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To my parents.

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Declaration of Co-Authorship and Publications

This dissertation consists of four research papers. Two papers were written in collaboration with one co-author and in one paper I had two co-authors. On paper is single authored. My contribution in conception, implementation and drafting can be summarized as follows:

1. **The Short-Run Effect of Monetary Policy on Credit Risk: An Analysis of the Euro Area**

   by Chi Hyun Kim and Lars Other

   Contribution: 50 percent

   An early version of this chapter was published as a DIW Berlin Working Paper:

   Kim, C. H. and Other, L. (2019). The Short-Run Effect of Monetary Policy Shocks on Credit Risk: An Analysis of the Euro Area. DIW Berlin Discussion Paper No.1781 (2019).

2. **The Term Structure of Redenomination Risk**

   by Christian Bayer, Chi Hyun Kim, and Alexander Kriwoluzky

   Contribution: 33.3 percent

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   Bayer, C., Kim C. and Kriwoluzky, A. (2018). The Term Structure of Redenomination Risk (May 2018). CEPR Discussion Paper No. DP12965.

3. **The Effect of Monetary Policy on Stock Market Investment Decisions: The Role of Gender and Marital Status**

   by Caterina Forti Grazzini und Chi Hyun Kim

   Contribution: 50 percent

   An early version of this chapter was published as a DIW Berlin Working Paper:

   Forti Grazzini, C. and C. H. Kim (2020). Is Monetary Policy Gender Neutral? Evidence from the Stock Market. DIW Berlin Discussion Paper No. 1841 (2020).

4. **Optimism Gone Bad? Persistent Effects of Traumatic Experiences on Households' Investment Decisions**

   by Chi Hyun Kim

   Contribution: 100 percent

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List of Abbreviations

AIF	Aktien-Informations-Forum
ARDL	Autoregressive Distributed Lag
BVerfG	Bundesverfassungsgericht
CAC	Collective Action Clause
CDS	Credit Default Swap
CJEU	European Court of Justice
CPI	Consumer Price Index
CSPP	Corporate Sector Purchase Programme
ECB	European Central Bank
e.g.	exempli gratia
EUR	Euro
FAZ	Frankfurter Allgemeine Zeitung
FE	Fixed Effects
FED	Federal Reserve
FOMC	Federal Open Market Commitee
FRG	Federal Republic of Germany
G7	Group of Seven
GDR	German Democratic Republic
GFC	Global Financial Crisis
GIPS	Greece, Ireland, Portugal, and Spain
FAVAR	Factor Augmented VAR
HFI	High Frequency Identification
HH	Household
i.e.	id est
IP	Industrial Production
IPO	Initial Public Offering
ISDA	International Swaps and Derivatives Association
LSAP	Large Scale Asset Purchases
MHH	Male-Headed Household
MM	Modified-Modified
MP	Monetary Policy
IRA	Individual Retirement Account
IRF	Impulse Response Function
LR	Likelihood Ratio
LTRO	Long-Term Refinancing Operation

OLS	Ordinary Least Squares
OMT	Outright Monetary Transactions
PSID	Panel Study of Income Dynamics
QE	Quantitative Easing
QOQ	Quarter on Quarter
SA	Seasonal Adjusted
S&P 500	Standard&Poor 500
SD	Standard Deviation
SFHH	Single Female-Headed Household
SHIW	Bank of Italy Survey on Household Income and Wealth
SMHH	Single Male-Headed Household
SMP	Stock Market Participation
SMP-1	First Phase of the Securities Market Program
SMP-2	Second Phase of the Securities Market Program
SOEP	Socio-Economic Panel
SVAR	Structural Vector Autoregression
TLTRO	Targeted Longer-Term Refinancing Operation
US	United States of America
USD	US Dollar
VAR	Vector Autoregression
WLS	Weighted Least Squares
ZLB	Zero Lower Bound

Summary

This dissertation comprises four chapters, each using micro-level data to empirically analyze dynamics in the financial markets. The first chapter, joint with Lars Other, studies the credit channel of monetary policy transmission mechanism in the euro area. We construct credit risk measures separately for France, Germany, Italy, and Spain using daily data on corporate bond yields of non-financial corporations. With these credit risk measures, we examine whether ECB monetary policy significantly affects credit conditions of the euro area countries. Monetary policy shocks are identified using the high frequency identification method. Our results show that, from 2000 to 2015, expansionary monetary policy shocks result in a short-run increase in the credit risk of non-financial corporations of all four countries. This dysfunctionality of the credit channel is driven by the crisis-dominated post-2009 period. During this time, market participants may have interpreted expansionary monetary policy shocks as a signal of worsening economic prospects. In order to understand the drivers of this dysfunctionality, we distinguish policy shocks aiming at short-run and long-run expectations of market participants, i.e. target and path shocks. Our analysis shows that, at least for the crisis countries, the adverse effect disappears when the ECB targets long-run rather than short-run expectations.

The second chapter, joint with Christian Bayer and Alexander Kriwoluzky, assesses redenomination risk in the euro area. As a first step, we estimate two distinct daily default-risk-free yield curves for French, German, and Italian bonds that can be redenominated in case of an exit and for those bonds that cannot, respectively. Thereafter, we extract the compensation for redenomination risk from the yield spreads between these two types of bonds. Redenomination risk primarily shows up at the short end of the yield curve. At the height of the euro crisis, spreads between first-year yields were close to 7% for Italy and up to -1% for Germany. The ECB's interventions, which were designed to reduce the risk of a breakup, did so for Italy, but increased it for France and Germany.

The third chapter, joint with Caterina Forti Grazzini, studies the role of gender and marital status on how monetary policy affects the investment decisions of households. We use the Panel Study of Income Dynamics (PSID) household survey data from 2001-2017 to investigate whether monetary policy has heterogeneous effects on stock market participation decisions across (i) single female-headed households and (ii) male-headed households (both single or married) in the US. The Federal Reserve's monetary policy shocks are identified using the high frequency identification method. On the one hand, we show that a monetary policy shock affects the stock market entry decisions of single female-headed households, while it neither impacts single- nor married male-headed households. On the other hand, monetary policy does not have a significantly different effect on either exit decisions or stock market investment rebalancing choices.

The fourth chapter explores whether memories of traumatic stock market crashes permanently affect the investment decisions of households. I address this question by focusing on the Initial Public Offerings (IPOs) of Deutsche Telekom during 1996-2000. This event provides an optimal base to address this question, as the dramatic crash of Telekom share prices shortly after the IPOs, followed by revelation of corruption scandals, traumatized the German public. Using Socio-Economic Panel (SOEP) household survey data, I show that having experienced this event leads to persistently lower stock market participation in the future. This effect is greater for households that had directly invested in Telekom shares. Finally, I show that such traumatic experiences with investment decisions have intergenerational consequences, significantly affecting how the next generation invests in the financial market.

Zusammenfassung

Die vorliegende Dissertation umfasst vier Kapitel. Alle beinhalten empirische Analysen von Finanzmarktdynamiken mit Hilfe mikroökonomischer Daten. Im Mittelpunkt des ersten Kapitels (Koautor: Lars Other) steht die Untersuchung des Kreditkanals als geldpolitischer Transmissionsmechanismus im Euroraum. Hierfür ermitteln wir anhand täglicher Renditedaten für Anleihen nichtfinanzieller Unternehmen das jeweilige Kreditrisiko für Deutschland, Frankreich, Italien und Spanien. Anhand dieser Kreditrisikomaße analysieren wir, ob geldpolitische Schocks signifikante Effekte auf die Kreditbedingungen der vier Länder haben. Nicht antizipierte geldpolitische Interventionen der Europäischen Zentralbank (EZB) identifizieren wir mithilfe der sogenannten Hochfrequenzidentifikation (HFI). Unsere Ergebnisse zeigen, dass expansive geldpolitische Schocks im Zeitraum von 2000 bis 2015 das kurzfristige Kreditrisiko für nichtfinanzielle Unternehmen aller vier Länder erhöhten. Diese Dysfunktionalität des Kreditkanals wird vor allem durch die von Finanz- und Schuldenkrise geprägte Periode nach 2009 getrieben. Marktteilnehmer:innen könnten während dieser Zeit expansive Geldpolitik als Anzeichen für schlechtere wirtschaftliche Aussichten interpretiert haben. In unserer Analyse differenzieren wir zudem, ob sich geldpolitische Schocks auf die langfristigen oder kurzfristigen Erwartungen von Marktteilnehmer:innen auswirken. Dafür identifizieren wir zwei unterschiedliche Dimensionen der geldpolitischen Schockreihe: einen „target shock“ (kurzfristige Dimension) und einen „path shock“ (langfristige Dimension). Von dieser Analyse lernen wir, dass die EZB mit einer an langfristigen Erwartungen ausgerichteten Geldpolitik das Funktionieren des Kreditkanals, zumindest in den stark von der Finanz- und Schuldenkrise betroffenen Ländern, wiederherstellen kann.

Im zweiten Kapitel (Koautoren: Christian Bayer und Alexander Kriwoluzky) berechnen wir das Redenominierungsrisiko im Euroraum für den Zeitraum 2010-2014. Im ersten Schritt ermitteln wir anhand von Tagesdaten Zinsstrukturkurven für zwei unterschiedliche Klassen ausfallrisikofreier Anleihen für Deutschland, Frankreich und Italien. Im Falle eines Eurozonentaustritts gestattet die eine Anleiheklasse eine Währungsumstellung des Nennwerts und Cashflows, die andere schließt dies rechtlich aus. Die Differenz der Renditen zwischen den unterschiedlichen Anleiheklassen interpretieren wir als Redenominierungsrisiko. Unsere Analyse zeigt, dass Redenominierungsrisiko vor allem am kürzeren Ende der Zinsstrukturkurve eine Rolle spielt. Auf dem Höhepunkt der Eurokrise betrug die Renditedifferenz im ersten Jahr in Italien nahezu 7% und in Deutschland bis zu -1%. Die geldpolitischen Interventionen der EZB zur Stabilisierung des Euroraums konnten in Italien das Redenominierungsrisiko senken. Für Deutschland und Frankreich erhöhten sie jedoch das Risiko.

Im dritten Kapitel (Koautorin: Caterina Forti Grazzini) untersuchen wir den Einfluss geldpolitischer Maßnahmen auf individuelle Investitionsentscheidungen, insbesondere die Rolle von Geschlecht und Familienstand. Mit Umfragedaten aus der „Panel Study of Income Dynamics“ (PSID) für die Jahre 2001 bis 2017 analysieren wir, ob Geldpolitik die Entscheidung am Aktienmarkt zu investieren von (i) alleinstehenden Frauen und (ii) Männern (alleinstehend oder als Haushaltsvorstand) unterschiedlich stark beeinflusst. Mithilfe des HFI-Verfahrens identifizieren wir geldpolitische Schocks der US-Notenbank. Unsere Ergebnisse zeigen einerseits, dass die Entscheidung alleinstehender Frauen erstmalig am Aktienmarkt zu investieren signifikant von nicht antizipierten geldpolitischen Interventionen beeinflusst wird. Bei Männern, ganz gleich ob alleinstehend oder als Haushaltsvorstand, ist dieser Zusammenhang nicht festzustellen. In Hinblick auf Austritts- und Investitionsentscheidungen können wir keine signifikanten Unterschiede zwischen den Untersuchungsgruppen am Aktienmarkt ausmachen.

Das vierte Kapitel ist der Frage gewidmet, ob Investitionsentscheidungen von Haushalten dauerhaft durch negative Erinnerungen an aufsehenerregende Börsencrashs beeinflusst werden. Im Fokus steht hierbei der vielbeachtete Börsengang der Deutschen Telekom zwischen 1996 und 2000. Die spektakulären Kursverluste der Telekomaktie kurz nach Börsengang und die in diesem Zusammenhang öffentlich gewordenen personellen Skandale waren für die breite öffentlichkeit in Deutschland ein prägendes Erlebnis und eignen sich aufgrund dieser Eigenschaft besonders gut für meine Fragestellung. Anhand von Umfragedaten aus dem Sozio-oekonomischen Panel (SOEP) zeige ich, dass Haushalte, die diesen Börsengang miterlebt haben, sich in der Zukunft dauerhaft geringer an der Börse beteiligen. Für Haushalte, die zwischen 1996 und 2000 direkt in Aktien der Telekom investiert haben, ist dieser Effekt ausgeprägter. In meiner Analyse zeige ich außerdem, dass solch traumatische Investitionserfahrungen auch generationenübergreifend die Beteiligung am Aktienmarkt reduzieren können.

Introduction

The Global Financial Crisis of 2007 and the subsequent European sovereign debt crisis of 2009 were sharp reminders of the importance of a well-informed understanding of the origins and consequences of financial frictions on the real economy. Within this context, *micro-level data* experienced a boom in both theoretical and empirical economic research in the field of macro-finance, as granular data on financial markets and its participants enable a more rigorous analysis of potential sources of stress in the financial markets. Such micro-level insights are key to correctly identify and react to threats to the stability of the financial system that can spill over to the real economy.

The four independent chapters of my dissertation are the outcomes of my interest in contributing to this field, aiming for a better understanding of the dynamics and heterogeneity in the financial system. In particular, I utilize diverse micro-level data sets to exploit various dimensions of *frictions in the financial market*. In the first two chapters of my dissertation, I focus on euro area financial markets during the European sovereign debt crisis, analyzing how market participants evaluated different risk components across euro area securities. In doing so, I also examine whether ECB monetary policy successfully stabilized euro area financial markets during the crisis period. The latter two chapters focus on heterogeneous investment decisions of households in the financial markets. Here I exploit how different household groups invest in the stock market, depending on their differences in socio-economic characteristics and personal experiences. These chapters shed light on those groups within society that miss out on equity premia crucial for long-term wealth accumulation and, in missing out, has the potential to worsen economic inequality.

Summary of the chapters

The **first chapter**, "*The Short-Run Effect of Monetary Policy Shocks on Credit Risk: An Analysis of the Euro Area*," is joint work with Lars Other and investigates the functionality of the euro area credit channel from 2000 to 2015. The credit channel is important for the monetary policy transmission mechanism as central banks are able to directly influence the cash flows and balance sheet positions of corporations, thus achieving a stronger impact on private sector borrowing rates than just through changes in riskless interest rates. During the European sovereign debt crisis, when credit conditions had severely deteriorated in the private sector, a well-functioning credit channel was especially critical for the success of accommodative monetary policy actions. Yet, it is not clear whether the credit channel was well-functioning during the crisis, especially as ECB policy rates eventually reached the effective lower bound and, therefore, were no longer able to support the illiquid financial markets in a conventional manner.

Therefore, we empirically analyze the effects of ECB monetary policy on the borrowing conditions of non-financial corporations of the four largest euro area economies, France, Germany, Italy, and Spain. A separate analysis of these countries is crucial for addressing the potential heterogeneous effects of monetary policy, as not all euro area countries were equally affected by the crisis (Italy and Spain were more severely hit than France and Germany). Therefore, ECB monetary policy actions may have had different effects on the respective private-sector borrowing conditions. For each country, we construct credit spreads using daily micro-level data on yields of non-financial corporate bonds by following the method of Gilchrist and Mojon (2018). Daily monetary policy shocks in the euro area are identified by applying the high frequency identification (HFI) method of Gürkaynak et al. (2005).

During the crisis-dominated post-2009 period, an expansionary monetary policy shock leads to a short-run increase in the credit risk of non-financial corporations of all four countries. In order to understand the nature of this dysfunctionality in the credit channel, we distinguish between monetary policy shocks aiming at short-run and long-run expectations of market participants, i.e., target and path shocks. When the ECB targets long-run expectations rather than the short-run expectations, the adverse effect of monetary policy shocks on credit risk disappears for Italy and Spain, which were severely hit by the debt crisis, but does not for France and Germany. This sheds light on two issues. First, ECB monetary policy seems to have heterogeneous effects across the member countries: while investors in the Italian and Spanish corporate bond market seem to evaluate lower interest rates positively, investors in the French and German corporate bond market have a rather pessimistic view of the same event. We interpret this as an indication of how the signal of worsening economic prospects can dominate the positive effect of low interest rates for non-crisis countries since their benefit of lower interest rates is much lower as compared to crisis countries. Second, forward guidance appears to be an effective instrument for the ECB when conducting monetary policy during times of high economic uncertainty.

The **second chapter**, "*The Term Structure of Redenomination Risk*," joint work with Christian Bayer and Alexander Kriwoluzky, assesses redenomination risk in the euro area, which is the additional risk premia euro area countries must bear due to positive exit expectations among financial market participants. This risk was especially present during the European sovereign debt crisis, where exit scenarios of crisis countries such as Greece and Italy were prevalent (Draghi, 2012).

We estimate a term structure of the risk premia charged for redenomination risk for France, Germany, and Italy over the 2010-2014 period. To do so, we use daily micro-level financial data to estimate market expectations of a euro area breakup from differences in yield curves of securities, which are differentially affected by a country leaving the euro area. On the one hand, we have euro-denominated securities that will redenominate in case the country of the issuer exits the euro area. On the other hand, we have euro-denominated securities that will not redenominate even if the country of the issuer exits. The yield spreads between these two types of securities are interpreted as the compensation for redenomination risk.

Results are twofold. First, we show that redenomination risk primarily shows up at the short end of yield curves. At the height of the euro crisis, spreads between first-year yields were close to 7% for Italy and up to -1% for Germany. Second- and third-year redenomination risk were significantly smaller. Second, the ECB's interventions designed to reduce the risk of a breakup successfully did so for Italy, but increased it for France and Germany. After the announcements of major policy interventions by the ECB, such as the Securities Markets Programme and Outright Monetary Transactions, premia charged for Italian redenomination risk decreased significantly, while the premia rather increased for France and Germany.

The **third chapter**, "*The Effect of Monetary Policy on Stock Market Investment Decisions: The Role of Gender and Marital Status*," joint work with Caterina Forti Grazzini, examines the impact of gender and marital status on monetary policy-driven financial portfolio decisions. By doing so, we set a special focus on single female-headed households. The gender wealth gap is one of the most important economic gaps, with single female households being the poorest in society. Besides gender-specific differences in income, a well-established strand in the literature documents that women are more risk averse and less confident in their investment decisions, thus leading to less investment in risky financial assets. This implies that women may be missing out the high equity premia that is crucial for long-term wealth accumulation. During a prolonged period of low interest rates (with high asset prices), the gender wealth gap may even increase - an important aspect in the ongoing economic debate regarding the possible distributional effects of monetary policy (Yellen, 2016; Draghi, 2016).

We utilize micro-level household survey data of the Panel Study of Income Dynamics (PSID) to investigate whether monetary policy has a heterogeneous impact on stock market investment decisions of single-female headed households in comparison to both single male-headed households and married male-headed households in the USA. Monetary policy shocks are identified by adopting the HFI method of Nakamura and Steinsson (2018). Results can be summarized as the following. On the one hand, single female-headed households are less likely to enter the stock market after a contractionary monetary policy shock, while male-headed households' entry decisions (both single and married) are not affected by monetary policy. On the other hand, as soon as they participate in the stock market, the effect of monetary policy does not differ across female-headed households and male-headed households: both exit decisions and active rebalancing behavior (in terms of selling/buying stocks) are equally affected by a monetary policy shock. Using these insights, we conduct a simulation study and quantify the missed-out capital gains stemming from monetary policy-driven stock market non-participation of single female-headed households.

The **fourth chapter**, "*Optimism Gone Bad? Persistent Effects of Traumatic Experiences on Households' Investment Decisions*," examines whether memories of traumatic stock market crashes can permanently affect investment decisions of households. To do so, I investigate the Initial Public Offerings (IPOs) of the Deutsche Telekom during 1996-2000. This event provides an optimal base to address this question, as the dramatic crash of Telekom share prices shortly after the IPOs, followed by revelation of corruption scandals, provoked strong emotional turbulence among the German public and, thus, has the reputation of being "the last time Germans invested in stocks" (Handelsblatt, 2016). I argue that this high emotional turbulence imprinted the memories of this event on German households, in combination with the emotion that was generated along the way, thus permanently influencing the way how they invest in the stock market, even more than fifteen years after the event.

Utilizing micro-level household survey data of the Socio-Economic Panel (SOEP), I confirm that having experienced this event leads to persistently lower stock market participation in the future. In addition, this effect is intensified for households with direct investment in Telekom shares, because they exhibit a higher probability of emotional attachment. Finally, I also show that such traumatic experiences on investment decisions are transferable intergenerationally, thus significantly affecting how the next generation invests in the financial market.

Literature review

All four chapters cover very different fields in empirical macro-finance – from financial frictions and monetary policy to behavioral finance. However, they all have one theme in common, namely that they all attempt to shed light on the nature of heterogeneity in the financial markets. In this section, I provide a literature review and connect the different topics of my dissertation together into one big picture, which I see as my personal research agenda in the future.

