that unconventional monetary policy is effective and is transmitted to the real economy mainly through interest rates (see Wright, 2012, Baumeister and Benati, 2013, Kapetanios et al., 2012). Quantitatively, the dynamics of output and prices implied by our estimates are more similar to the response of these variables to conventional interest rate innovations (see Christiano et al., 1999, Gertler and Karadi, 2015) than to unconventional monetary policy shocks identified through changes in the central bank balance sheet (see Gambacorta et al., 2014, Weale and Wieladek, 2016, and Boeckx et al., 2017).

In addition, our estimates complement existing studies on unconventional monetary policy by revealing several fiscal and distributional side effects of this policy. First, we provide evidence that primary fiscal expenditures rise significantly following a monetary surprise expansion. Second, we document a heterogeneous reaction of fiscal policy across the currency union to the common monetary policy shock and show that output and prices also respond differently. This heterogeneity, in turn, is associated with a divergence of relative prices and a widening of existing trade imbalances within the union.

All in all, our findings could be interpreted as containing a note of caution to monetary policymakers: policies that, in principle, support the economy might lead to laxer fiscal policy and a widening of internal imbalances, thereby creating the potential for increased risks to future financial and economic stability. On the other hand, the pro-cyclical response of fiscal variables could also be viewed as enhancing the effectiveness of monetary policy as it crowds in fiscal policy.

Appendix

4.A Data and sources

Table 4.A.1: Data construction and sources

Variable	Construction and source
Sovereign bond yields	Yield to redemption of sovereign bonds. Source: Datastream.
Euro area sovereign bond yields without Germany	Synthetic yields for euro area bonds are computed as weighted averages of nine individual countries: Austria, Belgium, Finland, France, Ireland, Italy, the Netherlands, Portugal, and Spain. The weights are taken from euro area benchmark bond yields in Datastream.
Stock market volatility	VStoxx option implied volatility. Source: Datastream.
Credit measures	Credit to non-financial firms, households, and monetary financial institutions. Source: ECB data warehouse. Seasonally adjusted with X-ARIMA-13.
Consumer price indices	Source: Datastream.
Real GDP and Industrial Production	Source: Datastream. Monthly IP series are seasonally adjusted with X-ARIMA-13. Quarterly GDP is interpolated using the series on IP and the method of Chow and Lin (1971).
Unemployment Rates	Source: Eurostat.
Inflation Expectations	Source of survey data: Centre for European Economic Research (ZEW), Germany, and European Commission. Source of inflation swaps: Datastream.
Real activity indicators	Retail sales, new car registrations, and new orders in manufacturing. Source: Datastream.
Other financial market variables	Eurepo, Euribor, EUR/USD spot exchange rate, Euro Stoxx 50, national stock price indices, yields of corporate bond indices with 2yr maturity and ratings AAA and BBB. Source: Datastream.

Oil Price	Price of Brent Crude Oil in US dollar. Source: Datastream.
Surprise component in economic data releases	Difference between the first-released data and the expected value (median expectation of a panel of experts surveyed by Bloomberg). The difference is divided by the standard deviation of the expectations. Source: Bloomberg. Variables from the following countries are included (see Table 4.A.2 for details): Euro Area, Germany, France, Italy, Spain, the UK, and the US.
Credit Rates	Source: ECB / Datastream. Cons. Credit: Personal Lending Rates, New Loans, Consumer Credit (Excluding Bank Overdrafts), 1-5 Years. House purchases: Personal Lending Rates, New Loans, House Purchases (Excluding Bank Overdrafts), 10 Years +. Loans to NFC (short): Prime Rates, New Loans, 1 Million Euro +, Excluding Bank Overdrafts, 1-5 Years.
Government budget balance / debt	Monthly euro area aggregated budget balance from Datastream. Seasonally adjusted with X-ARIMA-12. Converted to real terms using euro area CPI.
Government debt-to-GDP	Quarterly debt-to-GDP for ten individual countries (Austria, Belgium, Finalnd, France, Germany, Ireland, Italy, the Netherlands, Portugal, and Spain), aggregated using GDP weights. Source: Datastream. Seasonally adjusted with X-ARIMA-12 and linearly interpolated to monthly frequency.
Government revenues and expenditures (total and in detail)	Source: Eurostat (Consumption: intermediate consumption plus compensation of employees, social security expenditure: social benefits and social transfers in kind, gross investment: capital expenditure). Euro area aggregates based on data for ten individual countries: Austria, Belgium, Finalnd, France, Gremany, Ireland, Italy, the Netherlands, Portugal, and Spain. Quarterly data is seasonally adjusted with X-ARIMA-12/13 and then linearly interpolated to the monthly frequency. Converted to real terms using CPIs.
Net imports	Imports minus exports vis-a-vis Germany. Source: German Federal Statistical Office (Destatis). Seasonally adjusted with X-ARIMA-13.

Table 4.A.2: List of included economic data releases

Country	Economic data release
Euro area	EC Bus. Climate Ind., Current Account Net WDA SA, EC Cons. Conf. Ind., CPI YoY, CPI MoM, BOP Current Account Net NSA, New Orders (Manu.) YoY, Gr. Fx. Capital Formation QoQ, EC Serv. Conf. Ind., Markit Comp. PMI SA, Markit Serv. PMI SA, Retail Sales Vol. YoY WDA, Retail Sales Vol. MoM SA, ZEW Exp. of Econ. Growth (Econ. Sent.), Trade Bal. with non EZ Countries, M3 Money Supply 3 M. Mov. Avg., PPI Industry Ex Constr.YoY, PPI Industry Ex Constr. MoM, Unem. Rate, GDP SA QoQ (real SA)
France	CPI YoY, CPI MoM, Cons. Conf. Ind., Bus. Conf. Ind. (Manu.), Prod. Outlook Ind., Bus. Sent. Ind., Cons. Spending MoM, CPI ex Tobacco, real GDP QoQ, real GDP YoY, ML & OS Unemployment Rate, Markit Manu. PMI SA, Markit Serv. PMI SA, PPI MoM, PPI YoY, Jobseekers Total SA, Trade Balance EUR, Manu. Prod. MoM SA, Own-Comp. Prod. Outlook
Germany	CPI YoY, CPI MoM, Manu. Ord. YoY NSA, Manu. Ord. MoM SA, Trade Balance Val. Exp. MoM SA, Trade Balance Val. Imp. MoM SA, Trade Balance EUR NSA, Retail Sales NSA YoY, Retail Sales SWDA MoM, Prod. Prices MoM, Ind. Prod. YoY NSA WDA, Ind. Prod.n MoM SA, Ind. Prod. YoY SA, Unem. Rate SA, Unem. Change SA, Ifo Pan Bus. Climate, Ifo Pan Bus. Expectations, Current Account EUR, Import Price Index MoM, Markit Manu. PMI SA, GDP Priv.Cons. QoQ, GDP Gr. Fx. Capital Inv. QoQ, GDP Inv. in Const. QoQ

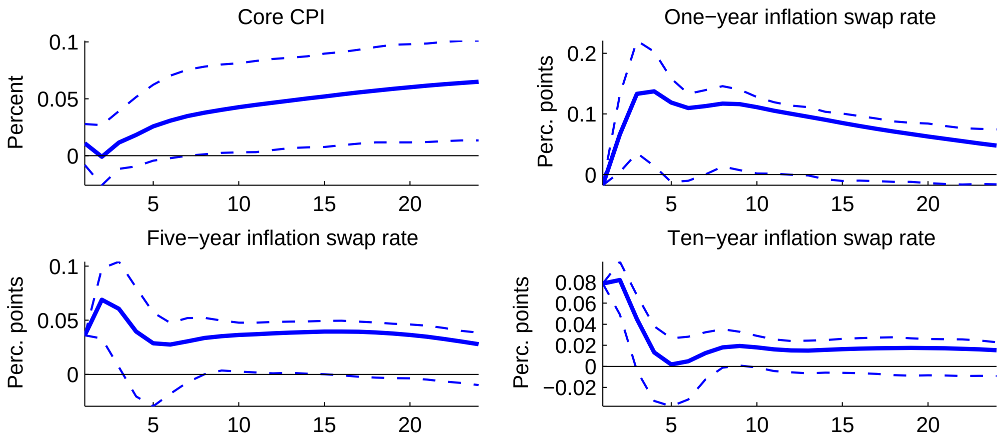
USA	CPI YoY NSA, CPI Ex. Fd. & En. YoY NSA, Pers. Cons. Exp. CPI YoY SA, CPI MoM SA, Un. of Michigan Cons. Conf. Ind., Gov. Budget Balance, Cons. Spend. GR MoM SA, Markit Manu. PMI SA, Ind. Prod. MoM SA, Core PPI, PPI - Fin. Goods, In. Jobless Claims SA, Housing Starts/Permits, Diff. between Exp. and Imp., GDP QoQ SAAR, PPI Fin. Goods SA MoM%, Cap. Util.n % of Tot. Cap., Bus. Inventories MoM SA, Avg. Hourly Earnings YoY% SA, Avg. Hourly Earnings MoM% SA, Constr. Spend. Total MoM SA, Dur. Goods New Orders Ind. MoM SA, Conf. Board Leading Ind. MoM, Prod. Output ph Nonfarm Bus. Sec. QoQ SA, Unit Labor Costs Nonfarm Bus. Sec. QoQ % SAAR, Retail Sales SA MoM % Change, Personal Income MoM SA, Nonfarm Payrolls Total MoM SA
Italy	CPI NIC Incl Tbc. YoY NSA, CPI NIC Incl Tbc. MoM NSA, Cons. Conf. Ind. SA, Bus. Conf. Manu. Sector, Hourly Wages MoM SA, Ind. Orders YoY NSA, Ind. Orders MoM SA, Ind. Prod. YoY WDA, Ind. Prod. MoM SA, Ind. Prod. YoY, Ind. Sales YoY, Ind. Sales MoM SA, Manu. PMI SA, Serv. PMI SA, PPI Manu. MoM, PPI Manu. YoY, PPI Manu. YoY, Priv. Cons. QoQ SA WDA, Retail Sales MoM SA, Retail Sales YoY, Trade Balance Total, Unem. Rate SA, Real GDP YoY SA WDA, Trade Balance Non EU NSA
UK	CPI Ex En. Fd. Alc.l & Tbc. YoY, GDP YoY, GDP MoM, Retail Sales Ex Auto. Fuel YoY SA, Retail Sales Ex Auto. Fuel MoM SA, PPI Manu.Prod. YoY NSA, PPI Manu. Prod. MoM NSA, PPI Input Prices MoM NSA, PPI Input Prices YoY NSA, Ind. Prod. YoY SA, Unem. Rate SA (Change), Markit/CIPS Const. PMI SA, Markit/CIPS Serv PMI SA, Govt. Budget Balance, Priv. Cons. QoQ, House Price Ind. MoM SA, Cons. Conf. Ind., Gov. Spending QoQ
Spain	CPI YoY, CPI Core YoY, PPI MoM, Trade Balance EUR, Reg. Unem. Level MoM Net Change, Avg Labor Costs per Worker YoY, PMI Manu. SA