There is a huge literature that utilizes micro-level data to explain different sources of frictions in the financial markets. One strand in this literature focuses on quantifying and analyzing credit risk in the financial markets. Great effort is made at improving existing credit risk measures by utilizing micro-level data on prices of individual securities (Gilchrist and Zakrajšek, 2012; Gilchrist and Mojon, 2018). Compared to other conventional indicators that are subject to maturity mismatch, micro-level information on individual securities allows for constructing a pure measure of credit risk that has high predictive power for future economic activity. Other studies focus on disentangling credit risk into distinct risk components to understand the drivers of high yield spreads during the Global Financial Crisis and the subsequent European sovereign debt crisis (Di Cesare et al., 2012; Ang and Longstaff, 2013; Krishnamurthy et al., 2018). In doing so, some studies focus on default risk (Aizenman et al., 2013; Arce et al., 2013; Cruces and Trebesch, 2013; Zettelmeyer et al., 2014; Trebesch and Zettelmeyer, 2018), while other studies examine the role of liquidity risk in the financial markets (Covitz and Downing, 2007; Beber et al., 2009; Schwarz, 2019). Additionally, for the euro area, the so-called "redenomination risk" was at the center of political debate, as deteriorating credit conditions of the member countries triggered fear among financial market participants that a sovereign default would inevitably lead to the exit of a euro area member (Choi et al., 2011; Clare and Schmidlin, 2014; Chamon et al., 2018; Kremens, 2018; De Santis, 2019). The **second chapter** of my dissertation contributes to this stream of literature by providing a term structure of redenomination risk of euro area countries during the European sovereign debt crisis.

In times of high financial stress, monetary policy plays a key role in stabilizing disruptive financial markets and mitigating negative spillover effects on the real economy. Further, exceptional times call for exceptional measures: in the aftermath of the Global Financial Crisis, central banks of major advanced economies sharply reduced interest rates, reaching the zero lower bound, expanded their liquidity facilities, and started large-scale asset purchase programs in an unprecedented manner. This unique monetary policy environment awakened a high demand for a better understanding of how such policy interventions transmit to the real economy. Within this context, micro-level data is very useful for examining different channels of monetary policy transmission mechanisms: household survey data are extensively used to analyze how heterogeneity in households' wealth portfolio and debt structures influences the way they respond to a monetary policy shock (Kaplan et al., 2014, 2018; Luetticke, 2018; Wong, 2019; Cloyne et al., 2020; Cumming and Hubert, 2020); micro-level financial data provide evidence on how monetary policy (both conventional and unconventional) affects credit conditions of market participants (Beckworth et al., 2010; Brand et al., 2010; Krishnamurthy and Vissing-Jorgensen, 2011; Cenesizoglu and Essid, 2012; Altavilla et al., 2016; Chakraborty et al., 2020; Bertsch et al., 2021). The **first chapter** of my dissertation contributes to this literature by exploiting different dimensions of ECB monetary policy shocks to shed light on the functionality of the credit channel of monetary policy transmission mechanism during the European sovereign debt crisis.

On the one hand, evidence in the literature shows that extra-loose monetary policies were successful at stabilizing global financial markets during the economic downturn. On the other hand, economic inequality in industrialized countries increased drastically and the public raised concerns that the long-enduring low interest rate environment is exacerbating this problem, since low interest rates might only benefit certain groups of households (Bivens, 2015). In particular, as expansionary monetary policy has positive effects on prices of financial assets, concerns were that only households that own financial wealth benefited from low interest rates, therefore increasing the wealth gap between households that invest in financial markets and those that do not. In the literature, several studies use household survey data to address this concern, while concentrating on different wealth- and/or income groups. Adam and Tzamourani (2016), Casiraghi et al. (2018), Ampudia et al. (2018), and Lenza and Slacalek (2018) show how changes in asset prices through unconventional monetary policy affect capital gains on households' financial portfolios, especially for the top end of the net wealth distribution; Forti Grazzini (2020) shows that loose monetary policy increased the risk appetite of households and, thus, their investment in financial assets, but only for the upper 25% of the income distribution. The empirical analysis of **chapter three** contributes to this debate by taking a different perspective for explaining the distributional impact of monetary policy, namely looking at household groups that differ in their heads' gender and marital status.

Insights from the debate on the distributional impacts of monetary policy show that during times of low interest rates, households that do not invest in stocks miss out equity premia, which can be crucial for their long-term wealth accumulation and, thus, financial well-being after retirement. Unfortunately, this applies to most households as their stock market participation is very low - a phenomenon well established in the literature as the “stock market participation puzzle” (Haliassos and Bertaut, 1995; Guiso et al., 2003). Great effort is made to understand the drivers of such behavior, for example through differences in their socio-economic characteristics such as age, gender, and income (Barber and Odean, 2001; Cocco et al., 2005; Beckmann and Menkhoff, 2008; Calvet et al., 2009; Christiansen et al., 2010), but also by using insights from behavioral and psychological factors such as risk aversion (Holt and Laury, 2002; Brunnermeier and Nagel, 2008; Eckel and Grossman, 2008; Guiso et al., 2018), social networks (Kaustia and Knüpfer, 2012; Heimer, 2016), and trust (Bohnet and Zeckhauser, 2004; Guiso et al., 2008; Georgarakos and Pasini, 2011; Pevzner et al., 2015). One key finding is that personal experiences play an important role on how households form their beliefs, which build a base for their economic decision making. Therefore, not only do households anticipate their life-time experiences when making investment decisions (Malmendier and Nagel, 2011; Ampudia and Ehrmann, 2017; Laudenbach et al., 2019b), but also when building expectations with regard to future inflations (Malmendier and Nagel, 2016; D’Acunto et al., 2019) and economic prospects (Kozlowski et al., 2019; Laudenbach et al., 2019a). The **fourth chapter** of my thesis contributes to this literature by shedding light on how persistent emotionally-attached stock market experiences can influence households’ decisions to invest in stocks in the future.

CHAPTER 1

The Short-Run Effect of Monetary Policy on Credit Risk: An Analysis of the Euro Area[1]

Chi Hyun Kim and Lars Other

This chapter investigates the credit channel of monetary policy in the euro area using daily monetary policy shock measures and credit risk measures in an autoregressive distributed lag model. From 2000 to 2015, we find that an expansionary monetary policy shock leads to a short-run increase in the credit risk of non-financial corporations based in France, Germany, Italy, and Spain. This dysfunctionality of the credit channel is driven by the crisis-dominated post-2009 period. During this time, market participants may have interpreted expansionary monetary policy shocks as a signal of worsening economic prospects. We further distinguish monetary policy shocks aiming at short-run and long-run expectations of market participants, i.e. target and path shocks, respectively. The adverse effect disappears for crisis countries such as Italy and Spain when the European Central Bank targets long-run rather than short-run expectations.

Keywords: credit channel, credit spreads, forward guidance, zero lower bound
JEL classication: C22, E44, E52, G12

[1]We thank the Halle Institute for Economic Research for providing us with the acccess to Datastream. Our special thanks also go to Jordi Gali, Oliver Holtemöller, Lutz Kilian, Alexander Kriwoluzky, Hans-Walter Lorenz, Tommaso Monacelli, Christian Pigorsch, Barbara Rossi, Maik Wolters, and participants of the DIW Macroeconomic Workshop in Berlin 2016, the SMYE 2017 in Halle, the Nordic Econometric Meeting in Tartu 2017, the EEA-ESEM in Lisbon 2017, and internal seminars for very helpful comments on earlier versions of the paper.

1.1 Introduction

During the European sovereign debt crisis, the importance of the *credit channel* for the ECB's monetary policy transmission mechanism has been highlighted in numerous speeches and remarks by Mario Draghi, President of the European Central Bank.[2] Through this channel, the ECB is able to influence directly the cash flows and the balance sheet positions of corporations, thus achieving a stronger impact on private sector borrowing rates than just through changes in riskless interest rates. Yet, it is not clear whether the credit channel functioned during the European sovereign debt crisis, as this period was characterized by important structural changes in the euro area financial markets. On the one hand, limited access to bank credit in the euro area resulted in increasing supply and demand for non-financial corporate bonds (De Fiore and Uhlig, 2015; Deutsche Bundesbank, 2017). This change in the nature of euro area corporate funding may have enhanced the demand-driven transmission mechanism of ECB monetary policy that works primarily through firm balance sheets (Ashcraft and Campello, 2007). On the other hand, as the ECB lowered its policy rates, eventually reaching the effective lower bound, it was no longer able to use conventional policy tools to further support the illiquid financial markets.

In this paper, we investigate the functionality of the credit channel in the euro area by analyzing the effect of monetary policy shocks on the borrowing conditions of non-financial corporations. We examine the credit channels of France, Germany, Italy, and Spain between January 2000 and November 2015. Our sample period enables us to analyze whether the efficacy of the credit channel changed during the Global Financial Crisis and its aftermath. Consequently, we divide the sample period into two distinct monetary policy regimes: (i) a period of "normal interest rates" and (ii) a period of low interest rates, characterized by prolonged expansionary monetary policy and the introduction of unconventional monetary policy operations. To investigate the effect of monetary policy on the credit risks, we use an autoregressive distributed lag model in combination with daily indicators of credit risk and high frequency measures of monetary policy shocks. We focus on the short-run dynamics since we want to capture the exogenous effect of monetary policy on fast-moving financial variables.

First, we construct an indicator of credit risk based on the spread between the borrowing rates of non-financial corporations and the riskless interest rate. This spread represents the credit risk assessment of market participants and, thus, reflects their expectations regarding the future economic activity. Bleaney et al. (2016) and Gilchrist and Mojon (2018) show, based on monthly data, that this approach yields a timely and reliable measure of borrowing conditions in the euro area. We extend this method to daily data and construct credit risk indicators of non-financial corporations of France, Germany, Italy, and Spain.

Second, we use the high frequency identification method of Gürkaynak et al. (2005) to identify two distinct dimensions of monetary policy shocks. As in the event study literature, we identify the surprise component of monetary policy actions by movements in the money market futures on the day of monetary policy announcements. By consid-

[2]See, *inter alia*, the hearing before the Plenary of the European Parliament in 2011 (Draghi, 2011).

ering the change in a sufficiently narrow time window, we can rule out other economic events that may have additionally influenced the futures rates. The high frequency identification method has an advantage over other conventional identification methods when using financial variables, e.g. recursive identification of SVAR, since it is able to account for the simultaneity problem between fast moving financial variables and policy shifts.

Since the prices of money market futures contracts are influenced by the expectations of investors regarding the future stance of monetary policy, we are able to capture the effect of monetary policy on the yield curve as a whole. Applying a factor model on a broad range of 3-month Euribor futures rates, we obtain a target shock and a path shock of monetary policy.[3] While the target shock exclusively represents the effect of current policy action on the short-end of the yield curve, the path shock represents the change in the expectations of market participants, which is induced by the information of monetary policy announcements beyond the change in the current policy rate.

Our dataset enables us to include various dimensions into our analysis. First, the long-time span of our dataset allows us to compare the effect of ECB monetary policy before and after the Global Financial Crisis. Indeed, if the tensions in the financial markets resulted in a dysfunctional transmission mechanism in the euro area, we would detect different effects of ECB monetary policy on the borrowing conditions of non-financial corporations. Second, we group the countries into two distinct categories. The first group consists of Italy and Spain, which were severely hit by the crisis. The second group comprises France and Germany, which were less hit by the crisis and, consequently, were not the target of ECB operations during that time. Third, using high frequency data, we are able to account for potential endogeneity between corporate credit spreads and monetary policy (Caldara and Herbst, 2019).

Our results are twofold. First, we provide evidence for a short-run adverse effect of ECB monetary policy on credit risk during the low interest rate period. This indicates a short-run dysfunctional credit channel, which is a relevant issue for the ECB since the Global Financial Crisis. Apparently, monetary policy actions are differently evaluated by market participants, depending on which signal they process from the decision of the central bank.[4] In the post-2009 low interest rate environment, which is also characterized by high financial stress in the euro area, investors may interpret an expansionary shock as a signal for worsening economic prospects. For example, recent studies find that conventional monetary policy actions in the aftermath of the Global Financial Crisis had relatively small effects on real activity but considerably in-

[3] We follow the standard event study approach to identify the surprise component of ECB monetary policy announcements (Bredin et al., 2009; León and Sebestyén, 2012; Haitsma et al., 2016). Different from the US, for which federal fund futures rates are available, there are no futures market instruments that track the euro area policy rate. Nevertheless, Bernoth and Von Hagen (2004) show that 3-month Euribor futures rates are a reliable predictor for the policy rates of the ECB.

[4] While the literature for the euro area is quite thin, Wright (2012), Beckworth et al. (2010), and Cenesizoglu and Essid (2012) show for the US that corporate bond yield spreads react significantly to monetary policy shocks. Javadi et al. (2018) provide evidence that the effect of the systematic component of monetary policy may lead to higher market uncertainty in crisis times and to an adverse response of corporate credit spreads.

creased market-based uncertainty measures (Hubrich and Tetlow, 2015; Jannsen et al., 2019).

Second, our results show that when ECB monetary policy solely influences the long-term expectations of market participants, the adverse effect on the credit spreads disappears for Italy and Spain. This finding sheds light on two issues. First, ECB monetary policy may have a different impact across the heterogeneous euro area countries. While investors in the Italian and Spanish corporate bond market seem to evaluate the lower interest rates in a positive way, investors in the French and German corporate bond market have a rather pessimistic view over the same event. This can show how the signal of worsening economic prospects can dominate the positive effect of low interest rates for these non-crisis countries because the benefit of lower interest rates for them is much weaker than for the crisis countries. Second, forward guidance appears to be an effective instrument for the ECB to conduct their monetary policy during times of high economic uncertainty.

The paper is organized in the following manner: in section two, we describe our data and the econometric framework. In the data section we provide a detailed explanation of our identification strategy for our monetary policy shocks and credit risk indicator measures. In section three, we present the empirical results. The last section draws conclusions based on our results.

1.2 Data and model

1.2.1 Data

We conduct our econometric exercise for four euro area countries: France, Germany, Italy, and Spain. Using daily data, our sample period is January 1, 2000, through November 23, 2015.

1.2.2 Credit spreads

We adapt the method of Gilchrist and Mojon (2018) and construct daily measures of country-specific credit spreads. In particular, we extract the corporate bond-specific credit risk component by eliminating the riskless component of micro-level corporate bond yields. This represents the risk premia an investor requires in addition to the riskless interest rate as compensation for holding higher risk.

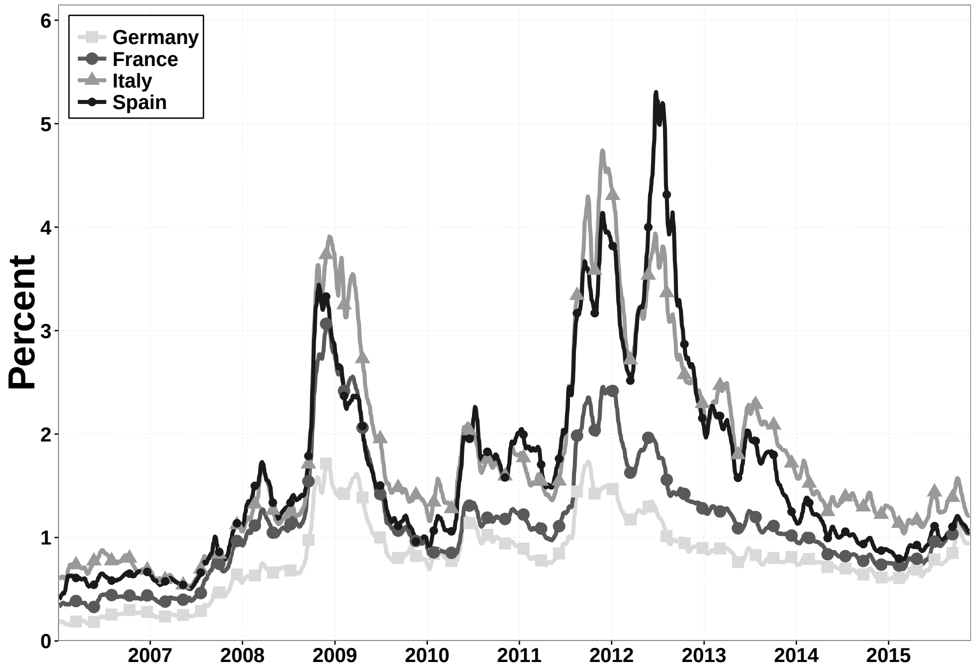
Specifically, the bond spread of a bond i of a country c at time t (cs_{it}^{c}) is defined as

$$cs_{it}^{c} = R_{ict} - ZCR_t^{DE}(Dur(i,c,t)), \tag{1.1}$$

where R_{ict} is the yield of bond i issued in country c at day t, and $ZCR_t^{DE}(Dur(I,c,t))$ is the corresponding risk-free yield with matched duration. We use interpolation methods whenever daily data of the same duration is not available. The country-level bond spread at time t is then calculated as the weighted average across all bond spreads in a given country: $cs_t^c = \sum_i \omega_{ict} cs_{it}^c$, where the weight ω_{ict} is the ratio of the market value of a bond i at issuance of the security relative to the total market value at date t.

We use effective yield data of fixed-coupon, euro-denominated, non-callable, and non-guaranteed securities of non-financial corporations. In total, we have micro-level information for 767 bonds from 122 non-financial corporations. We use the German bund zero-coupon bond yield rates as a proxy for riskless interest rates of the euro area. The credit spreads of the four euro area countries are illustrated in Figure 1.1.

Figure 1.1: Credit spreads, four euro area countries



Notes: The daily credit spreads of France, Germany, Italy, and Spain are constructed with the method of Gilchrist and Mojon (2018). For a better visualization of the data, we smooth the credit spread measures with the 10-day moving average. The unsmoothed series are presented in Figure 1.5 in the appendix. *Source*: Datastream.

1.2.3 Monetary policy shocks

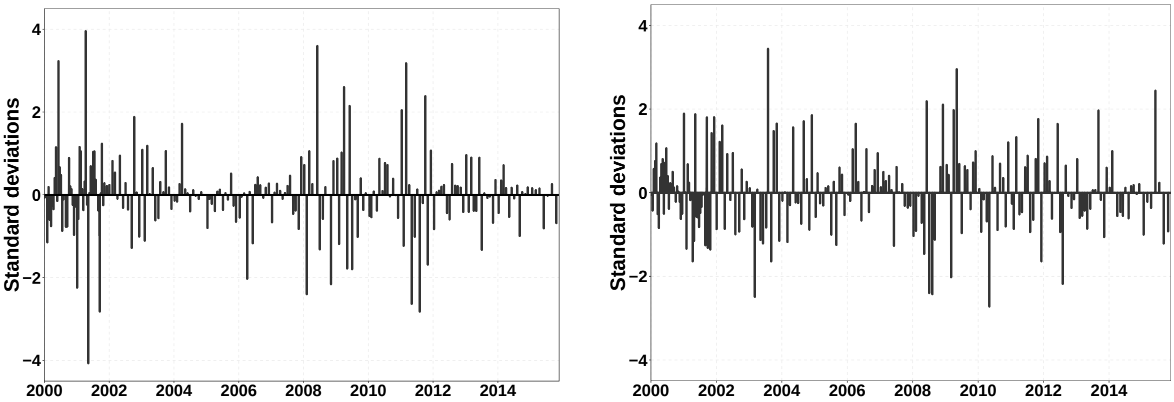
We apply the high frequency identification method to obtain measures of ECB monetary policy shocks. This method identifies monetary policy shocks as the surprise component of monetary policy actions, measured by the movements in asset prices on days of monetary policy announcements. Specifically, we apply the factor model of Gürkaynak et al. (2005) to extract two distinct dimensions of monetary policy by using the information of a broad range of money market futures.

Formally, the factor model representation of the $T \times N$ data matrix of the money market futures X can be expressed as

$$X = F\Lambda + \nu, \tag{1.2}$$

where F is the $T \times 2$ matrix of the two unobserved factors, Λ is a matrix of 2×2 factor loadings and ν is a $T \times N$ matrix of white noise errors. T represents the number

Figure 1.2: The two dimensions of monetary policy shocks



(a) Target shock

(b) Path shock

of ECB Board meetings in our sample and N is the number of money market futures rates included in the information set. After applying a principal component analysis, we use the rotation matrix of Gürkaynak et al. (2005) to obtain two factors $\tilde{f}_1$ and $\tilde{f}_2$.

The rotation builds on the assumption that the closest-to-deliver futures contract is not affected by changes in $\tilde{f}_2$. As a result, the obtained factors have a straightforward interpretation. The first factor, the target shock, can be interpreted as the surprise component of the current announcement. By construction, all the variation in the change of the Euribor futures rates with the shortest maturity is explained exclusively by this factor. As the two factors are orthogonal to each other, the second factor, the path shock, represents all other information released by the announcement above and beyond changes in the current short-term interest rate. Thus, the second factor is commonly interpreted as forward guidance.

Figure 1.2 visualizes the obtained two measures of monetary policy shocks. During our sample period, there were 212 meetings of the ECB Governing Council regarding monetary policy. The size of shocks and, in particular, the target shock is higher around the 2001, the 2008/2009, and the 2012/2013 recessions, than during the other periods. The identified shocks are, however, not systematically expansionary during recessions, but there are frequent positive as well as negative shocks.

1.2.4 Model

We use a daily flexible autoregressive distributed lag (ARDL) model since we are interested in the short-run effect of monetary policy shocks on the euro area credit conditions. Due to the forward-looking nature of the financial markets, monetary policy should have an immediate effect on credit spreads. For each country, we run a daily regression with the respective credit spread measures of the two dimensions of monetary policy shocks.

The Global Financial Crisis and the subsequent European sovereign debt crisis led to a structural change in the conduct of monetary policy and in the financing conditions of non-financial corporation. We take this into account in our analysis, distinguishing between two different regimes. The regimes are defined as:

1. regime A, where interest rates are *normal*, and
2. regime B, where interest rates are *low*.

We use the ECB Governing Council meeting on March 5, 2009, as the beginning of the low interest rate environment (Regime B). On this date, the ECB decided to cut the interest rate to a level below 2% for the first time in its history.[5]

Thus, the following two-regime ARDL model is:

$$\begin{aligned} cs_t^c = & I_t \left[\alpha_{A,0}^c + \sum_{i=1}^{p_A} \alpha_{A,i}^c cs_{t-i}^c + \sum_{j=0}^{q1_A} \beta_{A,1,j}^c shock_{1,t-j} + \sum_{k=0}^{q2_A} \beta_{A,2,k}^c shock_{2,t-k} \right] \\ & + (1 - I_t) \left[\alpha_{B,0}^c + \sum_{i=1}^{p_B} \alpha_{B,i}^c cs_{t-i}^c + \sum_{j=0}^{q1_B} \beta_{B,1,j}^c shock_{1,t-j} + \sum_{k=0}^{q2_B} \beta_{B,2,k}^c shock_{2,t-k} \right] + \varepsilon_t^c, \end{aligned} \tag{1.3}$$

where t denotes all working days in our sample period. cs_t^c is the credit spread of country c, $shock_{1t}$ and $shock_{2t}$ are the target- and path shocks, respectively. ε_c^t is the error term. The high frequency identification approach enables us to cleanly identify the impact of exogenous monetary policy shocks, so we do not need to include additional control variables for our model.

The dummy variable I_t takes the value 1 after March 5, 2009, and zero otherwise. To fully assess the response of the dependent variable over time, we model all three variables as dynamic. The maximum lag length p, $q1$, and $q2$ are determined by the Akaike information criteria for every country separately and we allow the lag length to differ between the two regimes. However, our results are robust to a fixed lag structure.[6] The lag orders are reported in Table 1.1.

1.3 Results

We separately examine the period of normal interest rates (regime A) and the period of low interest rates (regime B).[7] First, let us concentrate on the results of regime A. Figure 1.3 shows the impulse response functions of the credit spreads.[8] For all four countries, we observe a decrease in the credit spreads on impact. However, the response

[5]The shadow rate of the ECB, developed by Wu and Xia (2017), is below 0.5% since February 2009, reaching negative values directly after the MRO interest rate decreased to 1.5%. The shadow rate of the ECB's benchmark rate anticipates the effects of quantitative easing (QE) and central bank forward guidance and, thus, is not bounded below by 0%.

[6]See Appendix 1.H for the impulse response functions.

[7]To some degree, our results are sensitive to the inclusion of the ECB Governing Council meetings on October 2nd and 8th, 2008, which were directly after the collapse of Lehman Brothers. Therefore, we exclude these two meetings from our sample.

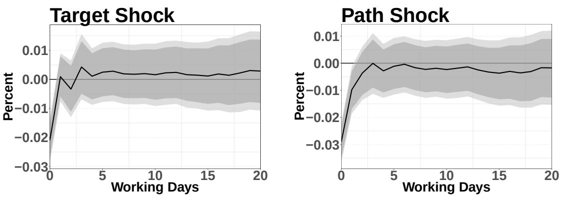
[8]As most of the dynamics happen in the first few response days, we show the impulse response functions for the first 20 business days. Longer IRFs are available upon request.

Table 1.1: Lag order of the two-regime ARDL model

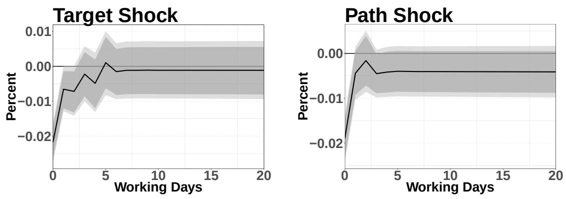
	Germany	France	Italy	Spain
p_A	4	19	10	16
$q1_A$	5	3	5	3
$q2_A$	2	3	2	3
p_B	13	12	12	7
$q1_B$	5	18	5	5
$q2_B$	5	7	4	6

immediately becomes insignificant. This holds for both target and path shocks, except for the positive response of the Spanish credit spread following an expansionary path shock. The mostly insignificant result is in line with the economic situation in regime A. At that time, bank loans were the primary financing instruments of European non-financial corporations (Ehrmann et al., 2003; Von Beschwitz and Howells, 2016; Deutsche Bundesbank, 2017). Since the supply of bank loans adjusts only very slowly to a change in the interest rates and the demand for bank loans of non-financial firms is rather fixed, an expansionary monetary policy shock in regime A should not affect the short-term credit conditions as measured by the corporate bond market of the euro area countries.

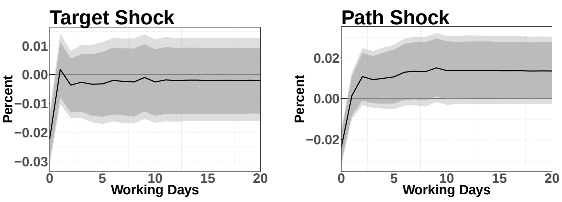
Figure 1.3: Estimated impulse responses in the two-regime ARDL model, regime A



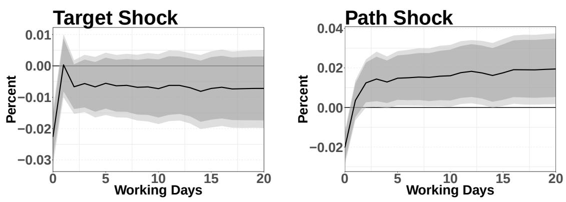
(a) France



(b) Germany



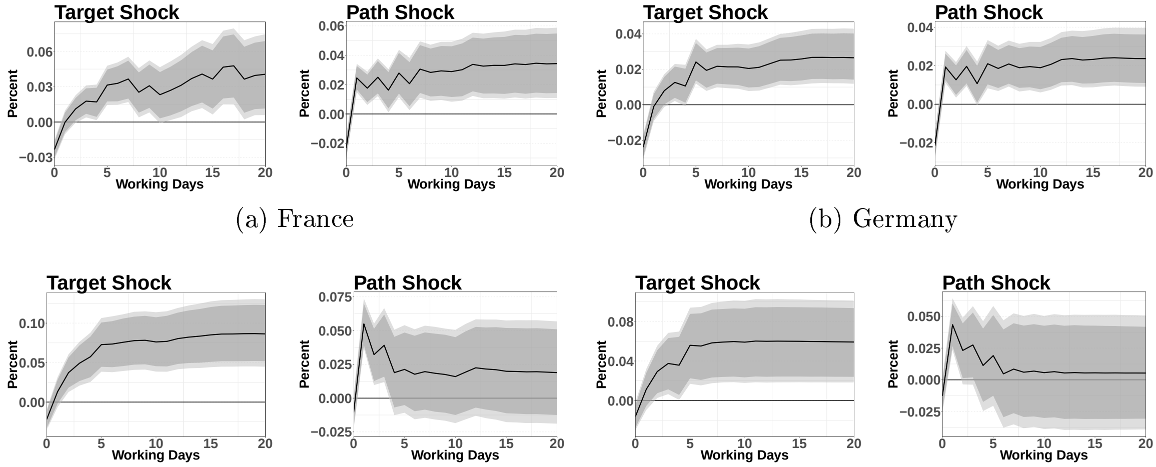
(c) Italy



(d) Spain

Notes: Impulse responses of the credit risk indicators to a one standard deviation expansionary monetary policy shock to the target factor and path factor, respectively. 90% (dark grey) and 95% (light grey) confidence intervals are produced by wild bootstrapping using the fixed design methodology (5000 replications). Sample period: March 6, 2009 - November 23, 2015.

Figure 1.4: Estimated impulse responses in the two-regime ARDL model, regime B



(a) France (b) Germany

(c) Italy (d) Spain

Notes: Impulse responses of the credit risk indicators to a one standard deviation expansionary monetary policy shock to the target factor and path factor, respectively. 90% (dark grey) and 95% (light grey) confidence intervals are produced by wild bootstrapping using the fixed design methodology (5000 replications). Sample period: March 6, 2009 - November 23, 2015.

In contrast, bank loans became limited and the markets for bonds of non-financial corporations in the euro area experienced strong growth in the aftermath of the Global Financial Crisis (De Fiore and Uhlig, 2015; Deutsche Bundesbank, 2017). Compared to bank loans, the cost of bond financing can vary on a daily basis because the demand for bonds in the financial markets is driven by the current expectations of the investors. This can explain our significant responses of the credit spreads in regime B, which are presented in Figure 1.4. While an expansionary target shock leads to an immediate decrease in the credit spreads of all four euro area countries, the responses become significantly positive for at least the next 20 trading days. For the path shock, we observe differences in the responses across countries. Following a path shock, we detect significantly positive responses for the French and German credit spreads, while there is no significant change for Italy and Spain.

According to the theory of the credit channel, an expansionary monetary policy shock should lower borrowing costs of non-financial corporations more than the fall in the risk-free rate. However, when the credit spread increases, we observe the opposite response. There is growing evidence in the literature on this adverse response of credit conditions of the private sector following a monetary policy shock during the crisis period. For instance, Bertsch et al. (2021) show that the liftoff of the Fed on December 16, 2015, led to an increase in the credit supply for households. They explain this phenomenon by the fact that an increase in the federal funds rates following a long lasting low interest rate environment may have provided a positive signal regarding the future solvency of the borrowers. Javadi et al. (2018) find that not only is it the actual policy rate decision of central banks that is important for the corporate bond market, but also the nature of the policy action. They analyze the systematic component of monetary policy and show, for example, how no-action by the FED (in terms of not changing the policy rate) during the Global Financial Crisis can lead to an increase in market uncertainty and widen corporate credit spreads.

One could think that the reassessment of investors' lending decision takes longer than the immediate change in the short-term interest rate following a monetary policy action. However, while the supply of bank loans adjust slowly to a change in the policy rate, the market price of corporate bonds should respond immediately to new information. Consequently, it seems unlikely that the adverse reaction of credit spreads is driven by a delayed response of investors.

Therefore, we consider how market participants evaluate the unexpected monetary policy action to explain this phenomenon. If interest rates are low due to weak economic conditions, how market participants evaluate a further surprising interest rate cut may instead be based on worsening economic prospects rather than on the ECB's intent of boosting economic activity. This can have a negative influence on the expectations regarding the creditworthiness of the non-financial corporations in the bond market and, in turn, affect the corporate borrowing rates. This interpretation matches with the growing literature that emphasizes the information effect of empirically identified monetary policy shocks (see i.e. Romer and Romer, 2000; Nakamura and Steinsson, 2018; Miranda-Agrippino and Ricco, 2018; Jarociński and Karadi, 2020).

In a robustness exercise reported in Appendix 1.C, we examine the macroeconomic consequences of a monetary policy shock identified with the high frequency identification approach to our credit spread indicator. Using a monthly proxy SVAR similar to Gertler and Karadi (2015a), the results show a significant increase in the credit spread that lasts about five months following a monetary policy easing shock, which is in line with our main analysis.[9] Moreover, the contractionary responses of industrial production on a similar horizon support the interpretation that monetary policy easing may have affected market participants' economic outlook.[10] After five months, however, the effect reverses: credit spreads decrease and macroeconomic variables are affected positively. Hence, while in the short run the credit channel is dysfunctional, in the medium run it works as the theory of monetary policy transmission predicts. Specifically, the pass-through of the decrease in the policy rate is amplified by the (i) improvement of the net worth of the borrower through the balance sheet channel and (ii) the increase in the liability of banks through the bank lending channel. These enhanced credit conditions have positive effects on economic activity, which is reflected in the increase of industrial production. Nevertheless, our results show that the transmission of ECB's

[9]The country-specific monthly VARs include the following endogenous variables: 1- or 2-year rates on German government bonds, industrial production (IP), the harmonized index of consumer prices (HICP), and credit spreads of non-financial corporations. The lag length is 12. As the post Global Financial Crisis sample is too short for meaningful inference, we estimate the VAR for the period January 2000 - November 2015. For details on the proxy SVAR see Appendix 1.C and 1.D.

[10]In another robustness exercise, we replaced all variables except the policy indicator and the credit spread by five factors obtained from a large panel of macroeconomic variables. This FAVAR specification also shows the adverse reaction of the credit spread indicator. Consequently, the response of the credit spread indicator is not driven by information insufficiencies in the VAR. Results are available in Appendix 1.G.

monetary policy is hampered in the short run due to its effect on market participants' economic outlook and risk assessment.

Our heterogeneous effect of monetary policy across crisis and non-crisis countries during the low interest rate environment can provide further evidence on how ECB monetary policy actions are evaluated by the European corporate bond market. Especially interesting is the fact that an expansionary path shock, which represents a flattening of the yield curve, negatively affects the credit conditions of non-crisis countries, while this is not the case for crisis countries. This result indicates that the investors in non-crisis countries evaluate the positive effect of lower interest rates in the future as less important than the negative signal of deteriorating economic prospects that may have led the ECB to the action. Different from this, an expansionary target shock leads to an increase in the credit spreads of all four countries.

1.4 Conclusion

Our results provide evidence that, in times of crises and low interest rates, expansionary ECB monetary policy interventions can have an adverse short-run effect on the credit conditions of euro area non-financial corporations. In addition, ECB monetary policy targeting long-run interest rate expectations appears to mitigate these adverse short-run effects for countries that were strongly affected by the crisis.

Our results suggest important policy implications for monetary policy in the euro area. First, we provide evidence for potential side effects of ECB monetary policy interventions on the bond market, which may dampen the originally intended effect of the interventions. Taking into account the increasing importance of market-based funding opportunities for firms in the euro area, the effect of monetary policy on this type of external funding must be taken more into consideration. Second, we show that the ECB is able to mitigate this distorting effect in the short-run, at least for the crisis countries, by relying on forward guidance and other measures that work primarily through the expectations channel of monetary policy. On March 10, 2016, the Governing Council of the ECB announced the Corporate Sector Purchase Programme. This operation aims to improve the financial conditions of corporations by buying their bonds on a large scale. In the light of our findings, this appears to be a promising venue to repair the monetary policy transmission mechanism of the euro area. In an early evaluation of the CSPP, De Santis et al. (2018) show that the introduction of this program improved the financing conditions of non-financial corporations by significantly reducing credit risk premia and, thus, corporate bond spreads.

1.A The construction of the ECB monetary policy shocks

We apply the factor model of Gürkaynak et al. (2005) to extract the surprise component of monetary policy announcements. Since we are analyzing the euro area, we use high frequency data of 3-month Euribor futures rates changes around an ECB monetary policy announcement date. Intra-daily data is unavailable to us, so we use the change in end-of-day closing prices surrounding ECB Governing Council decisions. In contrast to federal funds futures, we do not require a scale factor for the Euribor futures to account for the days remaining in the month after a policy action (Bredin et al., 2009; Brand et al., 2010). However, we account for illiquidity toward the maturity of the futures contracts and use the second closest-to-delivery contract instead of the current series whenever there are less than 5 days between the policy event and the next final settlement day. Moreover, we also include money market instruments with a longer time horizon. We consider German Treasury futures (Euro-Schatz, Euro-Bobl, and Euro-Bund futures as traded on the Eurex) in addition to the Euribor futures in the period after March 5, 2009, to account for a potential shift in the monetary policy regime since the ECB resorted to unconventional monetary policy measures. The yield changes of these Treasury futures are constructed as the daily return on the futures contract divided by the duration of the cheapest to deliver security in the deliverable basket. Table 1.2 reports the loadings of the two shocks. Both factors are normalized to have a unit standard deviation over the respective regime.

Table 1.2: Normalized loadings

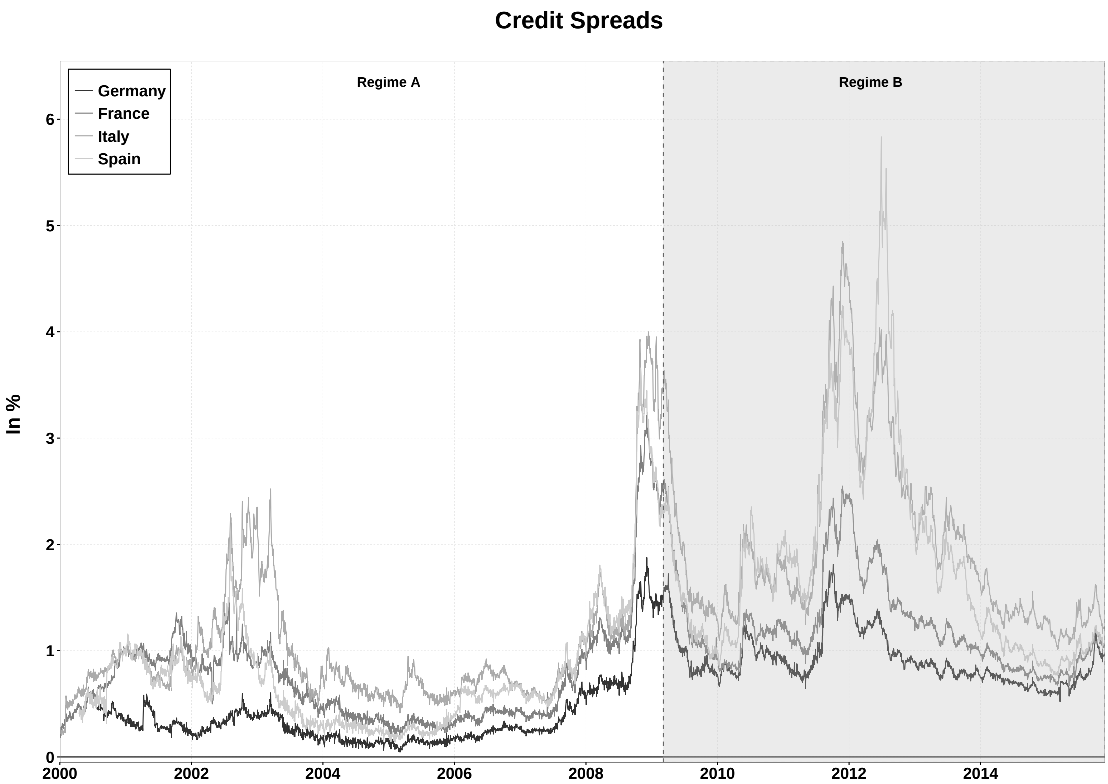
	Jan. 2000 - Mar. 2009		Mar. 2009 - Nov. 2015	
	Target Factor	Path Factor	Target Factor	Path Factor
1^{st} Futures	0.987	0	0.922	0
2^{nd} Futures	0.941	0.306	0.966	0.168
3^{rd} Futures	0.861	0.489	0.948	0.287
4^{th} Futures	0.779	0.622	0.901	0.401
5^{th} Futures	0.690	0.721	0.851	0.479
6^{th} Futures	0.619	0.774	0.802	0.539
Euro-Schatz			0.774	0.551
Euro-BOBL			0.522	0.838
Euro-Bund			0.280	0.913
Explained	89.8	8.6	82.2	11.6
Cum. Expl.	89.8	98.4	82.2	93.8

1.B Credit risk indicators

Table 1.3: Data description of the corporate bonds

	# Bonds	# Corp.	Observations	Avg. Issue Volume
France	295	41	500,000	EUR 600m
Germany	264	52	440,000	EUR 725m
Italy	138	21	245,000	EUR 750m
Spain	70	8	110,000	EUR 600m
Total	767	122	1,295,000	EUR 669m

Figure 1.5: Credit risk indicators of non-financial corporate bonds



1.C Macroeconomic implications

In this appendix, we investigate the macroeconomic implications of the effect of monetary policy on credit risk of non-financial corporations. We apply the SVAR approach proposed by Gertler and Karadi (2015a). They combine the traditional VAR analysis with high frequency identification using a monetary policy surprise measure as an external instrument to identify structural monetary policy shocks. We analyze France, Germany, Italy, and Spain separately using monthly data from January 2000 through November 2015.