Table 4.A.3: List of included ECB Monetary Policy Announcements

Date	Policy Announcement
22.08.2007	Supplementary liquidity-providing longer-term refinancing operation (LTRO) with a maturity of three months
28.03.2008	LTROs with a maturity of six months
29.09.2008	Special term refinancing operation
08.10.2008	Fixed rate tender procedure with full allotment on the main refinancing operation(MROs)
15.10.2008	List of assets eligible as collateral in Eurosystem credit operations extended
07.05.2009	LTROs with a maturity of one year
04.06.2009	Details on Purchase program for covered bonds (CBPP)
03.12.2009	Phasing out of 6-month LTROs, indexation of new one year LTROs
04.03.2010	Phasing out of 3-month LTROs, indexation of six month LTROs
10.05.2010	Securities Markets Program (SMP)
28.07.2010	Risk control measures in collateral framework reviewed
03.03.2011	Further LTROs
09.06.2011	MROs as fixed rate tender procedures with full allotment (FRFA) for as long as necessary, at least until October 2011
04.08.2011	Further LTROs with a maturity of three and six months
08.08.2011	ECB will actively implement its Securities Market Program
06.10.2011	New covered bond purchase program (CBPP2)
08.12.2011	Two additional LTROs with a maturity of three years
21.12.2011	Results of first three year LTRO
09.02.2012	ECB's Governing Council approves eligibility criteria for additional credit claims
28.02.2012	Results of second three year LTRO
06.06.2012	FRFA on MROs as long as necessary, and at least until January 2013
26.07.2012	'Whatever it takes...' speech by ECB President Mario Draghi in London
02.08.2012	Outright Monetary Transactions program (OMT)
06.09.2012	Technical features of OMT
06.12.2012	FRFA on MROs as long as necessary, and at least until July 2013
22.03.2013	Collateral rule changes for some uncovered government guaranteed bank bonds
02.05.2013	FRFA on MROs as long as necessary, and at least until July 2014
04.07.2013	Governing Council expects the key ECB interest rates to remain at present or lower levels for an extended period of time (open-ended forward guidance)
08.11.2013	FRFA on MROs as long as necessary, and at least until July 2015
05.06.2014	Targeted longer-term refinancing operations (TLTROs)
03.07.2014	Details on TLTROs published
22.01.2015	Expanded asset purchase program

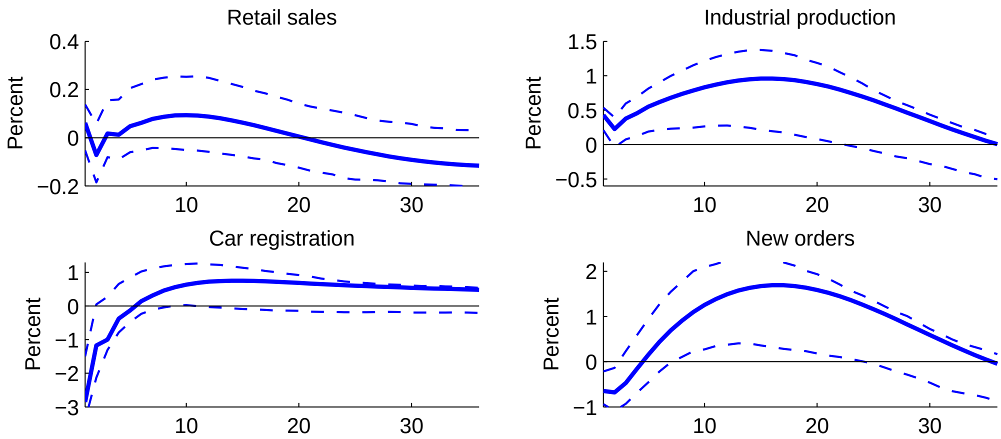
4.B Additional figures

Figure 4.B.1: Responses of alternative measures of prices and inflation expectations



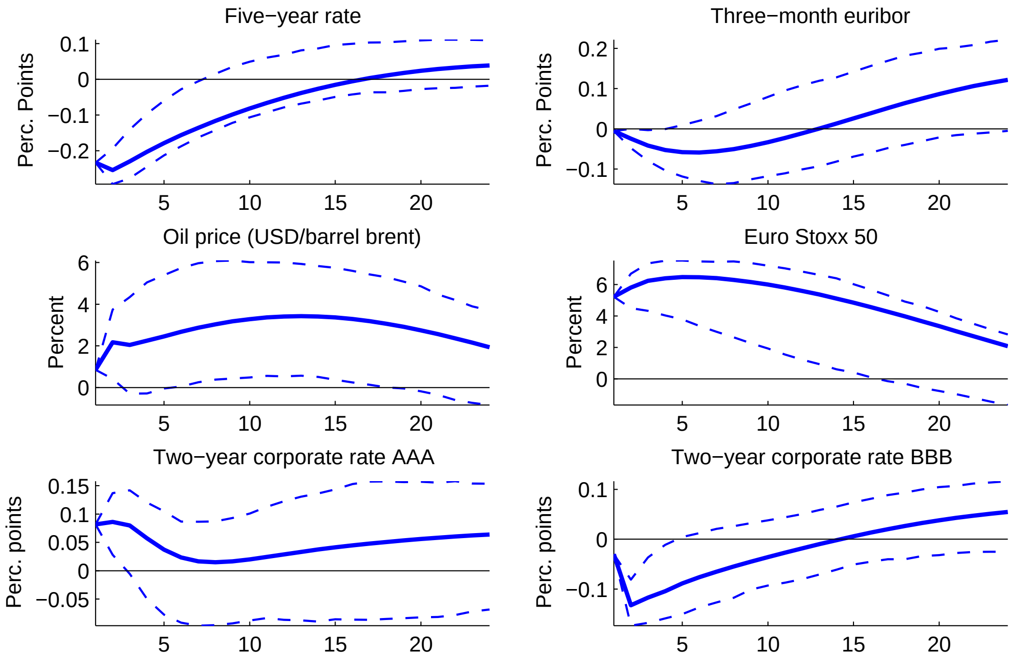
Note: The figure shows the estimated impulse responses, along with their 90 percent confidence bands obtained using 500 bootstrap replications, of selected euro area variables to a monetary policy shock that lowers the average two-year rate by 25 basis points. The sample is 1999M1 through 2015M6 for core inflation and 2008M9 through 2015M6 for the other variables.