The employed n-variate VAR model can be written as

$$y_t = a + B(L)y_{t-1} + u_t \quad u_t \sim iid\left(0, \Sigma_u\right), \tag{1.4}$$

where y_t is the n-variate vector, which contains economic and financial variables, a is a constant, $B(L)$ is a polynomial in the lag operator, and u_t is the reduced form error term with the variance-covariance matrix Σ_u. The VAR model of each country includes the following endogenous variables: Industrial production (IP), the Harmonized Index of Consumer Prices (HICP), the credit spread of non-financial corporations of Gilchrist and Mojon (2018), as well as the 1- or 2-year rates on German government bonds. Following Coibion (2012) and Gertler and Karadi (2015a), we set the lag order equal to 12 due to our monthly data.[11] The German government bond rate serves as the indicator for the stance of monetary policy, given that it is arguably a good proxy for the risk-free interest rate of the euro area. Furthermore, by using bonds with a maturity up to two years, we have a monetary policy indicator that also includes information regarding the change in expectations about the future path of monetary policy (Gertler and Karadi, 2015a).[12]

We are only interested in the effect of monetary policy shocks on economic activity, so that we partition the reduced form residuals in $u_t = \begin{bmatrix} u_t^p & u_t^{q'} \end{bmatrix}'$, where u_t^p is the reduced form residual of the policy indicator (in our case the German government bond rate) and u_t^q is the fraction of other reduced form residuals. Furthermore, we assume that $u_t = S\varepsilon_t$, where ε_t represents the structural shocks. The structural shock can be also decomposed into $\varepsilon_t = \begin{bmatrix} \varepsilon_t^p & \varepsilon_t^{q'} \end{bmatrix}'$, where ε_t^p represents the structural monetary policy shock and ε_t^q the other structural shocks. The relationship between u_t and ε_t can be explained by the unknown matrix S. The equation we estimate is then

$$y_t = c + B(L)y_{t-1} + s\varepsilon_t^p, \tag{1.5}$$

where s is an unknown vector that we need to identify from S.[13]

[11] For a textbook treatment of lag length selection in VARs, see Kilian and Lütkepohl (2017). In general, none of the results are sensitive to setting the lag length to a smaller value.

[12] The first-stage regression results show that the 1-year German government bond rate is a strong policy indicator for the analysis of Germany and France, while the 2-year German government bond rate is a better indicator for Italy and Spain. The first-stage results are available in Appendix 1.E.

[13] The whole model is $y_t = c + B(L)y_{t-1} + S\varepsilon_t,\ \varepsilon_t \sim \mathcal{N}\left(0, 1\right)$.

Following Gertler and Karadi (2015a), we apply the external instrument method to identify exogenous monetary policy shocks within the VAR model (Stock and Watson, 2012; Mertens and Ravn, 2013). The method requires an external instrument, Z_t, which fulfills the following assumptions:

$$E\left[Z_t \varepsilon_t^{p'}\right] = \Phi, \tag{1.6}$$

$$E\left[Z_t \varepsilon_t^{q'}\right] = 0, \tag{1.7}$$

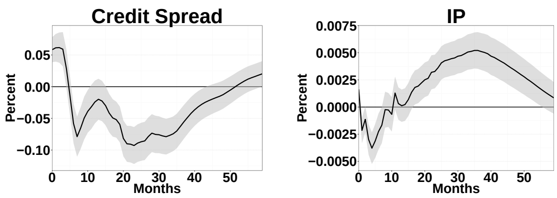
where $\Phi \neq 0$. These two assumptions show that a valid set of instruments must be correlated with the structural monetary policy shock, but not with other structural shocks.[14] We use daily changes in the 3-month Euribor futures around monetary policy announcement dates as an instrument. Given that we consider a very narrow time window around a monetary policy announcement, the change in the futures rates should exclusively represent the change in the expectations of financial market participants due to an unanticipated monetary policy action.

Figure 1.6 shows the impulse response functions of the credit risk indicators and industrial production from a one-unit expansionary monetary policy shock. In the mid- to long-run, the responses of the economic variables are consistent with the credit channel theory of monetary policy transmission and move in the expected direction. However, an expansionary monetary shock leads to an immediate increase in credit spreads for France, Germany, and Italy for up to 5 months.[15]

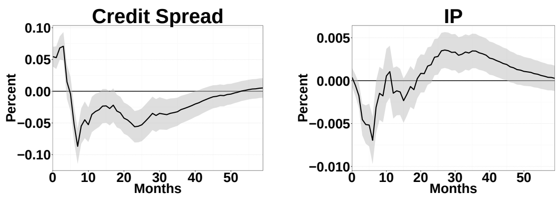
[14]Details on the derivation of the structural shocks and the impulse response functions are presented in Appendix 1.D.

[15]To exclude the possibility of misspecification of our econometric model, we conduct robustness exercises by (i) including a time trend in the model, (ii) changing the lag structure, and (iii) applying the Factor Augmented VAR framework to control for a potential information insufficiency. Our results are qualitatively robust with respect to these exercises. The results are available in Appendix 1.F and 1.G.

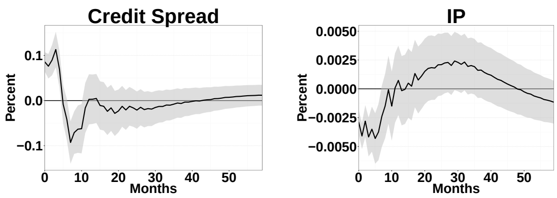
Figure 1.6: Estimated impulse responses in Proxy SVAR



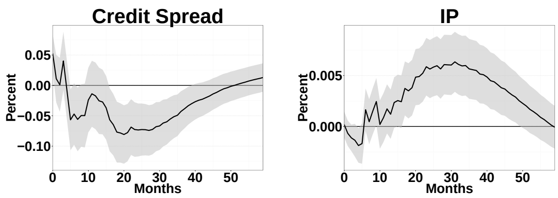
(a) France



(b) Germany



(c) Italy



(d) Spain

Notes: Impulse responses from an expansionary monetary policy shock. 95% bootstrap confidence intervals (5000 replications). Sample period: January 2000 to November 2015. Data source: Datastream.

1.D Details on the identification procedure: Proxy SVAR

There are n endogenous variables and k shocks of interest. We are only interested in the impact of monetary policy shocks, so we have $k = 1$. Given that the policy indicator variable is the first component of the vector y, we partition the first row s of the impulse matrix S, the reduced form and structural innovations in the following manner:

$$\underset{n\times 1}{s} = \begin{bmatrix} \underset{1\times 1}{s_{11}} & \underset{n-1\times 1}{s'_{21}} \end{bmatrix}', \quad \underset{n\times 1}{\varepsilon_t} = \begin{bmatrix} \underset{1\times 1}{\varepsilon_t^p} & \underset{n-1\times 1}{\varepsilon_t^{q\prime}} \end{bmatrix}', \quad \underset{n\times 1}{u_t} = \begin{bmatrix} \underset{1\times 1}{u_t^p} & \underset{n-1\times 1}{u_t^{q\prime}} \end{bmatrix}'.$$

The unknown matrix S and the variance-covariance matrix of the reduced form residuals are decomposed as follows:

$$S = \begin{bmatrix} \underset{1\times 1}{s_{11}} & \underset{1\times n-1}{s_{12}} \\ \underset{n-1\times 1}{s_{21}} & \underset{n-1\times n-1}{s_{22}} \end{bmatrix}, \quad \Sigma_u = \begin{bmatrix} \underset{1\times 1}{\Sigma_{u_p,1}} & \underset{1\times n-1}{\Sigma_{u_q,1}} \\ \underset{n-1\times 1}{\Sigma_{u_p,2}} & \underset{n-1\times n-1}{\Sigma_{u_q,2}} \end{bmatrix}.$$

The structural shock condition $u_t = S\varepsilon_t$ and the external instrument conditions (1.6) and (1.7) are necessary to identify the structural shocks of interest. Define s as the component of the $n\times n$ matrix S, which explains the impact of structural monetary policy shocks on the endogenous variables of the VAR model. Combining these conditions implies

$$\Phi s = \Sigma_{Zu'}, \tag{1.8}$$

where $\Sigma_{AB} \equiv E\left[A_t B_t\right]$ for any random vector or matrix A_t and B_t.

Partitioning $\Sigma_{Zu'} = [\Sigma_{Zu^{p\prime}} \quad \Sigma_{Zu^{q\prime}}]$, equation (1.8) can be expressed as

$$s_{21} = \left(\Sigma^{-1}_{Zu'_p}\Sigma_{Zu'_q}\right)' s_{11}, \tag{1.9}$$

where s_{11} and s_{21} are decomposed values of s. The dimension of s_{11} is 1×1 and s_{21} is $1 \times (n-1)$. The expression $\Sigma^{-1}_{Zu'_p}\Sigma_{Zu'_q}$ is estimable, so that we obtain additional restrictions for the identification of the structural shock ε^p.

Mertens and Ravn (2013) show that the estimation can proceed in the following procedure:

1. Estimate the reduced form VAR by least squares.
2. Estimate $\Sigma^{-1}_{Zu'_p}\Sigma_{Zu'_q}$ from the regressions of the VAR residuals u on Z_t.
3. Given $\Sigma_{Zu'}$ and Σ_u , calculate s using the equation (1.9) and the fact that $\Sigma_u = SS'$.

Further denote the standard deviation of ε^p by $\sigma_{\varepsilon,p}$.The vector s can be calculated as follows:

$$\begin{aligned}
s_{11}\sigma_{\varepsilon,p}^{-1} &= \left(I - s_{12}s_{22}^{-1}s_{21}s_{11}^{-1}\right)^{-1} \\
s_{21}\sigma_{\varepsilon,p}^{-1} &= s_{21}s_{11}^{-1}\left(I - s_{12}s_{22}^{-1}s_{21}s_{11}^{-1}\right)^{-1} \\
\sigma_{\varepsilon,p}^{2} &= \left(I - s_{12}s_{22}^{-1}s_{21}s_{11}^{-1}\right)s_{11}s_{11}'\left(I - s_{12}s_{22}^{-1}s_{21}s_{11}^{-1}\right)',
\end{aligned}$$

where

$$\begin{aligned}
s_{21}s_{11}^{-1} &= \left(\Sigma_{Zu_p'}^{-1}\Sigma_{Zu_q'}\right)' \\
s_{12}s_{22}^{-1} &= \left(s_{12}s_{12}'\left(s_{21}s_{11}^{-1}\right)' + \left(\Sigma_{u_p,2} - s_{21}s_{11}^{-1}\Sigma_{u_p,2}\right)'\right)\left(s_{22}s_{22}'^{-1}\right) \\
s_{12}s_{12}' &= \left(\Sigma_{u_p,2} - s_{21}s_{11}^{-1}\Sigma_{u_p,1}\right)' Q^{-1}\left(\Sigma_{u_p,2} - s_{21}s_{11}^{-1}\Sigma_{u_p,1}\right) \\
s_{22}s_{22}' &= \Sigma_{u_q,2} + s_{21}s_{11}^{-1}\left(s_{12}s_{12}' - \Sigma_{u_p,1}\right)\left(s_{21}s_{11}^{-1}\right)' \\
s_{11}s_{11}' &= \Sigma_{u_p,1} - s_{12}s_{12}' \\
Q &= s_{21}s_{11}^{-1}\Sigma_{u_p,1}\left(s_{21}s_{11}^{-1}\right)' - \left(\Sigma_{u_p,2}\left(s_{21}s_{11}^{-1}\right)' + s_{21}s_{11}^{-1}\Sigma_{u_p,2}'\right) + \Sigma_{u_q,2}.
\end{aligned}$$

1.E Policy indicator and instrument choice

We consider two policy indicators: the 1-year and 2-year German government bond rates. In addition, we also use the first twelve deliveries of 3-month Euribor futures rates. In order to obtain the best instrument, we apply the two-stage least squares: we first estimate the reduced-form VAR and then regress the reduced-form residuals of the policy indicator with the change of futures rates around a monetary policy announcement date. Stock et al. (2012) recommend a threshold value of ten for the F-statistic from the results of the first-stage regression. Results show that our optimal policy indicator and instrument are safely above this threshold.

Table 1.4: First-stage results

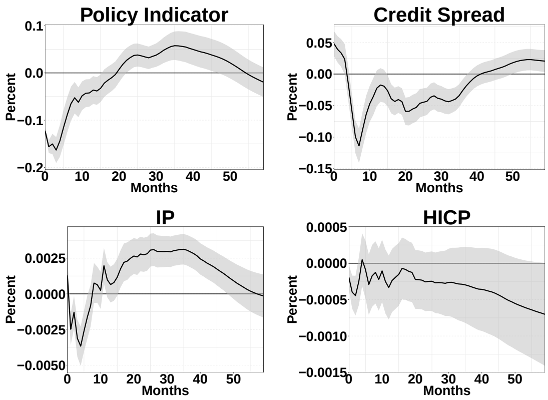
	Germany	France	Italy	Spain
F-statistic	31.8531	21.3714	36.4837	35.2774
PI	1-year	1-year	2-year	2-year
Futures	4	4	6	3

Notes: The first delivery of 3-month Euribor futures expires in the end of the current quarter, the second delivery expires in the end of the upcoming quarter, and so on.

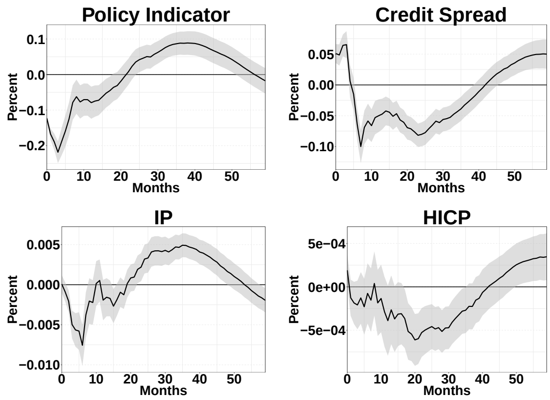
1.F Robustness check for proxy SVAR results

1.F.1 Proxy SVAR with time trend

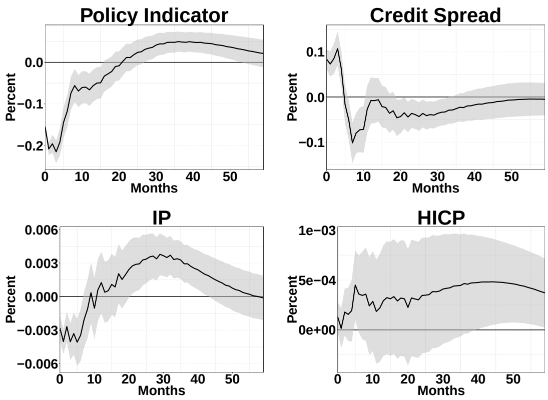
Figure 1.7: Estimated impulse responses in Proxy SVAR with time trend



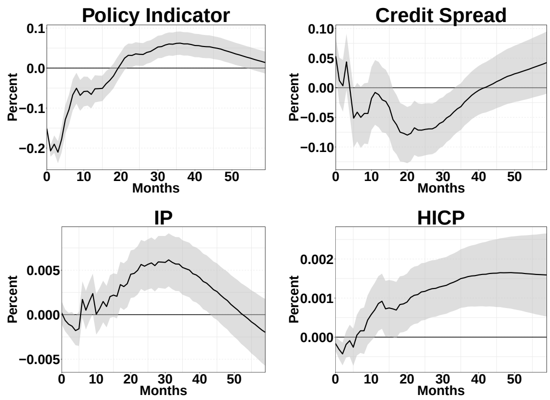
(a) France



(b) Germany



(c) Italy

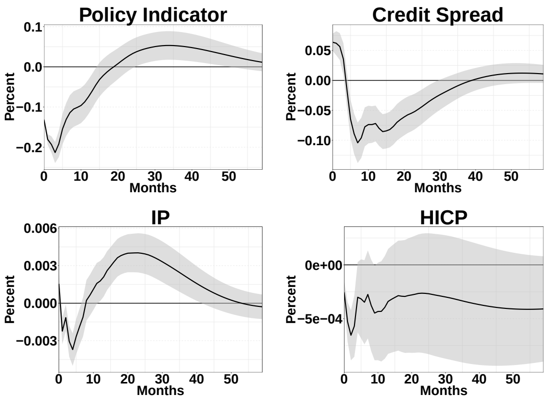


(d) Spain

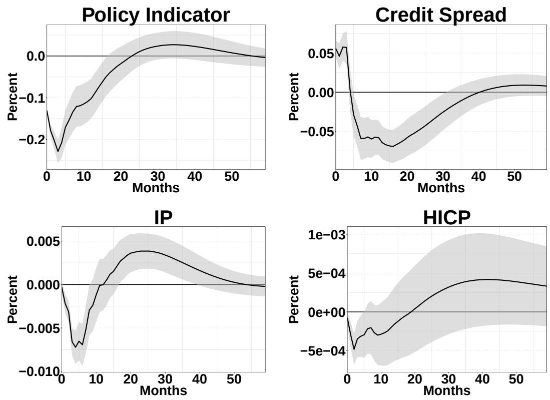
Notes: Impulse responses from an expansionary monetary policy shock. 95% bootstrap confidence intervals (5000 replications). Sample period: January 2000 to November 2015.

1.F.2 Proxy SVAR with 6 lags

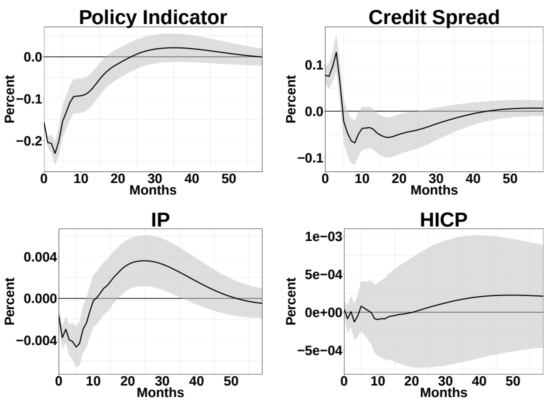
Figure 1.8: Estimated impulse responses in Proxy SVAR with 6 lags



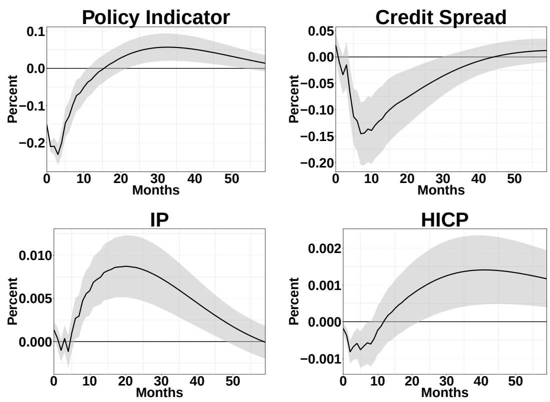
(a) France



(b) Germany



(c) Italy



(d) Spain

Notes: Impulse responses from an expansionary monetary policy shock. 95% bootstrap confidence intervals (5000 replications). Sample period: January 2000 to November 2015.

1.G Factor Augmented VAR with observable factors

We use the information set of a large panel of macroeconomic time series and apply a Factor Augmented VAR (FAVAR) model to detect the effect of monetary policy on the credit risk of France, Germany, Italy, and Spain (see Bernanke et al. (2005) and Yamamoto (2019)). By doing so, we address the potential information deficiency problem in small-scale structural VAR models and check the robustness of our results obtained with the proxy SVAR method with four variables. Within the FAVAR framework we extract common factors that explain most of the variation in the economy. In addition, we assume that the monetary policy indicator and the credit risk indicator are also factors that drive economic fluctuations. These are the *observable factors* of our model. We identify the exogenous monetary policy shocks by using again the external instrument approach.

We shortly explain how we extract the unknown common factors used in the FAVAR framework. Consider a $(n \times 1)$-matrix x_t that contains the n variables of the panel of country specific and euro area aggregated variables. We assume that x_t is driven by the unobservable factors f_t and the observable variables y_t (monetary policy indicator and credit risk indicator):

$$x_t = \Lambda^f f_t + \Lambda^y y_t + \varepsilon_t \tag{1.10}$$

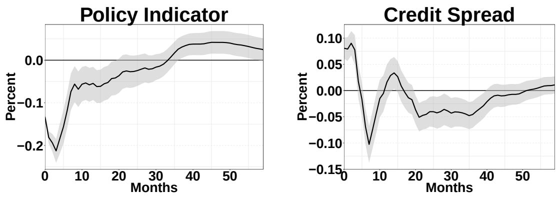
For the estimation of the unobservable factors f_t, we want to ensure the orthogonality of f_t and y_t by controlling for the variation of x_t driven by the policy variable and the credit spread indicator before extracting the unobservable factors. We do this by regressing x_t on y_t and apply the factor analysis on $z_t = x_t - \beta y_t$. The optimal number of latent factors is chosen with the method of Bai and Ng (2002) which yields 5 factors for Germany, Italy, and Spain and 3 factors for France.

Our panel consists of 84 monthly macroeconomic variables of the individual country as well as euro area aggregates. The balanced data set covers the time period January 2000 to November 2015. All time series are seasonally adjusted and appropriately transformed to ensure stationarity.

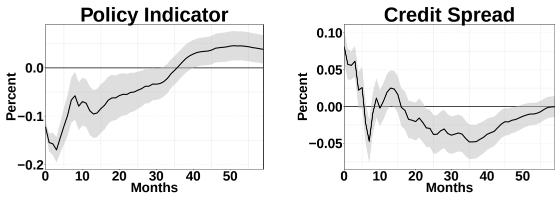
Since we are solely interested in the dynamic behavior of the observable factors, we do not focus on the structural identification of the whole model. Instead, we apply the partial identification method presented in section 2. The impulse response functions of the credit risk indicators from an expansionary monetary policy shock are shown in Figure 1.9.

Again, we observe an increase in the credit risk indicators of France, Germany, and Italy in the short run following an expansionary monetary policy shock. Additionally, we now observe a recovery of the Spanish credit conditions in the short run when there is an expansionary ECB monetary shock. In the standard proxy SVAR framework, we observe insignificant short-term responses of the Spanish credit risk indicator. Therefore, we are able to confirm our findings that an expansionary monetary policy shock leads to increasing credit risk indicators in the short run.

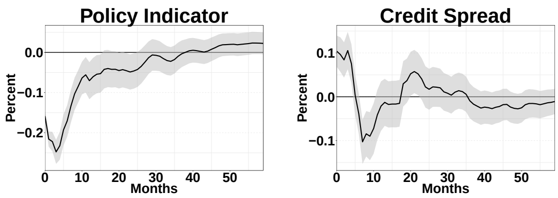
Figure 1.9: Impulse response functions in FAVAR



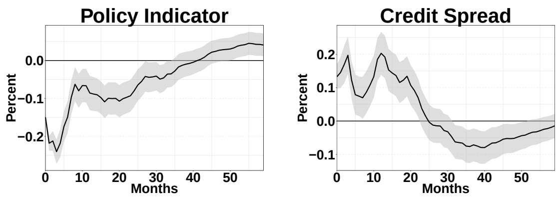
(a) France



(b) Germany



(c) Italy

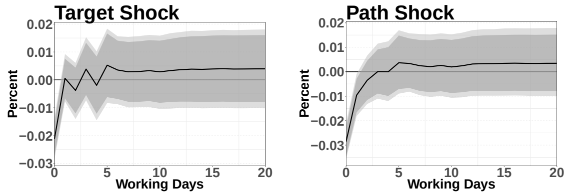


(d) Spain

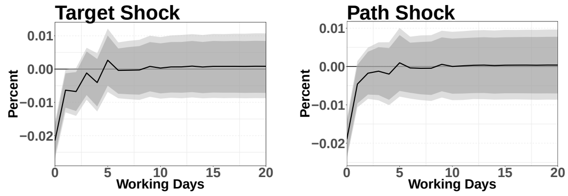
Notes: Impulse responses from an expansionary monetary policy shock. 95% bootstrap confidence intervals (5000 replications). Sample period: January 2000 to November 2015.

1.H ARDL model with a fixed lag structure (p=12, q1=q2=5)

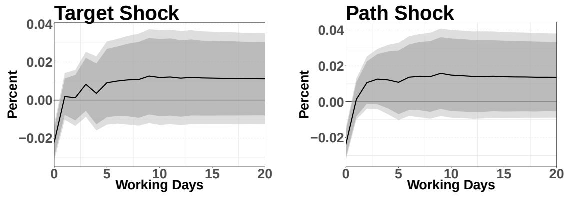
Figure 1.10: Impulse responses in the fixed lag two-regime ARDL model, regime A



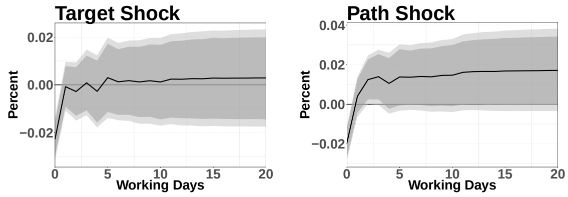
(a) France



(b) Germany



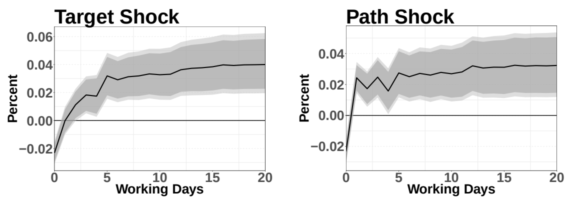
(c) Italy



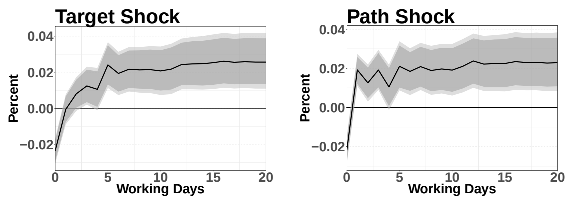
(d) Spain

Notes: Impulse responses of the credit risk indicators to a one standard deviation expansionary monetary policy shock to the target factor and path factor, respectively. 90% (dark grey) and 95% (light grey) confidence intervals are produced by wild bootstrapping using the fixed design methodology (5000 replications). Sample period: January 1, 2000 - March 5, 2009.

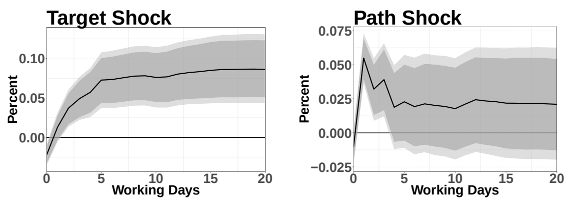
Figure 1.11: Impulse responses in the fixed lag two-regime ARDL model, regime B



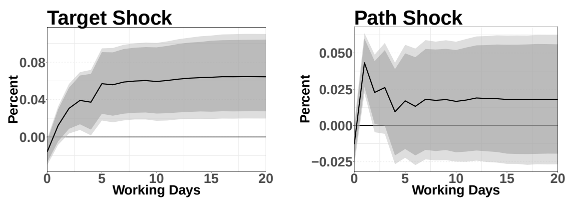
(a) France



(b) Germany



(c) Italy



(d) Spain

Notes: Impulse responses of the credit risk indicators to a one standard deviation expansionary monetary policy shock to the target factor and path factor, respectively. 90% (dark grey) and 95% (light grey) confidence intervals are produced by wild bootstrapping using the fixed design methodology (5000 replications). Sample period: March 6, 2009 - November 23, 2015.

CHAPTER 2

The Term Structure of Redenomination Risk[1]

Christian Bayer, Chi Hyun Kim, and Alexander Kriwoluzky

This chapter assesses redenomination risk in the euro area. We first estimate daily default-risk-free yield curves for French, German, and Italian bonds that can be redenominated in case of an exit and for bonds that cannot. Afterwards, we extract the compensation for redenomination risk from the yield spreads between these two types of bonds. Our results show that redenomination risk primarily shows up at the short end of the yield curve: at the height of the euro crisis, spreads between first-year yields were close to 7% for Italy and up to -1% for Germany. The ECB's interventions designed to reduce the risk of a breakup successfully did so for Italy, but increased it for France and Germany.

Keywords: euro crisis, redenomination risk, yield curve, ECB interventions
JEL classication: E44, F31, F45, G12, G14

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2.1 Introduction

Since the European sovereign debt crisis, a potential breakup of the euro area has been an ongoing concern, as we have again observed by the formation of the new Italian government in 2018. In particular, between 2010-2013, there were serious expectations that some countries might leave the monetary union with positive probability. Since such expectations can become self-fulfilling when they drive up sovereign yields (c.f. Obstfeld, 1986; Corsetti and Dedola, 2016; Coeuré, 2013; De Grauwe and Ji, 2013), these expectations of a breakup were one of the key motivations for the European Central Bank's (ECB's) interventions during the crisis.[2] Yet, even in hindsight, it is difficult to assess how likely the scenario was of a euro area breakup and how successful the ECB was in fighting expectations of a breakup. Our paper uses daily financial market data to estimate market expectations of a euro area breakup from differences in yield curves of securities, which are differentially affected by a country leaving the euro area.

A country (or a group of countries) breaking away from the euro area would introduce a new currency upon the event. This will be followed by a redenomination of contracts, likely including debt contracts, because the legal tender changes. However, the country is able to do so only for (debt) contracts that fall under that country's own domestic jurisdiction. Therefore, investors of assets governed by domestic jurisdiction will take this risk of a redenomination of their contracts into account. In particular, they will consider the possibility that a newly introduced currency might depreciate (or appreciate) vis-à-vis the euro. This introduces a spread between otherwise identical securities that differ only in jurisdiction. Choi et al. (2011), Clare and Schmidlin (2014), Corradin and Rodriguez-Moreno (2014), Trebesch and Zettelmeyer (2018), and Chamon et al. (2018) show that there are indeed systematic return differences in sovereign bonds issued under domestic and under foreign law. Importantly, this spread will have a term structure that reflects how the likelihood of a country's exit from the euro area changes over various horizons.

Therefore, we identify redenomination risk by comparing entire yield curves for default-risk-free returns on bonds issued under domestic and under foreign jurisdiction instead of bond yields for a given time to maturity.[3] The yield curves are estimated using a non-parametric approach following McCulloch (1971, 1975). The advantage of this approach is its flexibility and hence its potential to pick up redenomination risk at any time horizon. In order to apply the non-parametric approach, we have to collect an extensive data set of bond prices and coupon payments from Bloomberg and Datastream. We use *sovereign bonds* as bonds of domestic jurisdiction, i.e., as bonds that we expect to be redenominated in an exit event. As bonds that are always expected to be repaid in euro, we use *corporate bonds*, denominated in euro, emitted by an issuer in one country under another country's jurisdiction. To obtain riskless yield curves for these two groups of bonds, we control for default risk and subtract credit default swap (CDS) premia from

[2]See, e.g., the statement by the ECB's president, Mario Draghi, on July 26, 2012, in a speech at an investment bankers' conference

[3]Other authors, e.g. Krishnamurthy et al. (2018), have used a similar return-spread idea, however, using bonds with a fixed time-to-maturity (typically around five years).

all bond returns.[4] By doing so, on the one hand, we are able to obtain yield curves of default-risk-free domestic-law government bonds that only represent the riskless interest rate and redenomination risk. On the other hand, we obtain yield curves of default-risk-free foreign-law corporate bonds that only contain the riskless interest rate (without redenomination risk, from now on referred to as "safe international corporate €-bonds"). Comparing these two yield curves will give us our redenomination risk measure.

We have three main findings: First, for the period of the European sovereign debt crisis we find that default-risk-free Italian sovereign bonds show substantially higher yields compared to safe international corporate €-bonds. Interestingly, the opposite holds true for France and Germany, where the spread is negative. Second, these spreads show up mostly at the short end of the yield curve (mostly up to one year). Third, the spreads move systematically after the ECB's policy interventions. For Italy, default-risk-free sovereign short-term yields fell relative to the yield on safe international corporate €-bonds after the ECB's second Securities Market Program, the Longer-Term Refinancing Operations, and the announcement of the Outright Monetary Transactions. For France and Germany, the spreads decrease especially after the introduction of the first phase of the Securities Market Program.

Interpreting the spread along the lines of an uncovered interest rate parity as a compensation for expected changes in the exchange rate, financial markets were expecting Italy to potentially leave the euro area and depreciate its new currency. In particular, these expected changes in the exchange rate had substantial effects on the short end of the yield curve. They peak around the time of the collapse of the Berlusconi government at the end of 2011: the spread on one-year yields was roughly 7% in Italy. Even for France and Germany, the spreads are non-negligible with a *negative* one-year yield difference of, on average, around -1%. In other words, financial markets were considering the possibility that these countries might also leave the euro area, introducing a new currency that would then appreciate. Using this interpretation, the ECB's policy interventions have reduced redenomination risk in Italy, but increased it for France and Germany.

To demonstrate that our interpretation of the yield spread as redenomination risk is valid, we use an episode that we can expect to have no other impact on financial markets than through expectations of a breakup: the time when the German constitutional court examined the potential *ultra-vires* character of the ECB's Outright Monetary Transactions program. The court hearings took place between April 2013 and February 2014. At the start of the hearing, the court was surprisingly open to the complainants' case against the ECB's policy. At the same time, the one-year yield on German default-risk-free sovereign bonds fell substantially relative to safe international corporate €-bonds. Consequently, the spread quickly went to -1% and then slowly returned to zero when the constitutional court transferred the case to the European Court of Justice, which finally denied the *ultra-vires* character of the Outright Monetary Transactions

[4]Hull et al. (2004), Blanco et al. (2005), Ang and Longstaff (2013), Aizenman et al. (2013), and Arce et al. (2013) use credit default swaps as a direct measure of the price of default risk of debt issuers in the asset markets as well.

program. We view this episode as evidence that the spreads we construct measure redenomination risk indeed.

With these results, our paper contributes to a recent literature on the effect of ECB interventions on euro area financial markets. De Pooter et al. (2012), Eser and Schwaab (2013), Falagiarda and Reitz (2015), and Trebesch and Zettelmeyer (2018) show that ECB interventions are successful in decreasing the sovereign spreads of euro area crisis countries. Redenomination risk is a particularly important part of these spreads, because it necessarily affects all domestic interest rates in a country. In turn, this means that redenomination risk limits the ECB's capacity to fully affect the relevant short-term interest rates through its conventional monetary policy.

This is why a series of recent studies has focused on the prevailing breakup risk in the euro area. Some of this literature uses exclusively sovereign bonds and derivatives on them. Di Cesare et al. (2012) compare the sovereign yield spreads of euro area countries with their model-based values. Inter alia, they observe a strong divergence between these two measures during a time when the breakup of the euro area is frequently mentioned by market participants. In line with our results, they find evidence that market participants may have expected an appreciation of the new German currency and a depreciation of the currencies of "non-core" countries. De Santis (2019) constructs an empirical measure of redenomination risk for France, Italy, and Spain. Using a different approach from ours, which disentangles short-run from long-run redenomination risk, he examines redenomination risk at the five-year horizon. He uses five-year quanto government bond CDS of France, Italy, and Spain in relation to the quanto CDS of German government bonds as a benchmark, whereas we assume that German bonds under domestic law are also exposed to redenomination risk. Among other things, his analysis shows that redenomination risk has a significant impact on the sovereign yield spreads of the three countries. Kremens (2018) also utilizes CDS spreads of euro area securities to estimate redenomination risk and show how an exit scenario of France would involve a euro area break-up, while an Italian exit remains isolated. Krishnamurthy et al. (2018) assess the different channels of euro area sovereign bond yields. As in their paper, we use return differences between bonds traded under domestic and foreign law to estimate redenomination risk. They construct a rolling sample of bonds with three to five years to maturity and document the average return difference (after CDS premia) between the two types of bonds. We, in contrast, consider a wider range of bonds, using yield curve estimation to make the bond yields comparable. What is more, the yield curve estimates allow us to analyze the term structure of redenomination risk. Our findings suggest that it is important to look at short-term yields because redenomination risk is concentrated there. On a more theoretical level, our paper relates to Kriwoluzky et al. (2019). They set up a small open economy model in which a country is a member of a currency union at first, but where the possibility of an exit emerges and is reflected in return differences on sovereign bonds.

The remainder of the paper is organized as follows: Section 2.2 develops the empirical model we use to measure redenomination risk. Section 2.3 describes the data set we use. Section 2.4 presents the findings. In section 2.5 we provide robustness exercises to ensure the credibility of our identification strategy. Section 2.6 concludes. An appendix follows that describes the estimation method in detail and provides extensive robustness checks.

2.2 Identifying redenomination risk

Our measure of redenomination risk relies on estimates of yield curves for two sets of bonds: bonds issued under *domestic* and under *foreign* jurisdiction. For both types of bonds, we estimate default-risk-free yield curves out of bond prices, coupon payments, and credit default swap (CDS) premia. To illustrate our identification of redenomination risk, we start with the pricing of a risky bond by a risk-neutral investor.

2.2.1 Pricing a bond

A bond i is described by its promised coupon (and principal) payments $CF_i(\tau)$ at any payment date τ. We work in discrete time. The bond is subject to two fundamental risks: first, the issuer might default on the promised payments CF_i, or, second, the exchange rate $e(\tau)$ of the currency, in which the bond payments are denominated might change vis-à-vis the euro.

The price, $p_{i,t}$ (in €), which a risk-neutral investor is willing to pay for this bond at time t, is given by:

$$p_{i,t} = E_t \sum_{\tau > t}^{\infty} \frac{[1 - \pi_{i,t}(\tau)]\, e_t^{-1}(\tau)}{1 + r_t(\tau)} CF_i(\tau), \tag{2.1}$$

where $E_t\left[\pi_{i,t}(\tau)\right]$ is the expected probability the investor assigns at time t to the bond issuer's defaulting on $CF_i(\tau)$, $E_t\left[e_t(\tau)\right]$ is the exchange rate (in quantity quotation) the investor expects at trading-time t to hold at payment time τ to convert $CF_i(\tau)$ into €, and $E_t\left[r_t(\tau)\right]$ is the time value of money used to discount the future cash flows of the bond at payment date τ to their value at time t.

While expectations about the exchange rate and the time value of money should be the same across all bonds (of the same currency and under the same jurisdiction), expectations of default are bond specific. Therefore, we need to control for them to homogenize various bonds. In order to do this, we use CDS to directly identify the cost of the probability of default.[5]

A risk-neutral investor will be willing to buy a CDS if the premia to be paid on the swap, $CDS_{i,t}(\tau)$, equal the expected losses under default. Hence, the price of the bond with default risk should be equal to the price of a default-risk-free bond,

$$p_{i,t} = E_t \sum_{\tau > t}^{\infty} \frac{e_t^{-1}(\tau)}{1 + r_t(\tau)} cf_{i,t}(\tau), \tag{2.2}$$

whose cash flows $cf_{i,t}(\tau) \equiv CF_i(\tau) - CDS_{i,t}(\tau)$ are certain. Importantly for our measurement, the CDS premia we use are in terms of the principal's currency.

[5] Blanco et al. (2005) show that CDS prices have a valid relation to the theoretical default price of a bond and provide an upper bound of the price of credit risk.

2.2.2 The term structure of expected exchange rate changes

In this paper, we consider only bonds that promise payments in €. Therefore, as long as the country of the bond issuer remains in the euro area, the expected exchange rate $E_t[e(\tau)]$ is unity. Yet, a country leaving the euro area can redenominate contracts that are issued under domestic law, in particular its sovereign bonds, into the new currency it introduces. Therefore, when the investor assigns a positive probability that the country of the issuer will leave the euro area, exchange rate expectations can deviate from unity $E_t[e_t(\tau)] \neq 1$. Consequently, in the case when the probability of a country leaving the euro area is positive, $E_t[e_t(\tau)] < 1$ implies an expected appreciation and $E_t[e_t(\tau)] > 1$ an expected depreciation of its new currency vis-à-vis the euro. We group bonds by country of origin c issued under domestic law and estimate a discount rate for default-risk-free cash flows, $R^{dom}_{t,c}(\tau) = E_t\left[\frac{e^{-1}_{t,c}(\tau)}{1+r_t(\tau)}\right]$, from:

$$p_{i,t} = E_t \sum_{\tau>t}^{\infty} R^{dom}_{t,c}(\tau) cf_{i,t}(\tau). \tag{2.3}$$

Equation (2.3) implies that we can back out the exchange rate expectations by dividing the discount rate on domestic bonds $R^{dom}_{t,c}$ by a discount rate for bonds free of redenomination risk. To this end, we use bonds issued by a corporation in one euro area country under the jurisdiction of another country. Again, we control for default risk by using CDS premia and estimate the default-risk-free discount rate $R^{int}_t(\tau) = E_t\left[\frac{1}{1+r_t(\tau)}\right]$ as:

$$p_{i,t} = E_t \sum_{\tau>t}^{\infty} R^{int}_{t}(\tau) cf_{i,t}(\tau). \tag{2.4}$$

Thus, the exchange rate of country c expected at time t to be in place at a future time τ is

$$E_t[e_{t,c}(\tau)] = E_t\left[\frac{R^{int}_t(\tau)}{R^{dom}_{t,c}(\tau)}\right]. \tag{2.5}$$

Using these measures of expected exchange rates, we estimate the expected growth rate of the exchange rate between time τ_1 and τ_2 (expected at time t) as:

$$\Delta E_t[e_{t,c}(\tau_2,\tau_1)] \equiv E_t\left[\frac{e_{t,c}(\tau_2)}{e_{t,c}(\tau_1)} - 1\right] = E_t\left[\frac{R^{int}_t(\tau_2)}{R^{dom}_{t,c}(\tau_2)}\frac{R^{dom}_{t,c}(\tau_1)}{R^{int}_t(\tau_1)} - 1\right]. \tag{2.6}$$

One advantage of looking at these expected growth rate measures is that they correct for any fixed differences in domestic law and foreign law bonds that lead to proportionally higher discount factors for one type of bond or the other.

2.3 Data

We collect an extensive data set from Bloomberg and Datastream. The data contains prices for bonds, their coupon payments, and prices for CDS written on these bonds. Importantly, the CDS we use do not insure redenomination risk.[6] The data cover French, German, and Italian bonds. All bonds of our sample have a fixed-coupon and are euro-denominated, non-callable, and non-guaranteed. The sample runs from January 1, 2010 to September 21, 2014. The end-date of the sample is given by the introduction of new CDS that insure redenomination risk as well. Yet, also the old class is traded, but our data does not distinguish between both classes. A detailed report on how we collect the data is given in Appendix 2.B.

We consider sovereign bonds issued under domestic law as bonds that exhibit redenomination risk. We expect these bonds to be definitely redenominated into the new currency in the case of an exit from the monetary union because of their importance for the banking sector. Some of our bonds exhibit the Collective Action Clause (CAC), which allows for a supermajority of creditors to enforce the restructuring terms on minority holdout creditors. All euro area sovereign bonds issued after January 1, 2013 include these CACs. Nevertheless, this clause will not prevent the redenomination of sovereign debt under domestic law according to *lex monetae* (see Moore and Wigglesworth, 2017; Codogno and Galli, 2017).[7]

In Table 2.1 we present the bond data availability. For each country, we show the total number of bonds that we have in the sample and also the average fraction of bonds that will mature within three years. It is important that we have a sufficient amount of bonds during this three year horizon since we identify the risk of redenomination for the three year horizon.