Figure 4.B.2: Responses of alternative measures of economic activity



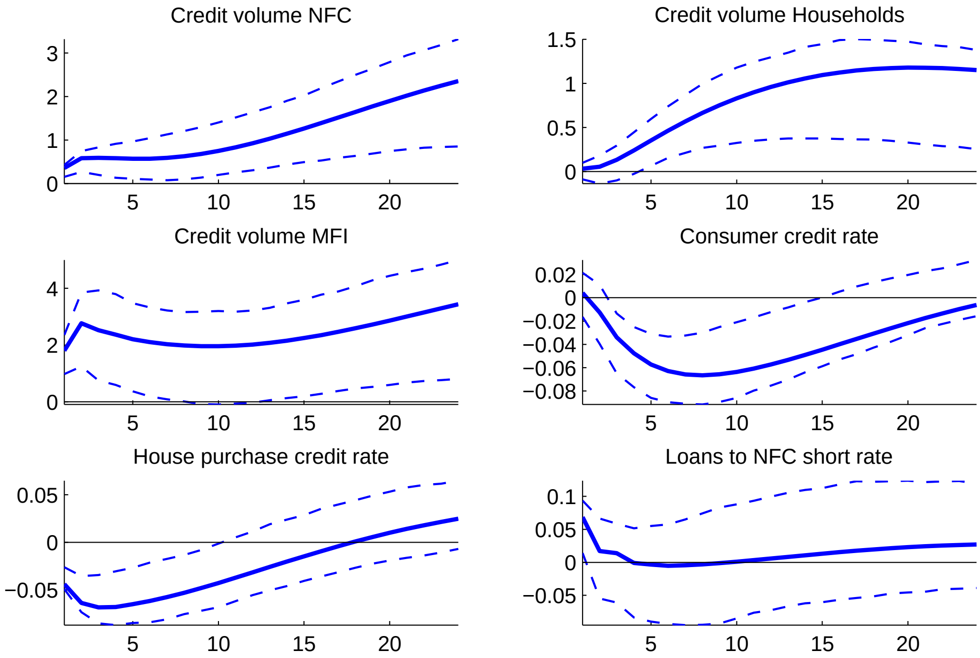
Note: The figure shows the estimated impulse responses, along with their 90 percent confidence bands obtained using 500 bootstrap replications, of selected euro area variables to a monetary policy shock that lowers the two-year rate on eurobonds by 25 basis points. The sample is 1999M1 through 2015M6.

Figure 4.B.3: Responses of other financial variables



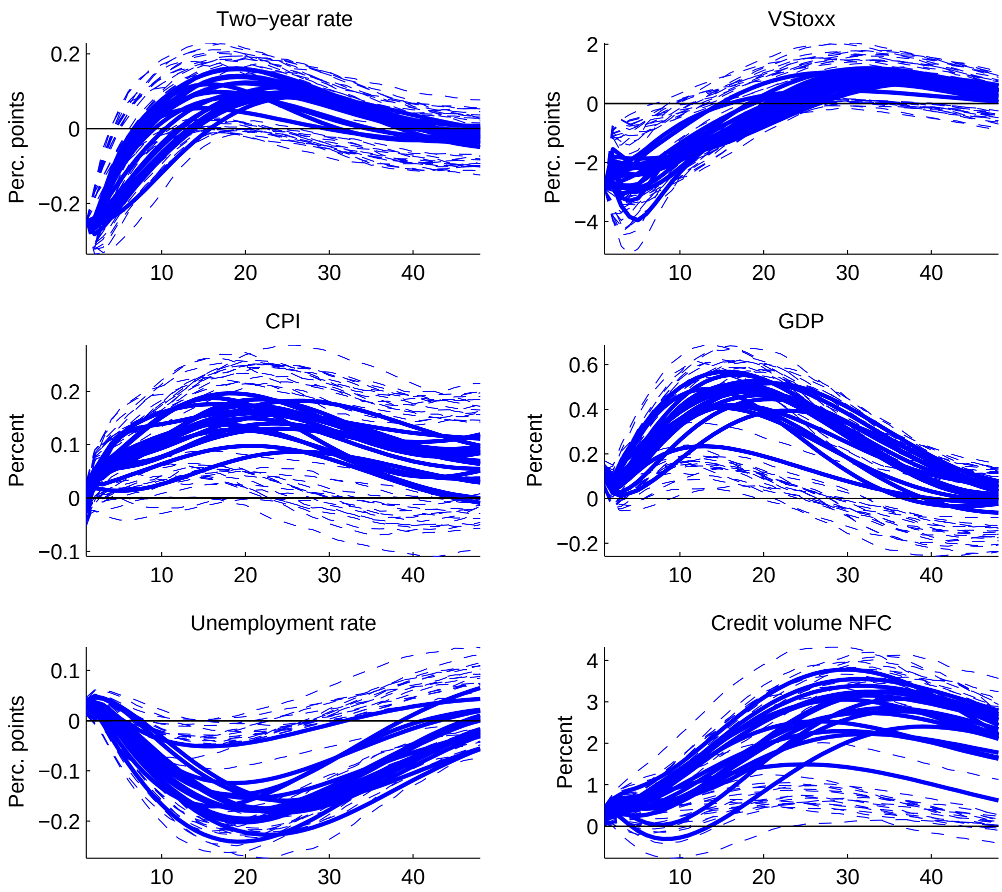
Note: The figure shows the estimated impulse responses, along with their 90 percent confidence bands obtained using 500 bootstrap replications, of selected euro area variables to a monetary policy shock that lowers the average two-year rate by 25 basis points. The sample is 1999M1 through 2015M6 for the first four variables and 2002M4 through 2015M6 for the corporate rates.

Figure 4.B.4: Responses of credit volume and credit rates



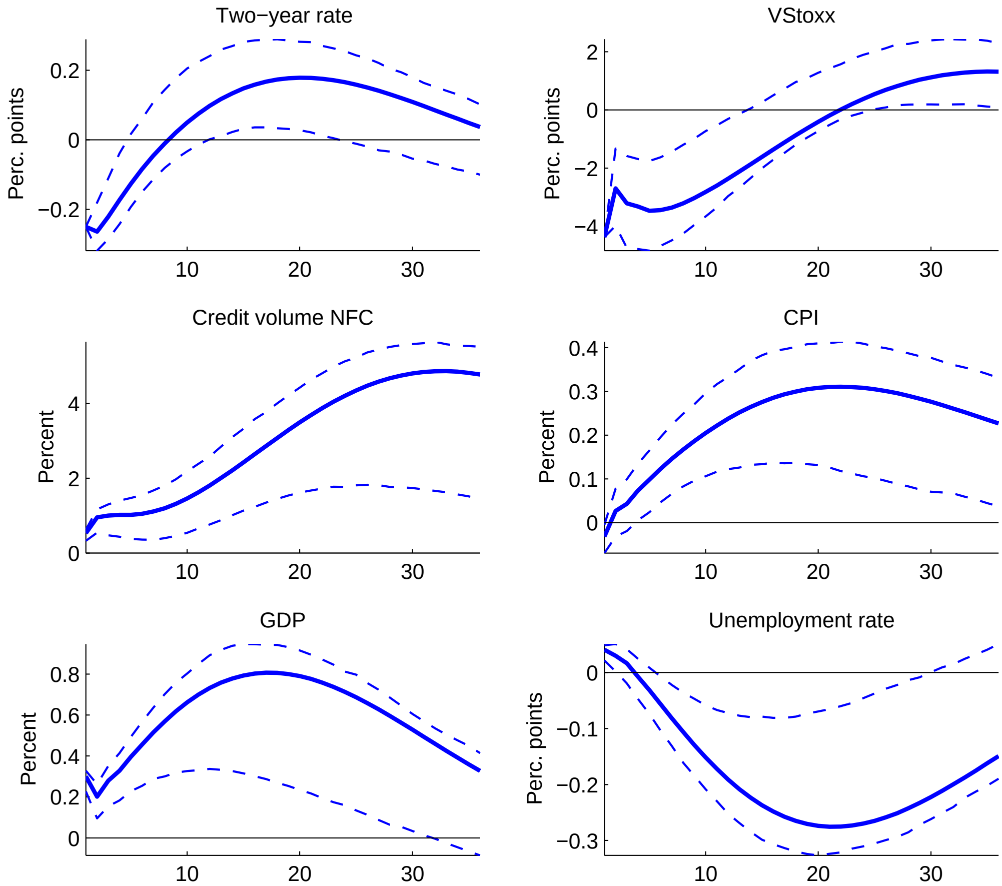
Note: The figure shows the estimated impulse responses, along with their 90 percent confidence bands obtained using 500 bootstrap replications, of selected euro area variables to a monetary policy shock that lowers the average two-year rate by 25 basis points. The sample is 2003M1 through 2015M6 for the house purchase rate, 2000M1 through 2015M6 for the other rates, and 1999M1 through 2015M6 for the credit volumes.

Figure 4.B.5: Robustness of baseline specification to including additional variables



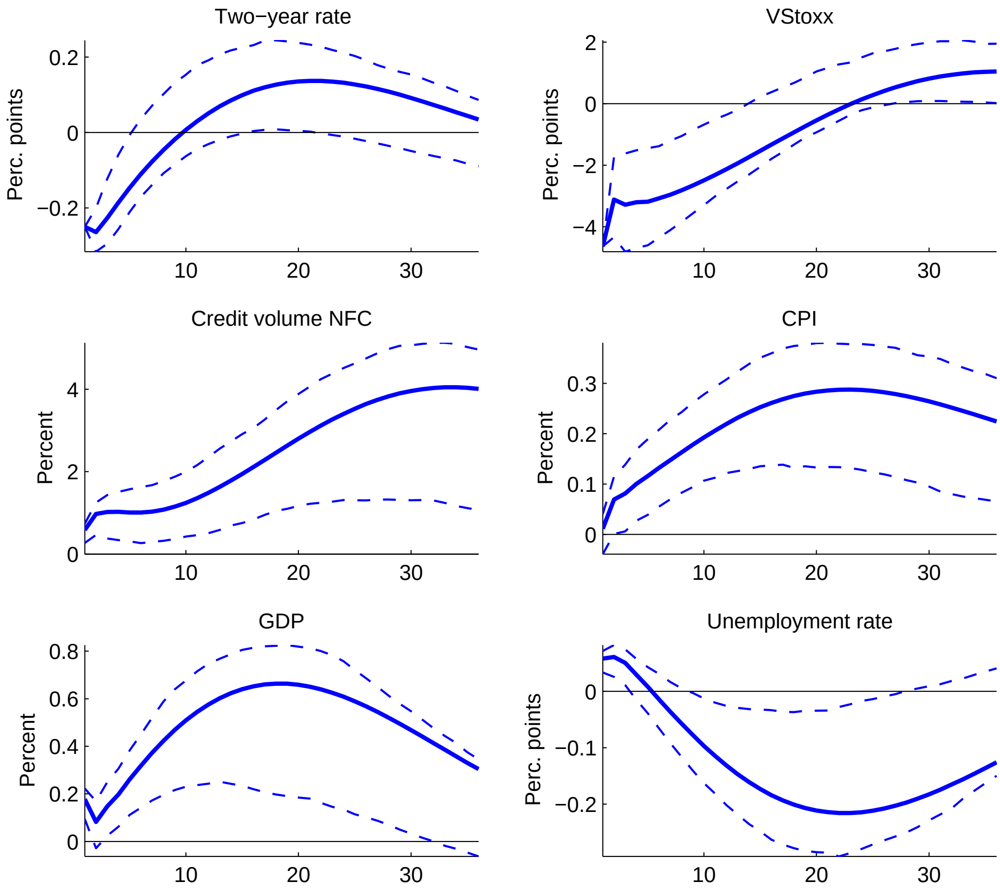
Note: The figure shows the estimated impulse responses, along with their 90 percent confidence bands obtained using 500 bootstrap replications, of the benchmark variables to a monetary policy shock that lowers the average two-year rate by 25 basis points when including additional variables, one at a time, to the benchmark VAR. Included are the variables from Figures 4.3.2 (expect for the swap rates), 4.3.3, 4.B.2, 4.B.3, and 4.B.4. The sample depends on the included marginal variable.

Figure 4.B.6: Robustness of baseline specification to not performing a Windsorization of the external instrument



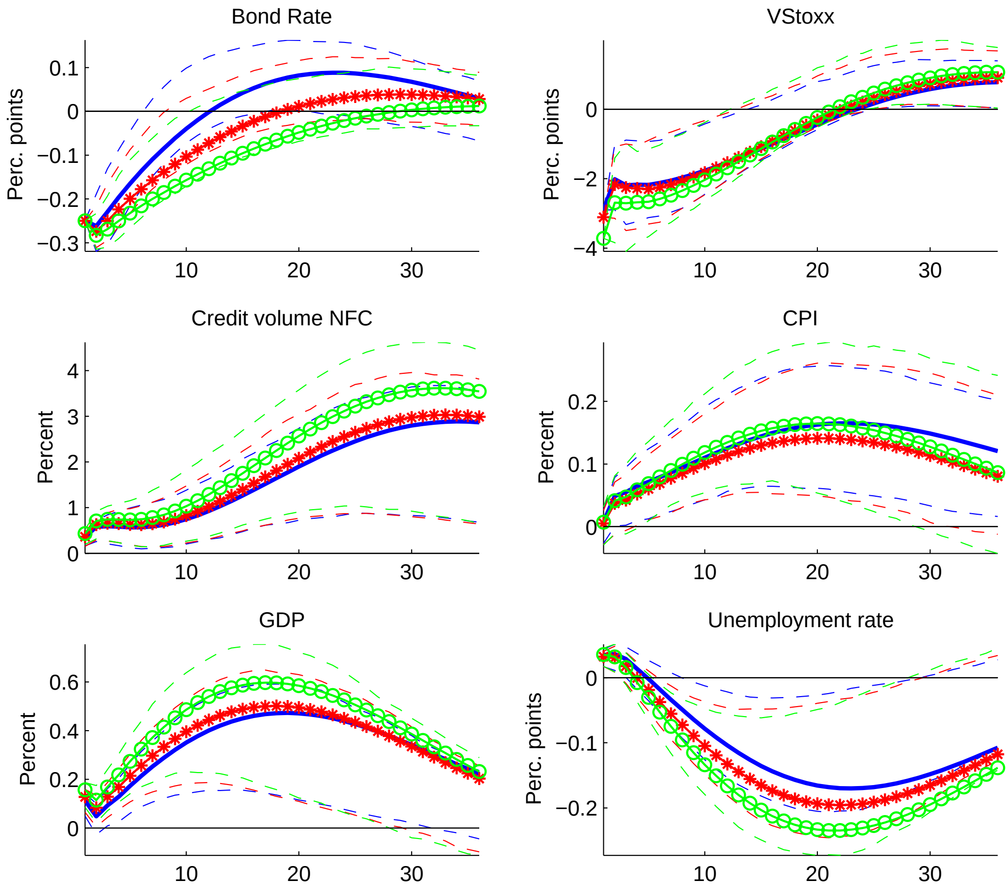
Note: The figure shows the estimated impulse responses, along with their 90 percent confidence bands obtained using 500 bootstrap replications, of selected euro area variables to a monetary policy shock that lowers the average two-year rate by 25 basis points. Different to the baseline specification, the instrument used to identify the monetary policy shock is not windsorized. The sample is 1999M1 through 2015M6.

Figure 4.B.7: Robustness of baseline specification to including German bond yields in the computation of the euro area average two-year rate



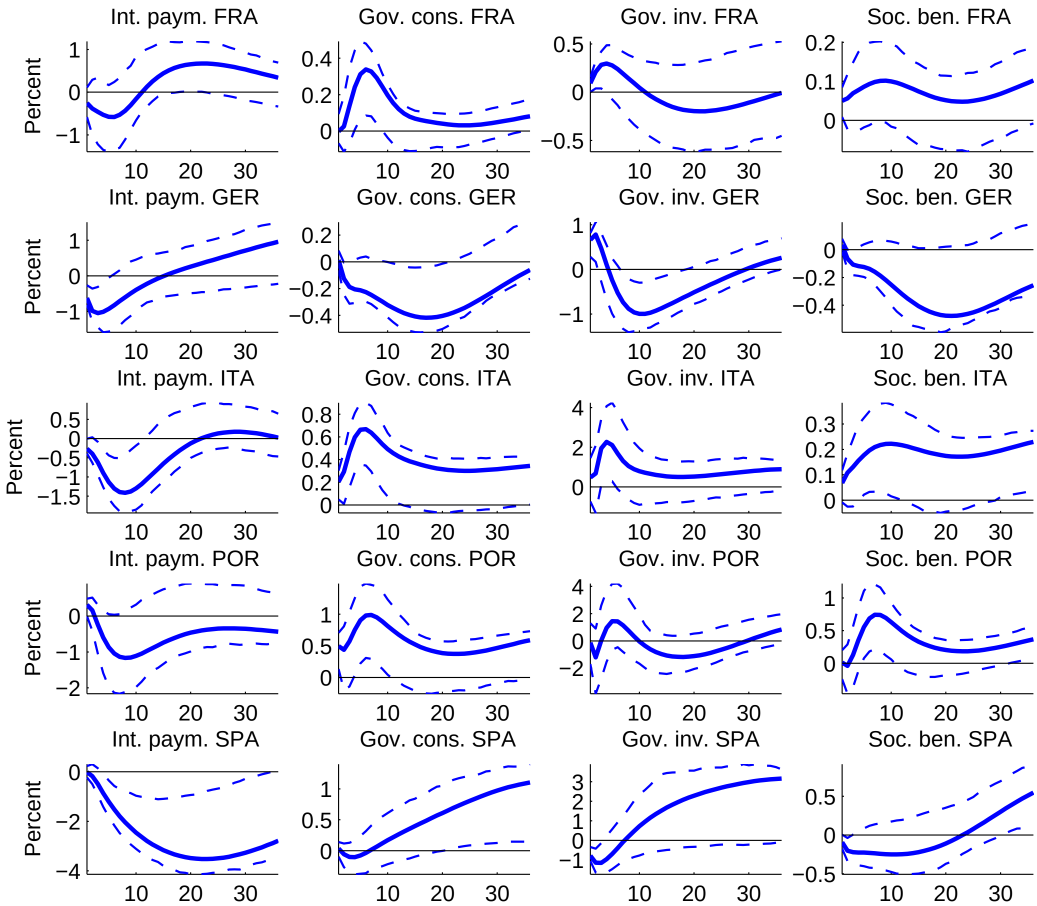
Note: The figure shows the estimated impulse responses, along with their 90 percent confidence bands obtained using 500 bootstrap replications, of selected euro area variables to a monetary policy shock that lowers the average two-year rate by 25 basis points. Different to the baseline specification, German yields are included in constructing the euro area bond rate. The sample is 1999M1 through 2015M6.