Collecting data to estimate the daily yield curve for bonds that do not contain redenomination risk is challenging. We address this challenge in the following ways. First, we consider bonds issued by a domestic issuer that are subject to foreign law. Bonds falling under this category are bonds issued by large French and Italian corporations. Among these corporations are Carrefour, Thales, Enel, and Fiat. All foreign-law French corporate bonds are under English law, while the issuer is either from France or a subsidiary of the French parent company in the Netherlands. All Italian bonds are under English law and the issuer is either from Italy or a subsidiary of the Italian parent company in Luxembourg and Belgium.

We do not find German corporations that issue bonds under non-German law. However, many of the large German corporations issue their bonds through a subsidiary outside Germany (under German law). We include these bonds in our sample as well, however, show that our results are robust to their exclusion. Examples are bonds from Volkswagen International Finance BV with limited liability in the Netherlands emitted under German law. Further German corporations that we include, among others,

[6]The ISDA Master Agreement of 2002 explicitly states that for G7 countries such as France, Germany, and Italy, CDS contracts do not cover losses from redenomination risk.

[7]We also run the estimation without the CAC bonds, which does not change our results significantly. The results are available upon request.

Table 2.1: Bond data availability

Country	France	Germany	Italy
Domestic law sovereign bonds			
Number of bonds in total	53	79	106
Fraction (%) bonds below 3 year maturity	21	32	25
International law corporate bonds			
Number of bonds in total	113	135	105
Fraction (%) bonds below 3 year maturity	19	22	17

Sources: Bloomberg and Datastream.
Notes: The fraction of bonds below three year maturity is the average over all trading days of the number of bonds that mature within three years from a given trading day relative to all bonds at that trading day.

are Deutsche Telekom and Siemens, and they all issue bonds through similar vehicles. Their bonds do not contain redenomination risk for the following reason. First, they will not be redenominated in the case Germany exits the monetary union because they are issued by a foreign subsidiary of Volkswagen. Second, they will not be redenominated in the case of a Dutch exit from the monetary union, because they are issued under (from a Dutch point of view) foreign, namely German, law. Even if both countries exit, the issued bonds would still be bonds of a foreign issuer from a German point of view and therefore unlikely to be redenominated. Even if we only consider bonds issued under English law, none of our results change; see the appendix. For a complete overview of the corporations, see Table 2.2.

Since foreign-law bonds do not include country-specific redenomination risk, we are able to pool all bonds issued under foreign jurisdiction, irrelevant of the origin of the issuer. We then estimate a euro-area-wide default-risk-free corporate bond yield curve instead of a country-specific one. This improves the precision of our estimation due to more data points. In total, we have 353 international law corporate bonds in our sample.[8]

We control for default risk by subtracting CDS premia from the coupon payments and principal. All CDS premia we use are contracted as a fraction of the face value of the bond and therefore have to be paid in the same currency as the bond. Conse-

[8]As a robustness check, we also estimate country-specific default-risk-free corporate bond yields and show that our results do not change. Detailed information about the safe international corporate €-bond yields are provided in Appendix 2.C.2

Table 2.2: The issuer of the corporate bonds

Corporation	Rating	Headquarters	Issuer country	Jurisdiction	# Bonds
Airbus Group	A2	France	Netherlands	English law	7
Carrefour S.A.	Baa1	France	France	English law	18
Saint Gobain S.A.	Baa2	France	France	English law	23
Électricité de France	A1	France	France	English law	15
Lafarge S.A.	Baa2	France	France	English law	9
Thales S.A.	A2	France	France	English law	5
Total S.A.	Aa3	France	France	English law	19
Veolia S.A.	Baa1	France	France	English law	16
Wendel	Baa2	France	France	English law	1
BASF SE	A1	Germany	Netherlands	German law	4
BMW Group	A1	Germany	Netherlands	German law	17
Deutsche Telekom AG	Baa1	Germany	Netherlands	German law	36
EnBW AG	A3	Germany	Netherlands	German law	9
E.ON SE	Baa2	Germany	Netherlands	German law	20
Lanxess AG	Baa2	Germany	Netherlands	German law	2
Metro Group	Baa3	Germany	Netherlands	German law	5
Innogy SE	Baa2	Germany	Netherlands	German law	11
Siemens AG	A1	Germany	Netherlands	German law	10
Suedzucker AG	Baa2	Germany	Netherlands	German law	3
Volkswagen Group	A3	Germany	Netherlands	German law	18
Atlantia S.p.A	Baa3	Italy	Italy/Luxembourg	English law	9
Edison S.p.A	Baa3	Italy	Italy	English law	3
Enel	Baa2	Italy	Italy/Netherlands	English law	26
Eni S.p.A.	Baa2	Italy	Italy/Belgium	English law	20
Telecom Italia S.p.A.	Ba1	Italy	Italy/Luxembourg	English law	24
Leonardo S.p.A.	Ba1	Italy	Italy/Luxembourg	English law	6
Fiat S.p.A.	Ba2	Italy	Luxembourg	English law	17

Sources: Bloomberg, Datastream, and Base Prospectus of the issued corporate debt.

quently, we can only include bonds in our analysis for which we are able to obtain CDS prices. The CDS data set covers daily data on CDS prices of ten different maturities (six months, 1-5, 7, 10, 20, and 30 years). Since the bonds have different maturity dates, we construct a precise CDS price measure for different maturity dates by interpolating between the CDS prices. Almost all our CDS include the modified-modified restructuring (MM) clause, which is the standard convention for European corporate contracts.[9] The CDS prices we use take the form of a fraction of the insured principal and are thus denominated in € and in the newly introduced currency in case of redenomination.

[9]For some corporations, we were not able to obtain CDS with an MM clause. For these cases, we use CDS with the CR clause (CR=Full Restructuring). This is the case for one French and two Italian corporations. As a robustness check, we exclude the corporate bonds with CDS under the CR clause. The results do not change (available upon request).

2.4 Results

In this section, we present our estimates of the term structure of redenomination risk of France, Germany, and Italy during 2010 - 2014. After a detailed descriptive analysis, we examine the effect of ECB unconventional monetary policy actions on redenomination risk. We show that the ECB interventions were indeed able to decrease our redenomination risk measure of Italy, however, at the expense of slightly increasing redenomination risk for France and Germany. We use Google Search trends as an additional proxy that visualizes public concern regarding the euro crisis and confirm that exit expectations of Italy decreased around the policy announcement dates, while exit expectations of France and Germany increased. Also, as the Outright Monetary Transactions of the ECB was challenged in the German Federal Constitutional Court, we observe a significant shift in our German redenomination risk measure, implying that exit expectations increased for Germany at that time.

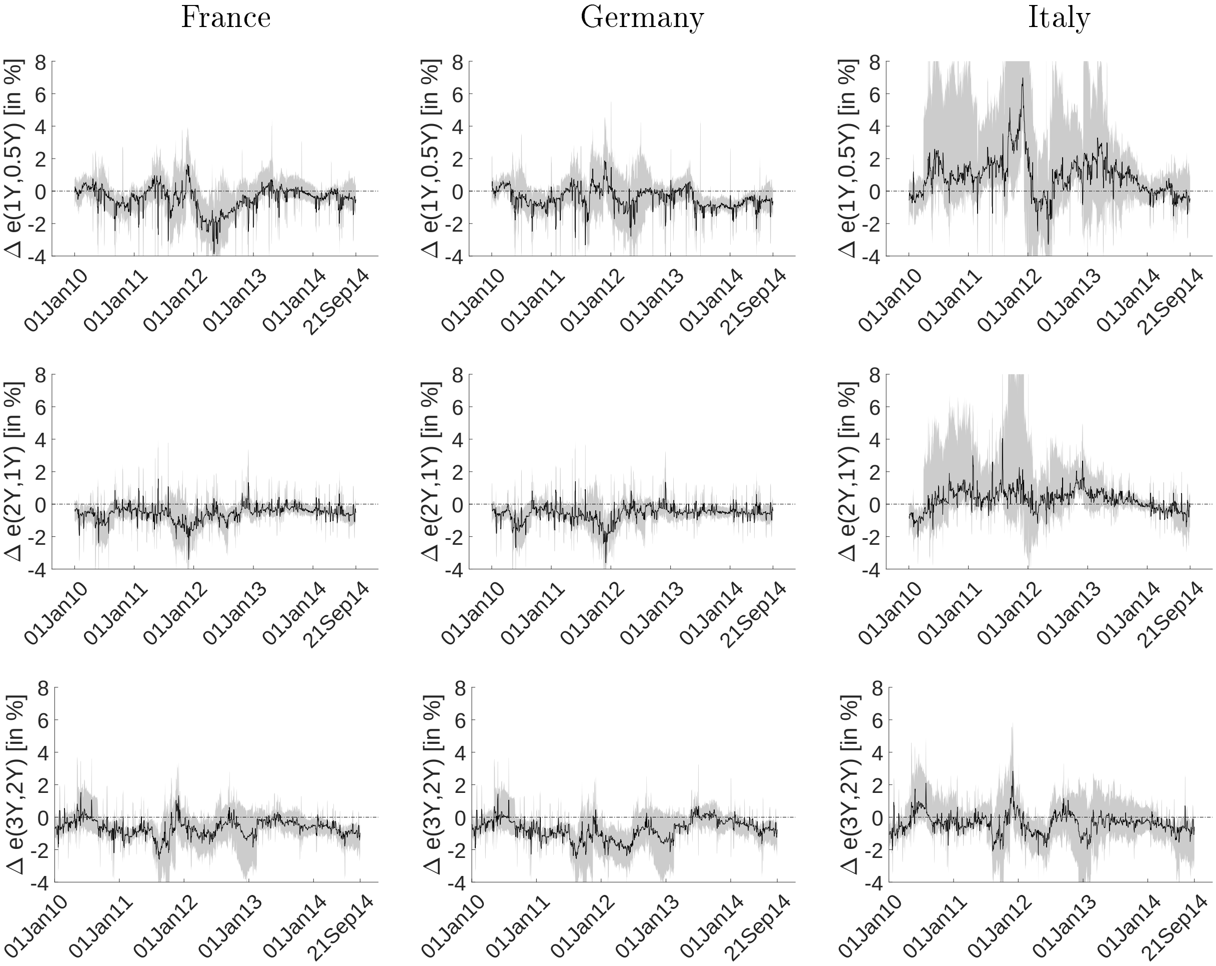
2.4.1 The term structure of redenomination risk

We identify the redenomination risk or, to be precise, the expected growth in the exchange rate, from the yield differences between default-risk-free sovereign bonds and the safe international corporate €-bond yield. First, we calculate the forward rates implied by our yield curve estimates for the first, second, and third year for each trading day. Then, using equation (2.6), we calculate redenomination risk from these spreads.[10] Figure 2.1 shows the result. The columns show the three countries, the rows the different horizons. For the interpretation, it is important to recall that the rows display spreads in the implied forward rates over the first, second, and third year from the trading day and not spreads of one-, two-, and three-year yields. Therefore, the first row reflects the expected exchange rate movement up until the end of the first year after the trading day, while the second row refers to the expected exchange rate movement between the end of the first and the end of the second year after the trading day. Analogously, the third row refers to expectations regarding the third year after the trading day. This means that one obtains, for example, the cumulated redenomination risk over the first three years from the trading day, i.e. the spread in the three-year yields, by summing over all three rows.

We find that redenomination risk influences the short end of the yield curve rather than the long end. For the first year and the second year after the trading date, the spread between the two yields is typically positive for Italy. This means that investors of Italian sovereign bonds had positive exit expectations for Italy and expected the new Italian currency to depreciate towards the euro. At the peak, which is around the collapse of the Berlusconi government at the end of 2011, the yield spread for Italy is around 7% for one-year yields and around 3% for second-year yields. For France and Germany, we observe negative spreads at least for the first year. The negative yield differences for France and Germany are also sizeable and around -0.75% on average,

[10]Note that our first year measure is the annualized expected growth in the exchange rate between the 183th day after the trading day and one year after. We do this to control for eventual illiquidity problems at the very short end of the yield curve. Nevertheless, if we calculate the spread between the first day after the trading day and one year after, we still obtain robust results.

Figure 2.1: Expected exchange rate movements



Notes: Expected changes in the exchange rate as implied by (2.6) for the estimated yield curves for CDS-insured sovereign bonds and CDS-insured international corporate €-bonds. The first row gives the annualized expected exchange rate movements between the 183th day after the trading day and the 365th day after. The second and third rows display the expected exchange rate movement between the 366th and the 730th day and between the 731st and the 1095th day after the trading day, respectively. In short, the rows display the expected exchange rate movement for the first, second and third year after the trading day. Shaded areas show bootstrapped 95-percent confidence bands.

with peaks close to -1.5% for one-year yields. The negative values show that exit expectations were also present for France and Germany, however, investors expected their new currencies to appreciate towards the euro.[11] Fluctuations and the differences from zero become smaller for horizons further away from the trading day. Within the third year from the trading day, the implied expected exchange rate movements for all three countries no longer differ systematically from zero. This result suggests that market participants' expectations of a breakup function as "expectation shocks" and have a significant influence on the sovereign bond yields of France, Germany, and Italy in the short run. This was exactly the concern of the ECB: the influence of such expectation shocks coming from expectations of a breakup, which can start a self-fulfilling spiral of redenomination risk and, thus, lead to an eventual real exit of the countries (European Central Bank, 2014). What is more, since redenomination risk affects mostly the short end of the yield curve, it affects exactly what conventional monetary policy uses as its instrument.

2.4.2 ECB policy interventions

In this section, we are interested in two issues. First, we examine whether the ECB's unconventional monetary policy interventions were able to reduce redenomination risk. Second, we also investigate the credibility of our measure by using Google Search trends that captures public concern regarding the euro crisis and redenomination risk.

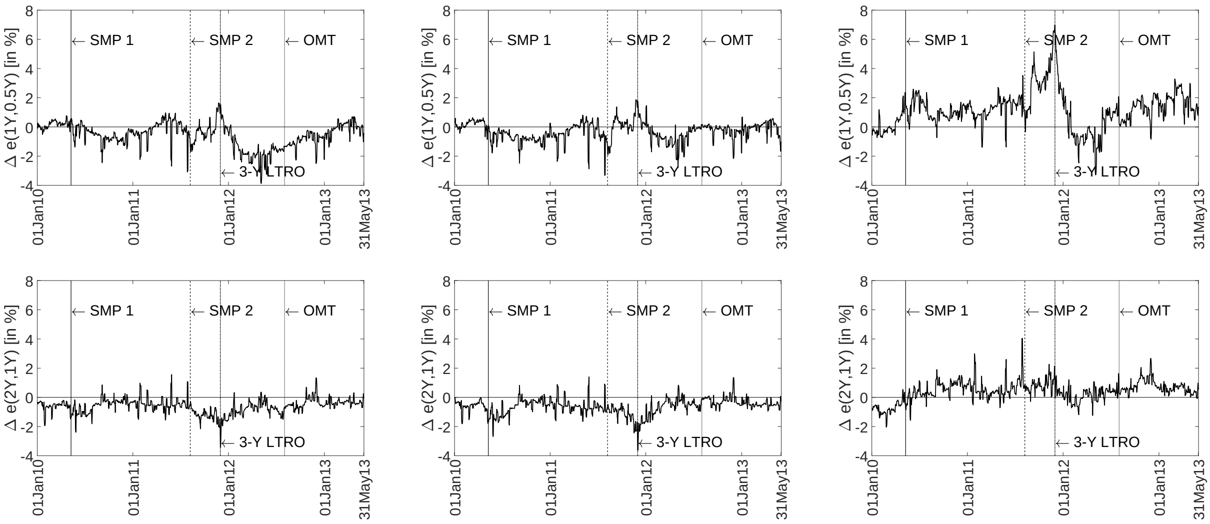
During the European sovereign debt crisis, the ECB intervened in the bond market by establishing four programs that aimed – among other things – to reduce expectations of a breakup as part of its unconventional monetary policy. On May 10, 2010 the ECB announced its first Securities Market Program (SMP-1), and the program was renewed (SMP-2) on August 8, 2011. On December 1, 2011 Mario Draghi, the ECB president, spoke to the European Parliament and signaled the introduction of a Longer-Term Refinancing Operations (LTRO) to stabilize the banking sector in the euro area, which was then officially introduced one weak later. Finally, on July 26, 2012 Mario Draghi gave his famous "whatever-it-takes" speech, which was followed by the ECB's official announcement of its Outright Monetary Transactions (OMT) program on August 2, 2012. The eventual modalities were made public on September 6, 2012.

Figure 2.2 displays the first- and second-year forward rate spreads from Figure 2.1, again interpreted as redenomination risk, together with the dates of the ECB's interventions. A first visual inspection suggests that all of the ECB's programs have brought down the spread between default-risk-free sovereign and safe international corporate €-bonds for Italy. Yet, they also seem to have decreased the yield of default-risk-free sovereign bonds relative to safe international corporate €-bonds in France and Germany. Given that the yield spread is typically negative for these countries before the ECB intervention, the spread increases in absolute value after the intervention.

[11] One might think that the Eurozone will break as a whole if France or Germany exits the monetary union. Nevertheless, it is reasonable to assume that the new French and German currency will appreciate towards the basket of the new European currencies.

Figure 2.2: ECB interventions and expected exchange rate movement.

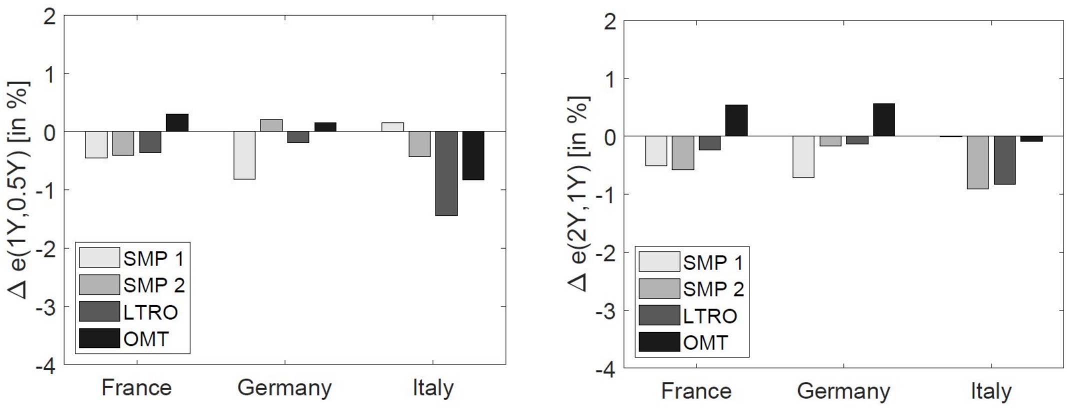
France Germany Italy



Notes: See Figure 2.1. The vertical dotted lines represent the days of ECB policy interventions. For the OMT program, we display both the date of the "whatever-it-takes" speech and the date when the details of the program are announced.

Figure 2.3: Expected changes in exchange rate around intervention dates

first year second year



Notes: The figure displays the average yield spreads between default-risk-free sovereign and safe international corporate €-bonds over the 20 days before and the 20 days after the ECB intervention. See Figure 2.1 for further information on the series of redenomination risk. The left panel refers to the yield spread over the first year, and the right panel refers to the spread in the second year yield. A negative number indicates a decline in the yield spread. For OMT we use August 2, 2012, the day of the official ECB announcement. For bootstrapped confidence bounds and significance, see Table 2.4 in Appendix 2.D.

Figure 2.3 shows how these spreads move around an ECB intervention. We calculate the average spread 20 trading days before and after an announcement/implementation of an ECB program. We find that the spread mostly declines after each program for all countries. For Italy, the effect of LTRO stands out by reducing redenomination risk of Italy by -1.5% at the first year and -0.8% at the second year. This can be explained by the fact that the LTRO not only alleviated financial constraints of Italy, but also may have had a soothing effect on the financial markets during the critial political situation in Italy, as the Prime Minister Silvio Berlusconi resigned on November 12, 2011.[12]

So far our results show that ECB interventions were able to reduce redenomination risk of Italy by decreasing the positive premia charged for the risk of Italy exiting the Eurozone and depreciating their new currency towards the euro. Also the spreads of France and Germany decrease after SMP. However, since investors expect an appreciation of the new French and German currencies in case of an exit, this indicates that the introduction of unconventional monetary policy interventions has rather increased the exit expectations of these two countries. OMT appear to have reduced redenomination risk in all three countries. In order to examine whether these scenarios are plausible, we use a measure that can proxy public concern regarding the euro crisis and thus exit expectations. For this, we use Google search trends.

We look at the search intensity for the term “euro crisis” in local languages (*crise euro, Eurokrise, crisi euro*) for France, Germany, and Italy and calculate how much the search intensity is affected by the ECB's intervention.[13] If the search intensity goes up, we view this as evidence that the public in the respective country is more concerned about the currency; if the search intensity goes down, public concern about the currency likely becomes less intense.[14] Since Google trends data come in the form of an index $s_t \in [0, 100]$, we calculate the rate of change as $\Delta s_t := 2\frac{s_t - s_{t-1}}{s_t + s_{t-1}}$.