Figure 4.B.8: Robustness of baseline specification to using average euro area sovereign rates for different maturities as policy indicator



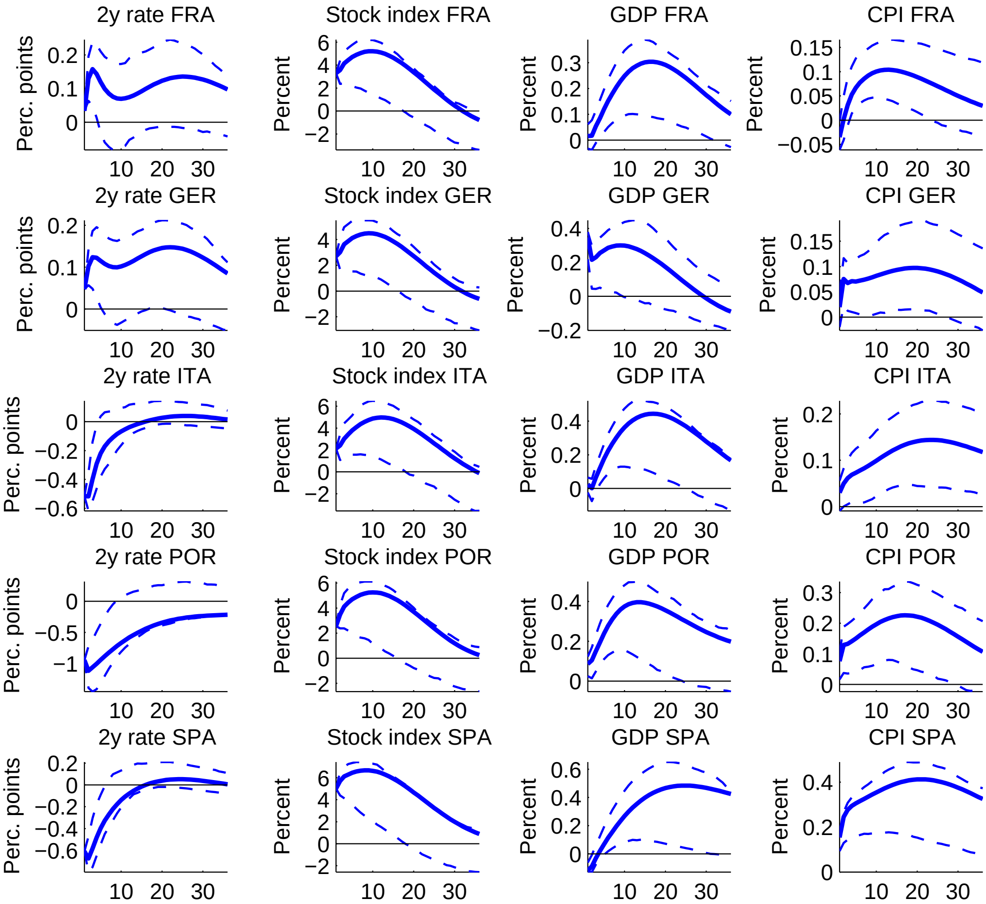
Note: The figure shows the estimated impulse responses, along with their 90 percent confidence bands obtained using 500 bootstrap replications, of selected euro area variables to a monetary policy shock that lowers either the average two-year rate (blue lines), the five-year rate (red lines, Asterisk) or the ten-year rate (green lines, Circles) by 25 basis points. The sample is 1999M1 through 2015M6.

Figure 4.B.9: Responses of national government expenditures



Note: The figure shows the estimated impulse responses, along with their 90 percent confidence bands, obtained using 500 bootstrap replications, of selected country-specific government expenditure components to a monetary policy shock that lowers the average two-year rate by 25 basis points. The sample is 2002M3 through 2015M6 for the German variables and 2000M3 through 2015M6 (interest payments) or 1999M1 through 2015M6 (consumption, investments, benefits) for the other countries.

Figure 4.B.10: Responses of selected country-specific variables



Note: The figure shows the estimated peak impulse responses of selected country-specific variables to a monetary policy shock that lowers the average two-year rate by 25 basis points. The sample is 2000M1 through 2015M for the stock indexes and 1999M1 through 2015M6 for the other variables.

Bibliography

Acharya, V.V., L.A. Lochstoer and T. Ramadorai (2013): Limits to arbitrage and hedging: Evidence from commodity markets, *Journal of Financial Economics*, 109(2), 441 – 465.

Ahmed, S. and A. Zlate (2014): Capital flows to emerging market economies: A brave new world?, *Journal of International Money and Finance*, 48, 221–248.

Aizenman, J., M. Binici and M.M. Hutchison (2016a): The Transmission of Federal Reserve Tapering News to Emerging Financial Markets, *International Journal of Central Banking*, 12(2), 317–356.

Aizenman, J., M.D. Chinn and H. Ito (2016b): Monetary policy spillovers and the trilemma in the new normal: Periphery country sensitivity to core country conditions, *Journal of International Money and Finance*, 68(C), 298 – 330.

Altavilla, C., D. Giannone and M. Lenza (2016): The Financial and Macroeconomic Effects of the OMT Announcements, *International Journal of Central Banking*, 12(3), 29–57.

Arias, J.E., C. Erceg and M. Trabandt (2016): The macroeconomic risks of undesirably low inflation, *European Economic Review*, 88(C), 88–107.

Arias, J.E., J.F. Rubio-Ramirez and D.F. Waggoner (2014): Inference Based on SVARs Identified with Sign and Zero Restrictions: Theory and Applications, International Finance Discussion Papers No. 1100, Board of Governors of the Federal Reserve System.

Aulerich, N.M., S.H. Irwin and P. Garcia (2013): Bubbles, Food Prices, and Speculation: Evidence from the CFTC's Daily Large Trader Data Files, in: *The Economics of Food Price Volatility*: National Bureau of Economic Research, Inc, NBER Chapters.

Autrup, S.L. and M. Grothe (2014): Economic surprises and inflation expectations: Has anchoring of expectations survived the crisis?, European Central Bank Working Paper Series No. 1671, European Central Bank.

Azis, I.J. and H.S. Shin (2015). *The Three Phases of Global Liquidity*, pp. 7–44. Singapore: Springer Singapore.

Baffes, J. (2007): Oil spills on other commodities, Policy Research Working Paper Series No. 4333, The World Bank.

Balduzzi, P., E.J. Elton and T.C. Green (2001): Economic News and Bond Prices: Evidence from the U.S. Treasury Market, *Journal of Financial and Quantitative Analysis*, 36(04), 523–543.

Basak, S. and A. Pavlova (2016): A Model of Financialization of Commodities, *Journal of Finance*, 71(4), 1511–1556.

Bauer, M.D. (2015): Inflation Expectations and the News, *International Journal of Central Banking*, 11(2), 1–40.

Baumeister, C. and L. Benati (2013): Unconventional Monetary Policy and the Great Recession: Estimating the Macroeconomic Effects of a Spread Compression at the Zero Lower Bound, *International Journal of Central Banking*, 9(2), 165–212.

Beechey, M.J., B.K. Johannsen and A.T. Levin (2011): Are Long-Run Inflation Expectations Anchored More Firmly in the Euro Area Than in the United States?, *American Economic Journal: Macroeconomics*, 3(2), 104–29.

Bekaert, G., M. Hoerova and M. Lo Duca (2013): Risk, uncertainty and monetary policy, *Journal of Monetary Economics*, 60(7), 771–788.

Bernanke, B.S. (2007): Inflation expectations and inflation forecasting, Speech No. 306, Board of Governors of the Federal Reserve System.

Bertaut, C. and R. Judson (2014): Estimating U.S. Cross-Border Securities Positions: New Data and New Methods, International Finance Discussion Papers No. 1113, Board of Governors of the Federal Reserve System.

Bertaut, C.C. and R.W. Tryon (2007): Monthly Estimates of U.S. Cross-border Securities Positions, International Finance Discussion Papers No. 910, Board of Governors of the Federal Reserve System.

Bhardwaj, G., G.B. Gorton and K.G. Rouwenhorst (2014): Fooling Some of the People All of the Time: The Inefficient Performance and Persistence of Commodity Trading Advisors, *Review of Financial Studies*, 27(11), 3099–3132.

Blanchard, O. and F. Giavazzi (2002): Current Account Deficits in the Euro Area: The End of the Feldstein Horioka Puzzle?, *Brookings Papers on Economic Activity*, 33(2), 147–210.

Blanchard, O.J. and D. Quah (1989): The Dynamic Effects of Aggregate Demand and Supply Disturbances, *American Economic Review*, 79(4), 655–673.

Blinder, A.S., M. Ehrmann, M. Fratzscher, J.D. Haan and D.J. Jansen (2008): Central Bank Communication and Monetary Policy: A Survey of Theory and Evidence, *Journal of Economic Literature*, 46(4), 910–45.

Boeckx, J., M. Dossche and G. Peersman (2017): Effectiveness and Transmission of the ECB's Balance Sheet Policies, *International Journal of Central Banking*, 13(1), 297–333.

Bouakez, H. and M. Normandin (2010): Fluctuations in the foreign exchange market: How important are monetary policy shocks?, *Journal of International Economics*, 81(1), 139 – 153.

Bowman, D., J.M. Londono and H. Sapriza (2015): U.S. unconventional monetary policy and transmission to emerging market economies, *Journal of International Money and Finance*, 55, 27–59.

Bruno, V. and H.S. Shin (2015): Capital flows and the risk-taking channel of monetary policy, *Journal of Monetary Economics*, 71, 119–132.

Bruno, V.G., B. Büyüksahin and M.A. Robe (2017): The Financialization of Food?, *American Journal of Agricultural Economics*, 99(1), 243.

Büyükşahin, B. and J.H. Harris (2011): Do Speculators Drive Crude Oil Futures Prices?, *The Energy Journal*, 0(2), 167–202.

Büyüksahin, B. and M.A. Robe (2014): Speculators, commodities and cross-market linkages, *Journal of International Money and Finance*, 42(0), 38–70.

Cesa-Bianchi, A., L.F. Cespedes and A. Rrebucci (2015): Global Liquidity, House Prices, and the Macroeconomy: Evidence from Advanced and Emerging Economies, *Journal of Money, Credit and Banking*, 47(S1), 301–335.

Cesa-Bianchi, A., G. Thwaites and A. Vicondoa (2016): Monetary Policy Transmission in an Open Economy: New Data and Evidence from the United Kingdom, Discussion Papers No. 1612, Centre for Macroeconomics (CFM).

Chen, Q., A. Filardo, D. He and F. Zhu (2016): Financial crisis, US unconventional monetary policy and international spillovers, *Journal of International Money and Finance*, 67, 62 – 81.

Chen, R., G.M. Milesi-Ferretti and T. Tressel (2013): External imbalances in the eurozone, *Economic Policy*, 28(73), 101–142.

Cheng, I.H., A. Kirilenko and W. Xiong (2015): Convective Risk Flows in Commodity Futures Markets, *Review of Finance*, 19(5), 1733–1781.

Cheng, I.H. and W. Xiong (2013): The Financialization of Commodity Markets, NBER Working Papers No. 19642, National Bureau of Economic Research, Inc.

Chinn, M.D. and H. Ito (2006): What matters for financial development? Capital controls, institutions, and interactions, *Journal of Development Economics*, 81(1), 163–192.

Chow, G.C. and A. Lin (1971): Best Linear Unbiased Interpolation, Distribution, and Extrapolation of Time Series by Related Series, *The Review of Economics and Statistics*, 53(4), 372–75.

Christensen, J.H. and G.D. Rudebusch (2012): The Response of Interest Rates to US and UK Quantitative Easing, *The Economic Journal*, 122(564), F385–F414.

Christensen, J.H.E. and J.M. Gillan (2012): Do Fed TIPS purchases affect market liquidity?, FRBSF Economic letter No. 2012-07, Federal Reserve Board.

Christiano, L.J., M. Eichenbaum and C.L. Evans (1999): Monetary policy shocks: What have we learned and to what end?, in: J. B. Taylor and M. Woodford (Eds.), *Handbook of Macroeconomics*: Elsevier, Vol. 1, Part A, Chapter 2, pp. 65–148.

Ciccarelli, M., A. Maddaloni and J.L. Peydró (2013): Heterogeneous transmission mechanism: monetary policy and financial fragility in the eurozone, *Economic Policy*, 28(75), 459–512.

Coibion, O. and Y. Gorodnichenko (2015): Is the Phillips Curve Alive and Well after All? Inflation Expectations and the Missing Disinflation, *American Economic Journal: Macroeconomics*, 7(1), 197–232.

Comunale, M. and J. Hessel (2014): Current account imbalances in the Euro area: Competitiveness or financial cycle?, DNB Working Papers No. 443, Netherlands Central Bank, Research Department.

Constancio, V. (2016): International headwinds and the effectiveness of Monetary Policy, Speech at the 25th Annual Hyman P. Minsky Conference on the State of the US and World Economies at the Levy Economics Institute of Bard College, European Central Bank.

Curcuru, S.E., T. Dvorak and F.E. Warnock (2010): Decomposing the U.S. external returns differential, *Journal of International Economics*, 80(1), 22–32.

Dahlhaus, T. and G. Vasishtha (2014): The Impact of U.S. Monetary Policy Normalization on Capital Flows to Emerging-Market Economies, Working Paper No. 2014-53, Bank of Canada.

D'Amico, S., D.H. Kim and M. Wei (2014): Tips from TIPS: the informational content of Treasury Inflation-Protected Security prices, Finance and Economics Discussion Series No. 2014-24, Board of Governors of the Federal Reserve System.

Darby, J. and J. Melitz (2008): Social spending and automatic stabilizers in the OECD, *Economic Policy*, 23(56), 716–756.

Davig, T. and E.M. Leeper (2011): Monetary–fiscal policy interactions and fiscal stimulus, *European Economic Review*, 55(2), 211–227.

de Wind, J. (2014): Time variation in the dynamic effects of unanticipated changes in tax policy, CPB Discussion Paper No. 271, CPB Netherlands Bureau for Economic Policy Analysis.

Dees, S., F.d. Mauro, M.H. Pesaran and L.V. Smith (2007): Exploring the international linkages of the euro area: a global VAR analysis, *Journal of Applied Econometrics*, 22(1), 1–38.

den Haan, W. (Ed.) (2013): *Forward guidance: perspectives from central bankers, scholars and market participants*: Vox eBook.

Department of the Treasury, Federal Reserve Bank of New York and Board of Governors of the Federal Reserve System (2016): U.S. Portfolio Holdings of Foreign Securities as of December 31, 2014, Report, U.S. Department of the Treasury.

Draghi, M. (2015a): Introductory statement to the press conference (with Q&A), Q&A, European Central Bank.

Draghi, M. (2015b): The ECB's recent monetary policy measures: Effectiveness and challenges : a speech at the 2015 Michel Camdessus Central Banking Lecture, International Monetary Fund, Washington, D.C., May 14, 2015, Speech, European Central Bank.

Draghi, M. (2016): Introductory statement to the press conference (with Q&A), Q&A, European Central Bank.

ECB (2015): Economic Bulletin Issue 6 - 2015, Economic Bulletin, European Central Bank.

Ederer, S., C. Heumeser and C. Staritz (2013): The role of fundamentals and financialisation in recent commodity price developments - an empirical analysis for wheat, coffee, cotton, and oil, Working Paper No. 42, Austrian Foundation for Development Research (ÖFSE).

Ehrmann, M. (2015): Targeting Inflation from Below: How Do Inflation Expectations Behave?, *International Journal of Central Banking*, 11(4), 213–249.

Ehrmann, M. and M. Fratzscher (2005): Equal Size, Equal Role? Interest Rate Interdependence Between the Euro Area and the United States, *Economic Journal*, 115(506), 928–948.

Ehrmann, M., M. Fratzscher and R. Rigobon (2011): Stocks, bonds, money markets and exchange rates: measuring international financial transmission, *Journal of Applied Econometrics*, 26(6), 948–974.

Eichengreen, B. and P. Gupta (2015): Tapering talk: The impact of expectations of reduced Federal Reserve security purchases on emerging markets, *Emerging Markets Review*, 25(C), 1–15.

Eickmeier, S. and T. Ng (2015): How do US credit supply shocks propagate internationally? A GVAR approach, *European Economic Review*, 74, 128–145.

European Commission (2015): Assessment of the 2015 Stability Programmes for various countries, Assessment of the 2015 Stability Programmes.

Fattouh, B., L. Kilian and L. Mahadeva (2013): The Role of Speculation in Oil Markets: What Have We Learned So Far?, *Energy Journal*, 0(3).

Forbes, K.J. and F.E. Warnock (2012): Capital flow waves: Surges, stops, flight, and retrenchment, *Journal of International Economics*, 88(2), 235–251.

Fragetta, M. and T. Kirsanova (2010): Strategic monetary and fiscal policy interactions: An empirical investigation, *European Economic Review*, 54(7), 855–879.

Fratzscher, M. (2012): Capital flows, push versus pull factors and the global financial crisis, *Journal of International Economics*, 88(2), 341–356.

Fratzscher, M., M. Lo Duca and R. Straub (2016a): ECB Unconventional Monetary Policy: Market Impact and International Spillovers, *IMF Economic Review*, 64(1), 36–74.

Fratzscher, M., M. Lo Duca and R. Straub (2016b): On the international spillovers of US Quantitative Easing, *Economic Journal*, Accepted Author Manuscript, available under: doi:10.1111/ecoj.12435.

Fry, R. and A. Pagan (2007): Some Issues in Using Sign Restrictions for Identifying Structural VARs, NCER Working Paper Series No. 14, National Centre for Econometric Research.

Gagnon, J., M. Raskin, J. Remache and B. Sack (2011): The financial market effects of the Federal Reserve's large-scale asset purchases, *International Journal of Central Banking*, 7(1), 3–43.

Galí, J. (2015): *Monetary Policy, Inflation, and the Business Cycle: An Introduction to the New Keynesian Framework and Its Applications Second edition*. Number 10495 in Economics Books. Princeton University Press.

Galati, G., S. Poelhekke and C. Zhou (2011): Did the Crisis Affect Inflation Expectations?, *International Journal of Central Banking*, 7(1), 167–207.

Gambacorta, L., B. Hofmann and G. Peersman (2014): The Effectiveness of Unconventional Monetary Policy at the Zero Lower Bound: A Cross-Country Analysis, *Journal of Money, Credit and Banking*, 46(4), 615–642.