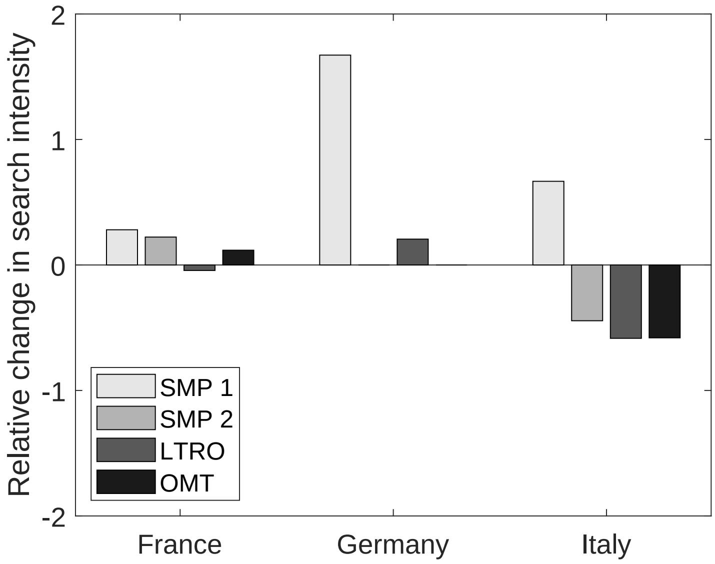
Figure 2.4 presents the results. In all three countries, search intensities go up after SMP-1, which is in line with our results for redenomination risk, where the absolute spread increases in all three countries, too. The effect of SMP-1 on German search intensities stands out, as did its impact on redenomination risk for Germany. Indeed, the introduction of the SMP-1 induced political turmoil in Germany: Axel Weber, the Bundesbank president back then, disagreed strongly with the ECB's decision to purchase sovereign bonds of crisis countries and resigned shortly after in February 2011. This disagreement in politics between the German Bundesbank and the ECB may have increased the probability of a German exit scenario. For SMP-2, LTRO, and OMT, we find again patterns that are in line with redenomination risk movements: search

[12] Results are significant as Table 2.4 in Appendix 2.D shows and robust to alternative specifications; see Appendix 2.C.

[13] If the intervention takes place before Wednesday, we compare the search intensity of the week before the intervention with the intervention week. If it takes place on Wednesday or later, we compare the week after the intervention with the week of the intervention.

[14] An issue here is, of course, that the language is not spoken just in the respective country. In addition, the term *crise euro* also means euro crisis in Portuguese. However, given the size of the countries relative to other European countries that use the same language or term, we can still attribute most of the change in search to the respective country.

Figure 2.4: Change in the search intensity for "euro crisis"



Notes: We display the relative change in the search intensity in Google for the term "euro crisis" in French, German and Italian. Data come from Google Trends for the period Jan 01, 2010 to Dec 31, 2012. The rate of change is calculated as $\Delta s_t = 2\frac{s_t - s_{t-1}}{s_t + s_{t-1}} \in [-2, 2]$ where s_t is the search intensity in the week of the intervention for interventions on Sunday to Wednesday and s_t is the search intensity in the week after the intervention for interventions between Thursday and Saturday.

intensity mildly goes up in France and Germany but it goes down in Italy. In other words, the ECB's policy interventions seem to have calmed the perception of the euro crisis in Italy as an interesting and urgent topic, but if anything, they have increased the perception of the euro being in crisis in the non-crisis countries.

2.4.3 The German constitutional court case regarding OMT

Finally, we conduct a case study on an event that should have exclusively influenced the market expectations about a country leaving the euro area. Examining the movements in our redenomination risk measure during this event will further support the credibility of our measures. The event we use is the German Federal Constitutional Court's (*Bundesverfassungsgericht,* BVerfG) decision regarding the legality of the Outright Monetary Transactions (OMT) program. Several individual members of the German parliament across all political parties had filed a complaint against the participation of any German government agency in both the European Stability Mechanism and the ECB's OMT program in 2012. The target of the complaint was the German federal government in general and the Bundesbank in particular. The BVerfG separated the two cases in 2012 and decided against the urgency of the complaint in the same year. In general, this was perceived as taking a pro-euro stance.

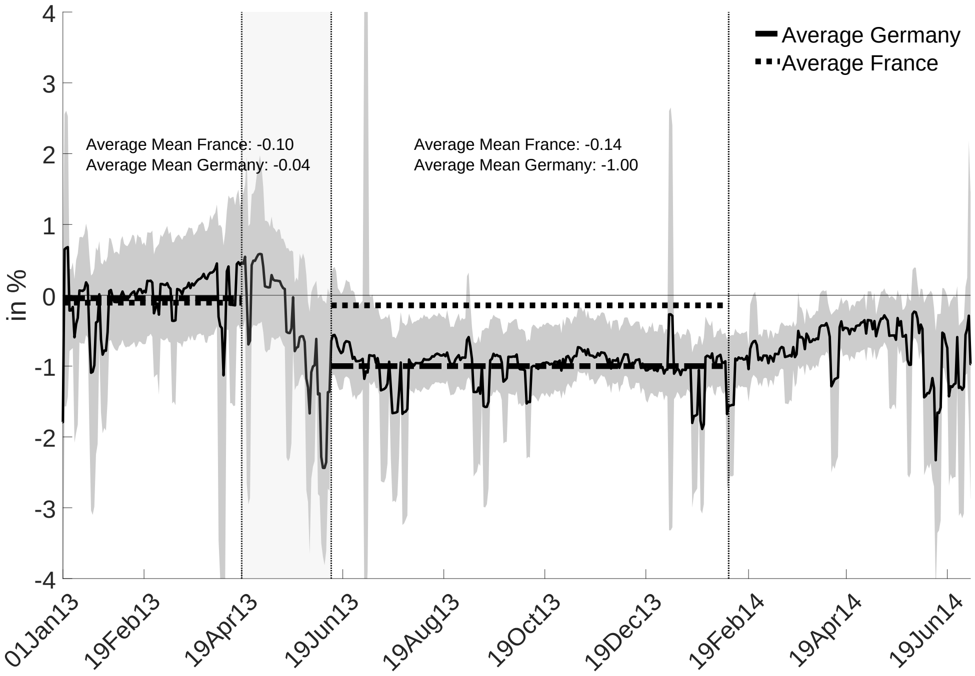
The court announced on April 19, 2013 that it would hold a two-day hearing on June 11/12, 2013, in order to prepare its final decision (press release No. 29/2013) on the OMT case. The hearing's agenda and the juridical topics to be discussed suggested that the court might now be leaning toward a critical view of the OMT. The court announced that it would discuss not only the potential *ultra-vires* character of the OMT, but also whether the program would touch the identity of the German constitution. Had the court ruled in favor of the former, the Bundesbank would have been obliged to stop any action that would directly or indirectly support the OMT program. Had the court ruled in favor of the latter, Germany would have needed to replace its existing constitution with a new one in order to remain in the euro area, which would have required a referendum. In both cases, a German exit from the EU would have been the likely inevitable consequence.

The court also decided to have economic experts testify during the hearing to assess the validity of the ECB's claim of an impediment to the monetary transmission channel (through the existence of redenomination risk). Importantly, the BVerfG's list of experts included some economists who had publicly announced their skepticism regarding the ECB's policy. Potentially the politically most important of those experts was the Bundesbank itself. On April 26, the bank's written statement to the court (from December 21, 2012) was leaked to the press. This statement was expected to be the basis of the Bundesbank president's upcoming testimony in court and it contained, among other things, a number of passages that were skeptical about sovereign spreads reflecting non-fundamental redenomination risk, the existence of which was a key argument for the ECB's program. Instead, the Bundesbank argued that redenomination risk was, if anything, an issue that reflected fundamental political uncertainty and as such is something that is and should remain outside the realm of monetary policy. Finally, in May 2013 a German think-tank produced a legal analysis by a former justice at the constitutional court (Di Fabio, 2013) that explicitly discussed a German exit from the EU as a potentially necessary consequence of the ECB's actions and the BVerfG's decision. The analysis was covered in the influential German quality newspaper, the Frankfurter Allgemeine Zeitung, on June 2 under the headline "Former constitutional court justice Di Fabio «In a pinch, Germany is obliged to leave the Euro»" ("Ehemaliger Bundesverfassungsrichter Di Fabio: «Notfalls ist Deutschland zum Euro-Austritt verpflichtet»").[15]

In sum, between the announcement of the hearing and the hearing itself, a number of pieces of information led to the conclusion that the court might come to a ruling that viewed the ECB program as being in conflict with the German constitution or the Treaty on the Functioning of the European Union. In fact, on February 7, 2014, the court ruled that the program, from the court's interpretation of the treaty, was indeed probably an *ultra-vires* act. Yet, the court did not reach a final conclusion, which would have forced Germany to exit the EU, but instead decided that it would ask the European Court of Justice (CJEU) for its judgment and interpretation of the treaty.

[15] The sharpest decline in the first-year yield spread happens on May 21, 2013. On that day the BVerfG officially announced that it would hear two additional experts (Harald Uhlig and Kai Konrad). Both these experts later happen to testify in a critical way regarding the ECB's OMT policy in the sense of it not being a monetary policy measure covered by the ECB's mandate (Konrad et al., 2013).

Figure 2.5: Expected exchange rate movements around the German Federal Constitutional Court hearings



Notes: See Figure 2.1. The shaded area covers the time period between the announcement of the hearing on April 19, 2013 and the BVerfG hearing on June 12, 2013. The case was widely debated during this period. One particular event is the publication of the Di Fabio (2013) paper at the end of May. The second vertical line is the day on which the BVerfG handed the case over to the CJEU. Shaded areas show bootstrapped 95-percent confidence intervals.

In his final statement on January 14, 2015, the advocate general of the CJEU suggested that the CJEU view the OMT program as being in line with the EU treaty, a line the CJEU followed in its decision on June 16, 2015. The BVerfG's final decision on the case was published on June 21, 2016. It follows the CJEU's ruling but qualifies it in setting limits on what the OMT program would need to look like in practice for German agencies to be allowed to participate.

In line with our interpretation of the estimated spreads as redenomination risk, the estimated spreads for Germany pick up this episode. Between the announcement of the hearing and the hearing itself, the spread between German default-risk-free sovereign bonds and safe international corporate €-bonds becomes more negative and declines to around -1%, where it stays until the BVerfG decides to have the case judged by the CJEU. From that point, the spread slowly disappears; see Figure 2.5. Remarkably, the persistent drop shows up only in the German spreads but not in the French spreads.[16] This supports our interpretation of the spread as indeed reflecting redenomination risk.

[16]Importantly, there is no change in the composition of the German sovereign bonds in our sample such that the strong and persistent decline in the one-year yield does not reflect unobserved heterogeneity in the bonds.

2.5 Robustness

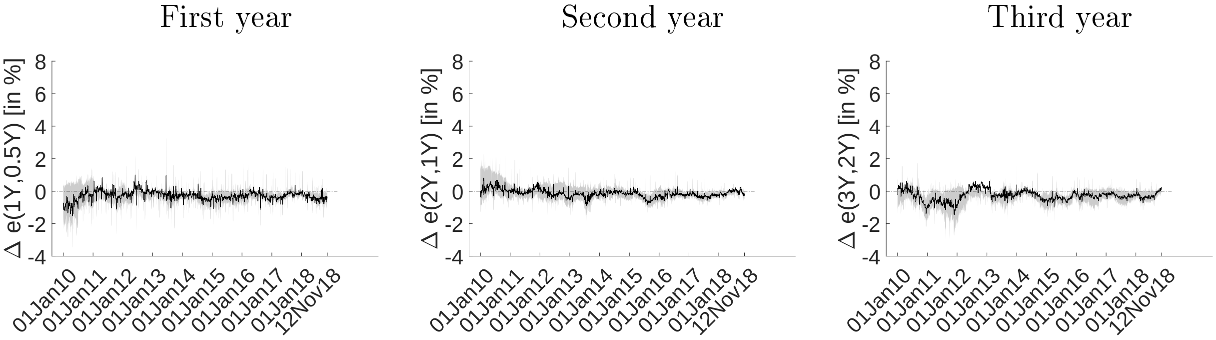
In this section, we examine the credibility of our identification strategy and redenomination risk measures. As a first step, we conduct a placebo test with US data, where we again compare sovereign yields to default-risk-free yields of large international corporate bonds. Since there is no redenomination risk in the US, we should find the spread to be close to zero. Indeed, this is the case. Second, we address the potential risk factors that may distort our measures of redenomination risk. Specifically, we consider the existence of CDS counterparty risk, liquidity risk, and legal risk of securities.

2.5.1 Placebo test with US data

We test whether our identification method picks up risk components other than redenomination risk. In order to do this, we run a placebo test with US corporate bonds and sovereign bonds and calculate a fictitious "redenomination risk" measure for the US as we did for the three Eurozone countries. Since investors of US securities are not exposed to redenomination risk, the estimated measure should fluctuate around zero and should not respond systematically to monetary policy actions of the Federal Reserve.

First, we estimate a default-risk-free yield curve of US corporate bonds. As for the euro area countries, we consider non-financial corporate bonds that are issued in the global market and for which we are able to find liquid CDS data. Table 2.3 lists the corporations that we have in our sample. In total, we have 663 bonds of 33 non-financial corporations. We use outright the yield curve measures of US government bonds that are available from the US Department of the Treasury since US government debt is free of default risk.[17]

Figure 2.6: Redenomination risk of the USA



Notes: Expected changes in the exchange rate as implied by (2.6) for US Treasury yield curves and CDS-insured global US corporate bonds. The first row gives the annualized expected exchange rate movements between the 182th day after the trading day and the 365th day after. The second and third rows display the expected exchange rate movement between the 366th and the 730th day and between the 731st and the 1095th day after the trading day, respectively. In short, the rows display the expected exchange rate movement for the first, second and third year after the trading day. Shaded areas show bootstrapped 95-percent confidence bands.

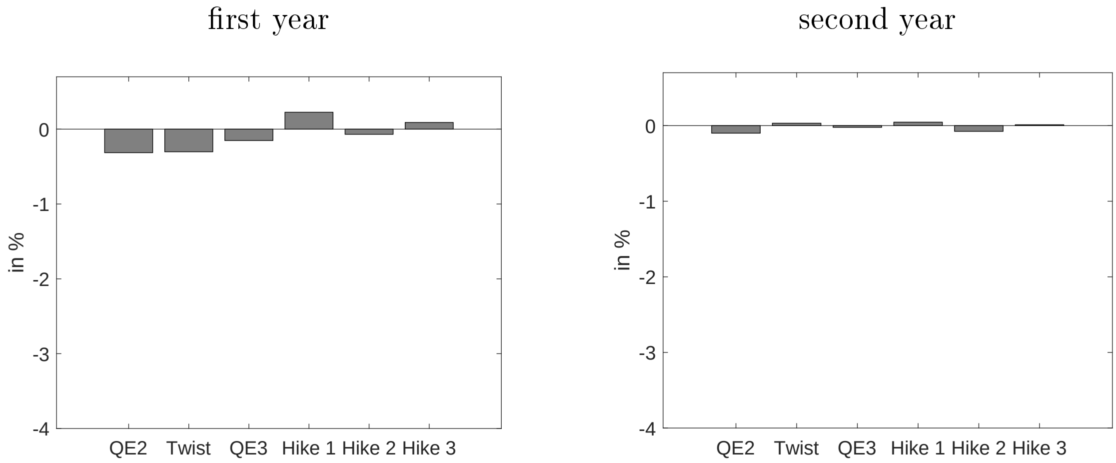
[17]Source: https://treasury.gov

Table 2.3: The issuer of the U.S. corporate bonds

Corporation	# Bonds	Rating
Amerisourcebergen	5	Baa2
AT&T Inc.	44	Baa2
Berkshire Hathaway Inc.	12	A3
Boeing Company	26	A2
Cardinal Health Inc.	13	Baa2
Caterpillar Financial Services Corporation	10	A3
Caterpillar Inc.	10	A3
Chevron Corporation	18	Aa2
Cisco Systems, Inc.	25	A1
Comcast Corporation	33	A3
Costco Wholesale Corporation	6	Aa3
Dell Inc.	10	Baa2
Dow Chemical Company	39	Baa2
Exelon Corporation	6	Baa2
Exxon Mobil Corporation	12	Aaa
Fedex Corporation	8	Baa2
Ford Motor Credit Co.	52	Baa3
Home Depot Inc.	18	A2
Honeywell International Inc.	13	A2
HP Inc.	8	Baa2
Johnson & Johnson	31	Aaa
Oracle Corporation	29	A1
PepsiCo	36	A1
Procter & Gamble Company	24	Aa3
Target Corporation	14	A2
The 3M Company	5	A1
The Coca-Cola Company	13	Aa3
The Kroger Company	26	Baa1
The Walt Disney Company	32	A2
TJX Companies Inc.	3	A2
Valero Energy Corporation	9	Baa2
Verizon Communications Inc.	32	Baa1
Walmart Inc.	35	Aa2
Total	663	

Sources: Datastream

Figure 2.7: Expected changes in exchange rate around FOMC meetings



Notes: The figure displays the average yield spreads between default-risk-free sovereign and default-risk-free corporate bonds over the 20 days before and the 20 days after the FOMC meetings. The left panel refers to the yield spread over the first year, and the right panel refers to the spread in the second year yield. A negative number indicates a decline in the yield spread. For bootstrapped confidence bounds and significance, see Table 2.7 in Appendix 2.D.

Figure 2.6 shows the ratio between the default-risk-free forward rates of U.S. corporate and government bonds at the first, second, and third year horizon. What strikes the eye is the magnitude of the measures. In the first year, the ratio seems to fluctuate around zero with no systematic movements. This is very different from the results that we obtain from all three euro area countries, for which we detect large redenomination risk at the first year horizon. As we move to the further time horizons, we still observe quite small values.[18]

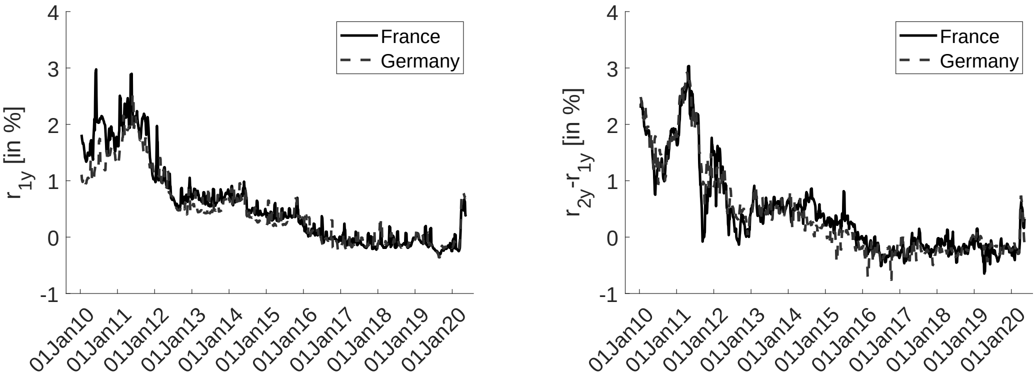
In addition, we are interested in whether our measures react to monetary policy announcements of the Federal Reserve (FED). We calculate the average spread 20 trading days before and after six FOMC meetings, when the Fed announced important insights regarding its monetary policy. The first three announcements are related to Quantitative Easing (QE). On September 13, 2012 the Fed announced the second phase of QE. On September 21, 2011 the Fed announced “Operation Twist”, which involved selling \$400 billion in short-term Treasuries in exchange for the same amount of longer-term bonds. The third phase of QE was announced a year after, which was on September 13, 2012. The latter three announcements reveal the Fed’s intention to keep the federal funds rate low, even though the markets were “ready” for a liftoff (December 17, 2014; March 18, 2015; September 17, 2015).

The results are plotted in Figure 2.7. Different from our euro area analysis, we do not detect systematic movements in the spreads after a FOMC meeting, neither for the first or the second year. In addition, the magnitude is very small, such that we can be confident of the fact that we are not measuring any systematic risk components that are influenced by monetary policy.

[18]For all time horizons, only approximately 20% of the measures is significant.

Figure 2.8: Difference between German and French international corporate bonds

first-year yield rate second-year forward rates



Notes: The first column displays our estimates of first-year yields (one-year yields minus one-day yields) of bonds issued by French corporations under English law and of Dutch subsidiaries of German corporations under German law (upper panel) and the difference thereof (lower panel). The second column displays the same for second-year forward rate (two-year yields minus one-year yields).

2.5.2 Counterparty, liquidity, and legal risk: Why the term structure of redenomination risk is important

Our measure of redenomination risk is a spread between default-risk-free sovereign bonds and safe international corporate €-bonds. Yet, this spread could potentially pick up additional factors other than redenomination risk. Liquidity risk, counterparty risk regarding the CDS, and legal risk are three potential culprits here.

Let us first address the counterparty risk of CDS. A CDS buyer is exposed to counterparty risk if the CDS does not fully insure against the risk of a default because the insurer itself might (partly) default. The basic principle of counterparty risk is very similar to the default risk of debt issuers: at each trading date t, the CDS buyer assigns a probability to the CDS seller's default on the payments and thus is willing to pay a lower CDS premium than without this risk.[19] This probability affects all CDS sold by the CDS seller similarly. More precisely, it affects even those with different maturities, because if the CDS seller is expected to default, then she should default on all payments, regardless of their maturity. Therefore, counterparty risk should be correlated across different time horizons, just as we observe this in annual CDS premia (see, e.g., De Santis, 2019). If our redenomination risk measure is exposed to this counterparty risk, we should observe this correlation across different maturities. However, this is not the case, as we do not observe these characteristics in our redenomination risk measure. They are very different over the three periods from the trading date.