Georgiadis, G. (2015): Examining asymmetries in the transmission of monetary policy in the euro area: Evidence from a mixed cross-section global VAR model, *European Economic Review*, 75, 195–215.

Georgiadis, G. (2016): Determinants of global spillovers from US monetary policy, *Journal of International Money and Finance*, 67, 41 – 61.

Gertler, M. and P. Karadi (2015): Monetary Policy Surprises, Credit Costs, and Economic Activity, *American Economic Journal: Macroeconomics*, 7(1), 44–76.

Gilbert, C.L. and S. Pfuderer (2014): The Role of Index Trading in Price Formation in the Grains and Oilseeds Markets, *Journal of Agricultural Economics*, 65(2), 303–322.

Girouard, N. and C. André (2006): Measuring cyclically-adjusted budget balances for OECD countries, Economics Department Working Paper No. 434, Organisation for Economic Co-operation and Development.

Goldstein, I. and L. Yang (2015): Information Diversity and Complementarities in Trading and Information Acquisition, *The Journal of Finance*, 70(4), 1723–1765.

Goldstein, I. and L. Yang (2016): Commodity Financialization: Risk Sharing and Price Discovery in Commodity Futures Markets, AFA 2016 Annual Meeting; Rotman School of Management Working Paper No. 2555996, The Wharton School Research Paper.

Gürkaynak, R.S., A. Levin and E. Swanson (2010b): Does Inflation Targeting Anchor Long-Run Inflation Expectations? Evidence from the U.S., UK, and Sweden, *Journal of the European Economic Association*, 8(6), 1208–1242.

Gürkaynak, R.S., B. Sack and E. Swanson (2005a): Do actions speak louder than words? The response of asset prices to monetary policy actions and statements, *International Journal of Central Banking*, 1(1), 55–93.

Gürkaynak, R.S., B. Sack and E. Swanson (2005b): The Sensitivity of Long-Term Interest Rates to Economic News: Evidence and Implications for Macroeconomic Models, *American Economic Review*, 95(1), 425–436.

Gürkaynak, R.S., B. Sack and J.H. Wright (2010a): The TIPS Yield Curve and Inflation Compensation, *American Economic Journal: Macroeconomics*, 2(1), 70–92.

Grossman, S.J. and J.E. Stiglitz (1980): On the impossibility of informationally efficient markets, *American Economic Review*, 393–408.

Hamilton, J.D. and J.C. Wu (2012): The effectiveness of alternative monetary policy tools in a zero lower bound environment, *Journal of Money, Credit and Banking*, 44(s1), 3–46.

Hamilton, J.D. and J.C. Wu (2014): Risk premia in crude oil futures prices, *Journal of International Money and Finance*, 42, 9–37.

Hamilton, J.D. and J.C. Wu (2015): Effects of Index-Fund Investing on Commodity Futures Prices, *International Economic Review*, 56(1), 187–205.

Hanlon, M., E.L. Maydew and J.R. Thornock (2015): Taking the Long Way Home: U.S. Tax Evasion and Offshore Investments in U.S. Equity and Debt Markets, *Journal of Finance*, 70(1), 257–287.

He, Z. and A. Krishnamurthy (2013): Intermediary Asset Pricing, *American Economic Review*, 103(2), 732–770.

Hellwig, M.F. (1980): On the aggregation of information in competitive markets, *Journal of Economic Theory*, 22(3), 477–498.

Henderson, B.J., N.D. Pearson and L. Wang (2015): New Evidence on the Financialization of Commodity Markets, *Review of Financial Studies*, 28(5), 1285–1311.

Herwartz, H. and H. Lütkepohl (2014): Structural vector autoregressions with Markov switching: Combining conventional with statistical identification of shocks, *Journal of Econometrics*, 183(1), 104–116.

Heumesser, C. and C. Staritz (2013): Financialisation and the microstructure of commodity markets: A qualitative investigation of trading strategies of financial investors and commercial traders, Working Paper No. 44, Austrian Foundation for Development Research (ÖFSE).

Hördahl, P. and O. Tristani (2012): Inflation Risk Premia In The Term Structure Of Interest Rates, *Journal of the European Economic Association*, 10(3), 634–657.

Hördahl, P. and O. Tristani (2014): Inflation Risk Premia in the Euro Area and the United States, *International Journal of Central Banking*, 10(3), 1–47.

International Monetary Fund (2013): Global Impact and Challenges of Unconventional Monetary Policies - Background Paper, IMF Policy Paper, October.

Irwin, S.H. and D.R. Sanders (2012): Testing the Masters Hypothesis in commodity futures markets, *Energy Economics*, 34(1), 256–269.

Jochmann, M., G. Koop and S.M. Potter (2010): Modeling the dynamics of inflation compensation, *Journal of Empirical Finance*, 17(1), 157–167.

Joyce, M., D. Miles, A. Scott and D. Vayanos (2012): Quantitative Easing and Unconventional Monetary Policy - An Introduction, *Economic Journal*, 122(564), F271–F288.

Kang, J.S. and J.C. Shambaugh (2016): The rise and fall of European current account deficits, *Economic Policy*, 31(85), 153–199.

Kang, W., K.G. Rouwenhorst and K. Tang (2014): The role of hedgers and speculators in liquidity provision to commodity futures markets, Yale International Center for Finance Working Paper No. 14-24.

Kapetanios, G., H. Mumtaz, I. Stevens and K. Theodoridis (2012): Assessing the Economy-wide Effects of Quantitative Easing, *Economic Journal*, 122(564), F316–F347.

Kaufmann, D., A. Kraay and M. Mastruzzi (2010): The worldwide governance indicators : methodology and analytical issues, Policy Research Working Paper Series No. 5430, The World Bank.

Keynes, J. (1930): *A Treatise on Money*, Vol. 2, London: Macmillan.

Kilian, L. and H. Lütkepohl (2017): *Structural Vector Autoregressive Analysis*, 1^{st} Edition: Cambridge University Press, forthcoming.

Kilian, L. and D.P. Murphy (2014): The Role of Inventories and Speculative Trading in the Global Market for Crude Oil, *Journal of Applied Econometrics*, 29(3), 454–478.

Klein, M.W. and J.C. Shambaugh (2008): The dynamics of exchange rate regimes: Fixes, floats, and flips, *Journal of International Economics*, 75(1), 70–92.

Klein, M.W. and J.C. Shambaugh (2015): Rounding the Corners of the Policy Trilemma: Sources of Monetary Policy Autonomy, *American Economic Journal: Macroeconomics*, 7(4), 33–66.

Koepke, R. (2014): Fed Policy Expectations and Portfolio Flows to Emerging Markets, MPRA Paper No. 63519, University Library of Munich, Germany.

Krishnamurthy, A. and A. Vissing-Jorgensen (2011): The Effects of Quantitative Easing on Interest Rates: Channels and Implications for Policy, NBER Working Papers No. 17555, National Bureau of Economic Research, Inc.

Kuttner, K.N. (2001): Monetary policy surprises and interest rates: Evidence from the Fed funds futures market, *Journal of Monetary Economics*, 47(3), 523–544.

Lanne, M. and H. Lütkepohl (2008): Identifying Monetary Policy Shocks via Changes in Volatility, *Journal of Money, Credit and Banking*, 40(6), 1131–1149.

Leeper, E.M. (1991): Equilibria under activeand passive monetary and fiscal policies, *Journal of Monetary Economics*, 27(1), 129–147.

Liikanen, E. (2015): Independence of monetary policy and the banking union, Speech at the Lamfalussy Lecture Conference (the Central Bank of Hungary), Budapest, Governor of the Bank of Finland.

Lim, J.J., S. Mohapatra and M. Stocker (2014): Tinker, taper, QE, bye? The effect of quantitative easing on financial flows to developing countries, Policy Research Working Paper Series No. 6820, The World Bank.

Lo Duca, M., G. Nicoletti and A. Vidal Martínez (2016): Global corporate bond issuance: What role for US quantitative easing?, *Journal of International Money and Finance*, 60(C), 114–150.

Lütkepohl, H. and A. Netšunajev (2014): Disentangling demand and supply shocks in the crude oil market: How to check sign restrictions in structural VARs, *Journal of Applied Econometrics*, 29(3), 479–496.

Masters, M.W. and A.K. White (2008): The accidental Hunt brothers: How institutional investors are driving up food and energy prices, The accidental Hunt Brothers Blog, special report postet on July, 31.

Mayer, J. (2009): The Growing Interdependence Between Financial And Commodity Markets, UNCTAD Discussion Papers No. 195, United Nations Conference on Trade and Development.

Mehrotra, A. and J. Yetman (2014): Decaying expectations: what inflation forecasts tell us about the anchoring of inflation expectations, BIS Working Papers No. 464, Bank for International Settlements.

Mertens, K. and M.O. Ravn (2013): The dynamic effects of personal and corporate income tax changes in the United States, *American Economic Review*, 103(4), 1212–1247.

Miranda-Agrippino, S. and H. Rey (2015): World Asset Markets and the Global Financial Cycle, NBER Working Papers No. 21722, National Bureau of Economic Research, Inc.

Moore, J., S. Nam, M. Suh and A. Tepper (2013): Estimating the Impacts of U.S. LSAPs on Emerging Market Economies' Local Currency Bond Markets, Federal Reserve Bank of New York Staff Report No. 595, Federal Reserve Bank of New York.

Mou, Y. (2011). *Limits to Arbitrage and Commodity Index Investment.* Ph. D. thesis, Columbia School.

Mountford, A. and H. Uhlig (2009): What are the effects of fiscal policy shocks?, *Journal of Applied Econometrics*, 24(6), 960–992.

Nagel, S. (2016): Long-Run Inflation Uncertainty, *International Journal of Central Banking*, 12(3), 207 – 217.

Nautz, D., A. Netsunajev and T. Strohsal (2016): The Anchoring of Inflation Expectations in the Short and in the Long Run, SFB 649 Discussion Papers No. 2016-015, Sonderforschungsbereich 649, Humboldt University, Berlin, Germany.

Nautz, D., L. Pagenhardt and T. Strohsal (2017): The (De-)Anchoring of Inflation Expectations: New Evidence from the Euro Area, *North American Journal of Economics and Finance* (40), 103–115.

Nautz, D. and T. Strohsal (2015): Are US inflation expectations re-anchored?, *Economics Letters*, 127(C), 6–9.

Netsunajev, A. and L. Winkelmann (2016): International dynamics of inflation expectations, SFB 649 Discussion Papers No. 2016-019, Sonderforschungsbereich 649, Humboldt University, Berlin, Germany.

Nier, E., T.S. Sedik and T. Mondino (2014): Gross Private Capital Flows to Emerging Markets: Can the Global Financial Cycle Be Tamed?, IMF Working Papers No. 14/196, International Monetary Fund.

Obstfeld, M., J.C. Shambaugh and A.M. Taylor (2005): The Trilemma in History: Tradeoffs Among Exchange Rates, Monetary Policies, and Capital Mobility, *The Review of Economics and Statistics*, 87(3), 423–438.

Olea, J.L.M., J.H. Stock and M.W. Watson (2013): Inference in Structural VARs with External Instruments, 3rd VALE-EPGE Global Economic Conference Business Cycles, May 9-10, 2013.

Orphanides, A. (2016): Fiscal Implications of Central Bank Balance Sheet Policies, Discussion Paper No. 11383, Centre for Economic Policy Research.

Passari, E. and H. Rey (2015): Financial Flows and the International Monetary System, *Economic Journal*, 125(584), 675–698.

Peersman, G. (2011): Macroeconomic Effects of Unconventional Monetary Policy in the Euro Area, CEPR Discussion Papers No. 8348, C.E.P.R. Discussion Papers.

Rai, V. and L. Suchanek (2014): The Effect of the Federal Reserve‘s Tapering Announcements on Emerging Markets, Working Paper No. 2014-50, Bank of Canada.

Reinhart, C.M. and K.S. Rogoff (2004): The Modern History of Exchange Rate Arrangements: A Reinterpretation, *Quarterly Journal of Economics*, 119(1), 1–48.

Rey, H. (2013): Dilemma not Trilemma: the Global Financial Cycle and Monetary Policy Independence, Proceedings of the Economic Policy Symposium at Jackson Hole, Federal Reserve Bank of Kansas City.

Rey, H. (2016): International Channels of Transmission of Monetary Policy and the Mundellian Trilemma, *IMF Economic Review*, 64(1), 6–35.

Rigobon, R. (2003): Identification through heteroskedasticity, *Review of Economics and Statistics*, 85(4), 777–792.

Rogers, J.H., C. Scotti and J.H. Wright (2014): Evaluating Asset-Market Effects of Unconventional Monetary Policy: A Cross-Country Comparison, International Finance Discussion Papers No. 1101, Board of Governors of the Federal Reserve System.

Rogers, J.H., C. Scotti and J.H. Wright (2016): Unconventional Monetary Policy and International Risk Premia, International Finance Discussion Papers No. 1172, Board of Governors of the Federal Reserve System.

Rossi, B. and S. Zubairy (2011): What is the importance of monetary and fiscal shocks in explaining US macroeconomic fluctuations?, *Journal of Money, Credit and Banking*, 43(6), 1247–1270.

Rouwenhorst, K.G. (1998): International Momentum Strategies, *Journal of Finance*, 53(1), 267–284.

Schmidt, C., P. Bofinger, I. Schnabel, L. Feld and V. Wieland (2015): Annual Economic Report, Annual Economic Report No. 2015/16, German council of economic experts.

Sentana, E. and G. Fiorentini (2001): Identification, estimation and testing of conditionally heteroskedastic factor models, *Journal of Econometrics*, 102(2), 143–164.

Shin, H.S. (2013): The second phase of global liquidity and its impact on emerging economies, in: *Remarks at 2013 Federal Reserve Bank of San Francisco Asia Economic Policy Conference*.

Shleifer, A. and L.H. Summers (1990): The noise trader approach to finance, *Journal of Economic Perspectives*, 19–33.

Shleifer, A. and R.W. Vishny (1997): The Limits of Arbitrage, *The Journal of Finance*, 52(1), 35–55.

Sims, C.A. (1980): Macroeconomics and Reality, *Econometrica*, 48(1), 1–48.

Singleton, K.J. (2014): Investor Flows and the 2008 Boom/Bust in Oil Prices, *Management Science*, 60(2), 300–318.

Sockin, M. and W. Xiong (2015): Informational Frictions and Commodity Markets, *Journal of Finance*, 70(5), 2063–2098.

Speck, C. (2016): Inflation anchoring in the euro area, Discussion Papers No. 04/2016, Deutsche Bundesbank, Research Centre.

Stock, J.H. and M.W. Watson (2012): Disentangling the Channels of the 2007-09 Recession, *Brookings Papers on Economic Activity*, 44(1 (Spring), 81–156.

Stock, J.H., J.H. Wright and M. Yogo (2002): A Survey of Weak Instruments and Weak Identification in Generalized Method of Moments, *Journal of Business & Economic Statistics*, 20(4), 518–529.

Stoll, H.R. and R.E. Whaley (2010): Commodity index investing and commodity futures prices, *Journal of Applied Finance* (20), 7–46.

Strohsal, T., R. Melnick and D. Nautz (2016): The time-varying degree of inflation expectations anchoring, *Journal of Macroeconomics*, 48, 62 – 71.

Strohsal, T. and L. Winkelmann (2015): Assessing the anchoring of inflation expectations, *Journal of International Money and Finance*, 50, 33 – 48.

Tang, K. and W. Xiong (2010): Index Investment and Financialization of Commodities, NBER Working Papers No. 16385, National Bureau of Economic Research, Inc.

Tillmann, P. (2016): Unconventional monetary policy and the spillovers to emerging markets, *Journal of International Money and Finance*, 66, 136 – 156.

Traum, N. and S.C.S. Yang (2011): Monetary and fiscal policy interactions in the post-war US, *European Economic Review*, 55(1), 140–164.

Unger, R. (2017): Asymmetric credit growth and current account imbalances in the euro area, *Journal of International Money and Finance*, forthcoming.

Van den Noord, P. (2000): The size and role of automatic fiscal stabilizers in the 1990s and beyond, Economics Department Working Paper No. 230, Organisation for Economic Co-operation and Development.

Wang, Y., C. Wu and L. Yang (2014): Oil price shocks and agricultural commodity prices, *Energy Economics*, 44(0), 22 – 35.

Weale, M. and T. Wieladek (2016): What are the macroeconomic effects of asset purchases?, *Journal of Monetary Economics*, 79(C), 81–93.

Weidmann, J. and K. Knot (2015): Mario Draghi hits back at QE hawks, Article, Financial Times.

Wright, J.H. (2012): What does Monetary Policy do to Long-term Interest Rates at the Zero Lower Bound?, *Economic Journal*, 122(564), F447–F466.

Wu, J.C. and F.D. Xia (2016): Measuring the Macroeconomic Impact of Monetary Policy at the Zero Lower Bound, *Journal of Money, Credit and Banking*, 48(2-3), 253–291.

Wyplosz, C. (2013): The Eurozone Crisis and the Competitiveness Legend, *Asian Economic Papers*, 12(3), 63–81.

Yellen, J.L. (2015): Inflation Dynamics and Monetary Policy : A speech at the Philip Gamble Memorial Lecture, University of Massachusetts, Amherst, Amherst, Massachusetts, September 24, 2015, Speech No. 863, Board of Governors of the Federal Reserve System (U.S.).

Ehrenwörtliche Erklärung

Ich habe die vorgelegte Dissertation selbst verfasst und dabei nur die von mir angegebenen Quellen und Hilfsmittel benutzt. Alle Textstellen, die wörtlich oder sinngemäß aus veröffentlichten oder nicht veröffentlichten Schriften entnommen sind, sowie alle Angaben, die auf mündlichen Auskünften beruhen, sind als solche kenntlich gemacht.

Michael Hachula
Berlin, den 31. März 2017

Liste verwendeter Hilfsmittel

- Matlab 8.1.0.604 (R2013a)
  - Optimization Toolbox
  - Financial Toolbox
  - Statistics Toolbox
- Eviews 8.0
- Stata 13
- Microsoft Excel
- LaTeX
- Siehe auch Literatur- und Quellenangaben