Similarly, our spread measure might pick up differences in liquidity between sovereign bonds and the corporate bonds in our sample. In particular, one would expect that the premium for liquidity risk is strongest for bonds far away from maturity, not for bonds close to maturity (c.f. Longstaff et al., 2005; Covitz and Downing, 2007; Schwarz, 2019).

[19]See equation 2.1 for the details on how we define default risk in our bond pricing model.

However, the redenomination risk movements are strongest for the spread between the short end of the yield curve and not for the forward rates implied from the longer end of the curve.

Finally, we also take into consideration the fact that the bonds in our sample are governed by different legal rules. If some markets are "riskier" than others, our spreads might pick up these differences. In order to address this problem, we compare the first-year and second-year (forward) rates of corporate bonds issued by Dutch subsidiaries of German firms and bonds issued by French firms under English law. The differences are plotted in Figure 2.8. This analysis gives an idea of which order of magnitude legal risk should have in our redenomination risk measure. Most of the time, the differences are small and substantially below the estimated redenomination risks. In addition, the second-year spread is more volatile than the first-year spread, while this is the opposite for our redenomination risk measure. All this speaks against the hypothesis that our redenomination risk measures reflect primarily time variations in the premia for legal risk.

2.6 Conclusion

In this paper we document the term structure of redenomination risk. The term structure enables us to show how expectations of a euro area breakup have an influence on the short-run dynamics of sovereign bond yields. We identify redenomination risk by comparing yields of bonds that differ in jurisdiction. First, we estimate daily yield curves for default-risk-free sovereign bonds issued by France, Germany, and Italy under their respective domestic jurisdictions. These bonds will be redenominated in the case of an exit of the respective country from the euro area into the new currency the country issues. Second, we estimate daily yield curves for safe international corporate €-bonds issued by corporations in these countries under a foreign jurisdiction. This legal setting prohibits the redenomination of the bond into a newly issued currency.

Our results confirm that the ECB's fear of (self-fulfilling) expectations of a breakup disrupting its control over the short end of the yield curve and potentially leading to a real exit of the countries was a justified concern. At the peak of the crisis, market participants' expectations of a breakup had a considerable impact on the sovereign bond yields in the short end of the yield curve. Furthermore, we are able to show that the ECB was indeed able to reduce the implied redenomination risk for Italy, which was one of the countries hit by the crisis. Nevertheless, there was a downside to this: expectations that France and Germany would exit increased. This is hardly surprising in light of the fact that France and Germany as the two largest euro-area and non-crisis countries bear the fiscal risks of the ECB's unconventional policies.

2.A Estimation method

We apply the cubic spline method of McCulloch (1971, 1975) to estimate the default-risk-free discount rates $R_{t,c}^{dom}(\tau)$ and $R_t^{int}(\tau)$. Compared to the parametric alternatives, which specify a single functional form of the forward rates over the entire maturity domain, this approach models the instantaneous forward rate curve with piecewise cubic polynomials joined at predetermined knot points. This enables high flexibility of the estimation method, which is useful for our purpose because our goal is not to obtain a smooth representation of the yield curve, but rather to obtain an accurate measure of the riskless interest rate *and* the cost of redenomination risk in the prices of the euro-denominated bonds.

The following notation will be used. Let $p_{i,t}$, $i = 1, \ldots, K$ denote the observed dirty prices of K bonds at time t from which the term structure is to be inferred. Bond i has fixed payments, $c_i(\tau_j)$, where τ_j, $j = 1, \ldots, m_i$ are the coupon payment dates of a bond i with maturity m_i. The payment $c_i(\tau_j)$ consists of coupon and repayment of principal at maturity net of CDS premia. According to the bond pricing formula, the dirty price of a bond i is the discounted future cash flows of the bond until maturity:

$$p_{i,t} = \sum_{j=1}^{m_i} c_i(\tau_j) d_t(\tau_j). \tag{2.7}$$

where $d_t(\tau_j)$ is the discount factor of maturity τ_j and is identical for all bonds but can change over trading days t.

We estimate the unknown discount curve, $d(t_j)$ with a piecewise cubic spline model:

$$d_t(\tau_j, \beta) = 1 + \sum_{l=1}^{L} \beta_t^l g^l(\tau_j).$$

Here $g^l(t_j), l = 1, \cdots, L$, defines a set of piecewise cubic basis functions, which satisfy $g^l(0) = 0$. For $l < L$, the basis functions are defined as

$$g^l(\tau_j) = \begin{cases} 0 & \tau_j < q_{l-1}, \\ \frac{(\tau_j - q_{l-1})^3}{6(q_l - q_{l-1})} & q_{l-1} \leq m_{ij} < q_l, \\ \frac{(q_l - q_{l-1})^2}{6} + \frac{(q_l - q_{l-1}(\tau_j - q_l))}{2} + \frac{(\tau_j - q_l)^2}{2} - \frac{(\tau_j - q_l)^3}{6(q_{l+1} - q_l)} & q_l \leq \tau_j < q_{l+1}, \\ (q_{l+1} - q_{l-1}) \left[\frac{2q_{l+1} - q_l - q_{l-1}}{6} + \frac{\tau_j - q_{l+1}}{2} \right] & q_{l+1} \leq \tau_j. \end{cases}$$

These functions are twice-differentiable at each knot point to ensure a smooth and continuous curve around the points. For $l = 1$ we set $q_{l-1} = q_l = 0$ and for $l = L$, the basis function is given by $g^l(\tau_j) = \tau_j$.

The cubic-splines-based term structure estimation divides the term structure into segments with knot points. We use an L-parameter spline with $L - 1$ knot points q_l. For $1 < l < L - 1$ the knot points are defined as

$$q_l = \tau_h + \theta(\tau_{h+1} - \tau_h), \tag{2.8}$$

with $h = \frac{(l-1)K}{L-2}$ and $\theta = \frac{(l-1)K}{L-2} - h$. The first knot point is $q_l = 0$ and the last knot corresponds to the longest maturity of the bond in the sample. The number of basis functions n are set to the nearest square root of the number of observed bonds N. Cubic polynomial functions are then used to fit the term structure over these segments.

We can rewrite equation (2.7) in a vector notation

$$\mathbf{p}_t = \iota_m' \left(\mathbf{C} \cdot \mathbf{D}_t\right) + \epsilon_t, \tag{2.9}$$

where ι_m is an $m \times 1$ vector of ones, and $\cdot$ represents a element-wise multiplication of the matrices $\mathbf{C}$ and $\mathbf{D}_t$. m is the longest maturity of the sample. $\mathbf{C}$ is an $m \times K$ matrix, with $c_i(\tau_j)$ in cell i, j. Note that if bond i has no payment on date τ_j, $c_i(\tau_j) = 0$. The dirty prices of the K bonds are listed in the $1 \times K$ vector $\mathbf{p}$. The discount factor matrix, also n $m \times K$ matrix, is the weighted sum of the $l = 1, \ldots, L$ basis functions

$$\mathbf{D}_t = \iota_m \iota_K' + \beta_t^1 \mathbf{G}^1 + \cdots + \beta_t^L \mathbf{G}^L. \tag{2.10}$$

If we combine and rearrange equations (2.9) and (2.10), we obtain the following equation:

$$\begin{aligned} \underbrace{\mathbf{p}_t - \iota' \mathbf{C}}_{z_t} &= \underbrace{\beta_t^1 \iota' \mathbf{C} \cdot \mathbf{G}^1 + \cdots + \beta_t^L \iota' \mathbf{C} \cdot \mathbf{G}^L}_{\beta_t X_t} + \epsilon_t \\ \mathbf{z}_t &= \beta_t \mathbf{X}_t + \epsilon_t \end{aligned} \tag{2.11}$$

We estimate the unknown parameter vector β_t with weightes least squares (WLS) for each individual trading day separately. Since isseud volume differs across bonds, in some cases by a substantial amount. Since it could be that prices fluctuate more for low volume bonds, OLS might not be efficient. Therefore, we use the squared root of the issued volume of the bonds to weights the bonds differently.

2.B Data availability

In this section we provide a detailed report on how we collect our data set. In particular, we show how we come to the decision of using both sovereign and corporate bonds for our analysis, even though the comparison of two different bond types can be problematic. In addition, we also explain why we do not analyze redenomination risk of other euro area countries such as Greece, Portugal, and Spain, which were also exposed to very high exit expectations during the European sovereign debt crisis.

2.B.1 Why both corporate and sovereign bonds?

For each country we need data on securities that only differ in jurisdiction. In the optimal case, all other characteristics of the securities should be equal. Therefore, we only consider bonds that have a fixed coupon rate, are euro denominated, and neither callable nor guaranteed. In addition, it is important to have a sufficient amount of bonds to precisely estimate the yield curve non-parametrically.

We collect our data set using Bloomberg and Datastream. As a start, we search only for sovereign bonds that are either governed under domestic law or foreign law. If we are able to obtain a sufficient amount of domestic-law *and* foreign-law sovereign bonds for each country, we would be able to estimate yield curves for these two groups and identify redenomination risk. However, for neither of the three countries we were able to find sovereign bonds that are denominated in euro *and* governed under a foreign law. Note that we were able to find data on Italian government bonds under US law and denominated in US Dollars (USD). However, we are not able to compare euro-denominated bonds with USD-denominated bonds since we would capture, additional to redenomination risk, exchange rate risk between the new currency towards USD in case the country exits the Eurozone.

As a next step, we examine whether we are able to execute our identification strategy by using only corporate bonds. We search for non-financial corporate bonds of France, Germany, and Italy that are governed under different jurisdictions. However, we again have the same problem as with government bonds, only that the corporations in our sample mostly issue foreign-law bonds but do not have a sufficient amount of domestic-law bonds. Therefore, we are not able to precisely estimate the yield curve non-parametrically using only corporate bonds. Consequently, we decide to take advantage of the fact that we have sufficient amount of data on (i) government bonds that are issued under domestic law and (ii) corporate bonds that are issued under foreign law for all three euro area countries and combine these two data sets.

2.B.2 Why not other countries?

Besides France, Germany, and Italy it would be very interesting to identify redenomination risk of other euro area countries such as Greece, Portugal, and Spain, which were exposed to high exit expectations during the peak of the European sovereign debt crisis. Nevertheless, we decide not to include these countries due to the following reasons.

First, for these three countries it is difficult to find sufficient amount of data on bonds of our interest. For instance, even though they have some foreign-law government bonds that are denominated in euro, the amount is not enough to non-parametrically estimate a yield curve. For Greece, we were able to find two government bonds under English law, for Portugal and Spain only one under English law. The data availability of corporate bonds is also problematic. Especially for Greece, we were not able to find a single non-financial corporation that issues foreign-law corporate bonds *and* has actively traded CDS. For Portugal, we find in total 22 corporate bonds issued by Energias de Portugal and 3 bonds of Altice Portugal. For Spain we find more corporate bonds that fulfill our requirements: in total, we find one bond issued by Abertis, 13 bonds by Iberdrola International B.V., and 27 bonds issued by Telefonica Emisiones, S.A.U..

Second, for our identification strategy it is important to distinguish between G7 and non-G7 countries due to the ISDA Master Agreement of 2002 that governs CDS contracts. It states that the CDS contracts must pay if there is "any change in the currency and composition of any payment of interest or principal to any currency which is not a Permitted Currency. "Permitted currency" means (1) the legal tender of any Group of 7 country, or (2) the legal tender of any country which has a local currency long-term debt rating of either Aaa or higher". Therefore, if we apply our identification strategy on non-G7 countries, by subtracting the CDS premia from domestic-law bonds we would also get rid of redenomination risk.

This, in first sight, should not be a problem with our identification strategy. This would mean that we would obtain (i) a riskless discount rate when estimating a yield curve of CDS-adjusted domestic-law government bonds and (ii) a discount rate *over-compensated* for redenomination risk when estimating a yield curve of CDS-adjusted foreign-law corporate bonds. Taking the ratio would also give us a measure that captures redenomination risk. However, in order for the corporate CDS contracts to measure default *and* redenomination risk as given in the ISDA Master Agreement 2002, the corporations of the CDS contracts must have euro-denominated local debt. Otherwise they will not cover for redenomination risk. However, we were not able to confirm this for all corporations of Portugal and Spain[20]

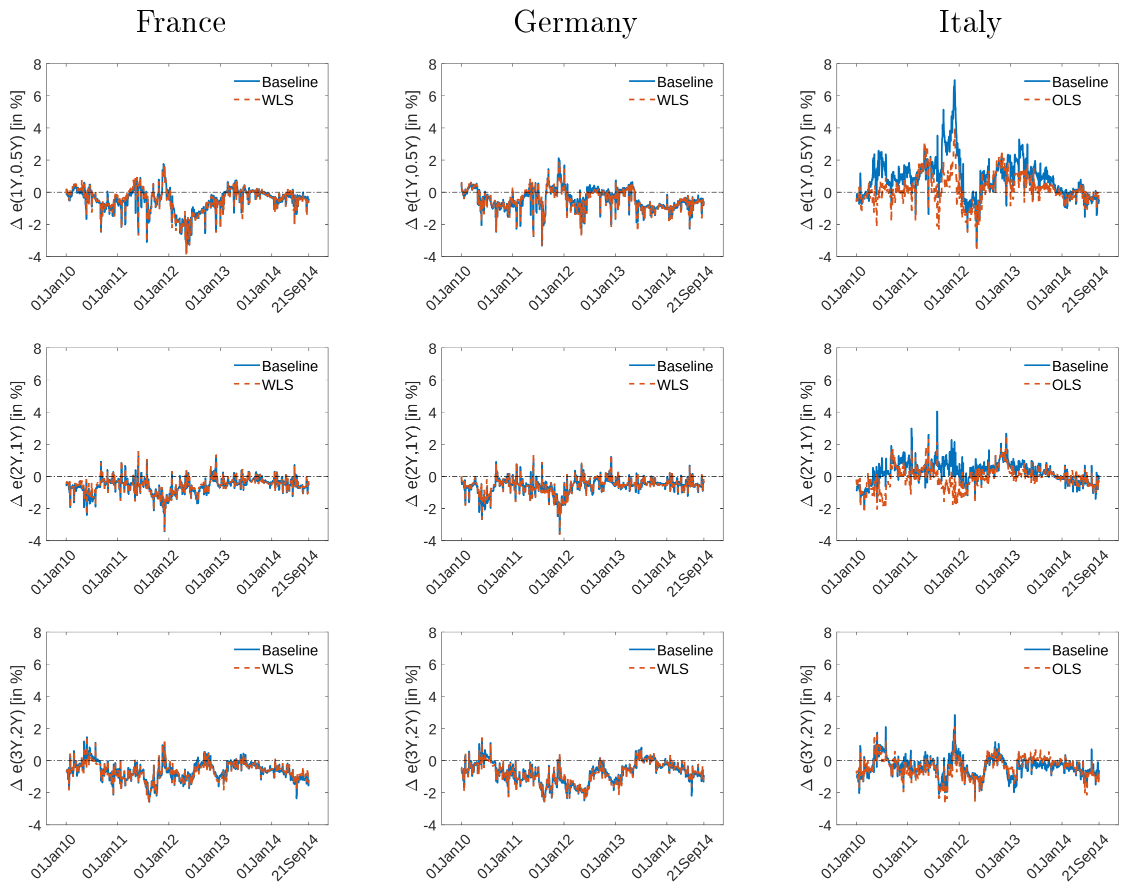
[20]Krishnamurthy et al. (2018) show that Telefonica Emisiones and Energias de Portugal do have outstanding local-law euro-denominated debt.

2.C Robustness and further results

2.C.1 Yield curve estimation with ordinary least squares

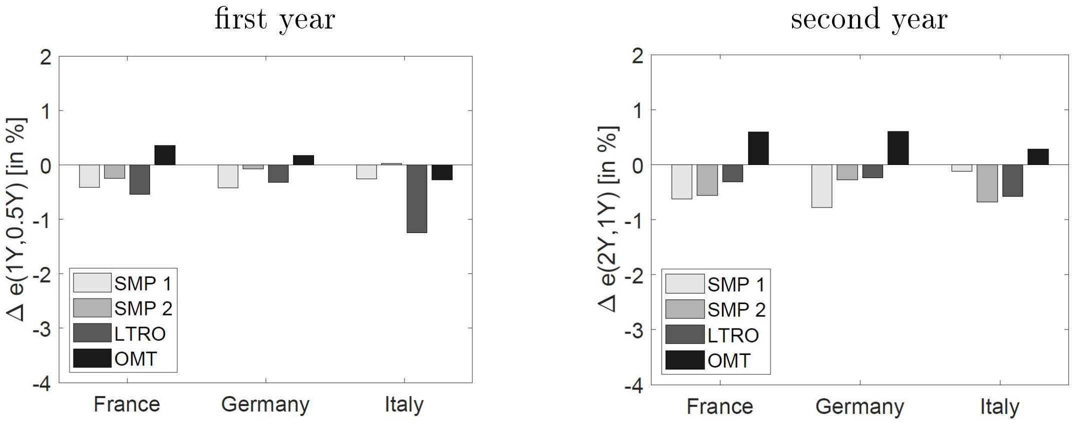
Our baseline model estimates the yield curve by weighted least squares by using the issue volume of bonds as weights. As a robustness check, we also estimate the yield curve by ordinary least squares. The results do not differ much from our baseline treatment; see Figures 2.9 and 2.10.

Figure 2.9: Expected exchange rate movement, first, second and third year from trading date, OLS



Notes: See Figure 2.11 for baseline.

Figure 2.10: Changes in expected exchange rate changes (20-day window, OLS)



Notes: See Figure 2.3. The effect is computed from the OLS estimates displayed in Figure 2.9. For bootstrapped standard errors and significance, see Table 2.5 in Appendix 2.D.

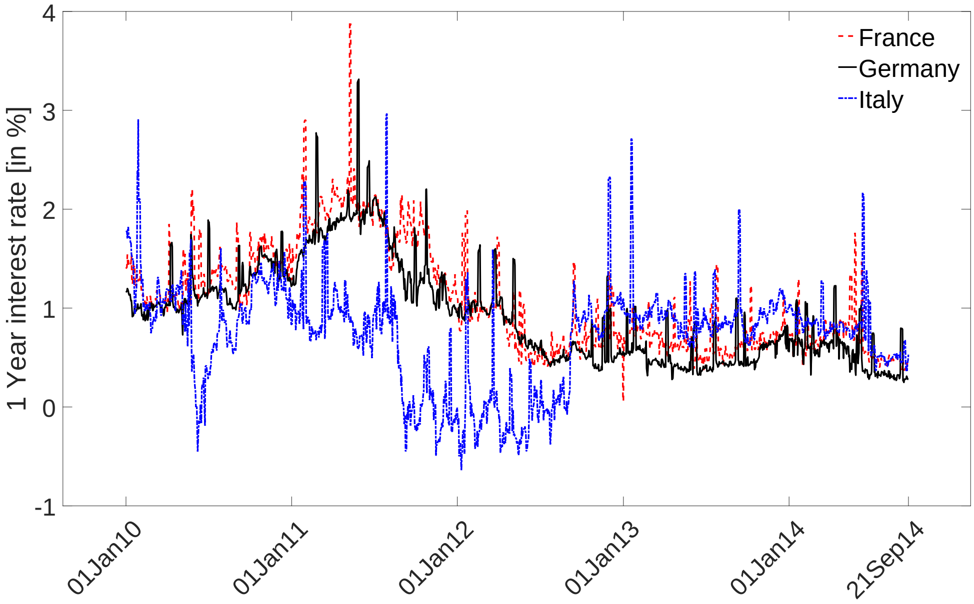
2.C.2 The safe international corporate €-bond yield

The assumption that the spread between the yield on the safe international corporate €-bond and the default-risk-free sovereign bond is only redenomination risk is key for our approach. Two potential issues might arise with respect to the safe international corporate €-bond in this respect. First, it might be that markets for Italian, French and German bonds are segmented along the geographical dimension. Second, it might be that the corporate bonds issued under German law still pick up German redenomination risk.

We address the first issue by estimating safe international corporate €-bond yields for French, German, and Italian corporations separately. By contrast, our baseline approach pools all international corporate €-bonds of companies from all three countries. If markets are efficient and not geographically segmented, the separate safe yields should be identical across countries and differences in the estimated yields should only result from estimation uncertainty. Figure 2.11 represents the country-specific safe corporate €-bond yield estimate of France, Germany, and Italy. French and German bonds show very similar yields over the entire horizon. For Italian safe corporate €-bonds, i.e., after CDS premia, the yields are *lower* than the French and German yields between mid-2010 and mid-2012. This means that we cannot exclude the possibility that markets are regionally separated and that at the height of the Italian crisis some of the decreased demand for Italian sovereign bonds flows into corporate bond demand. However, compared to the default-risk-free sovereign/safe international corporate €-bond spread, the differences between the safe €-bond yield and corporate bonds issued in the three countries are minor.

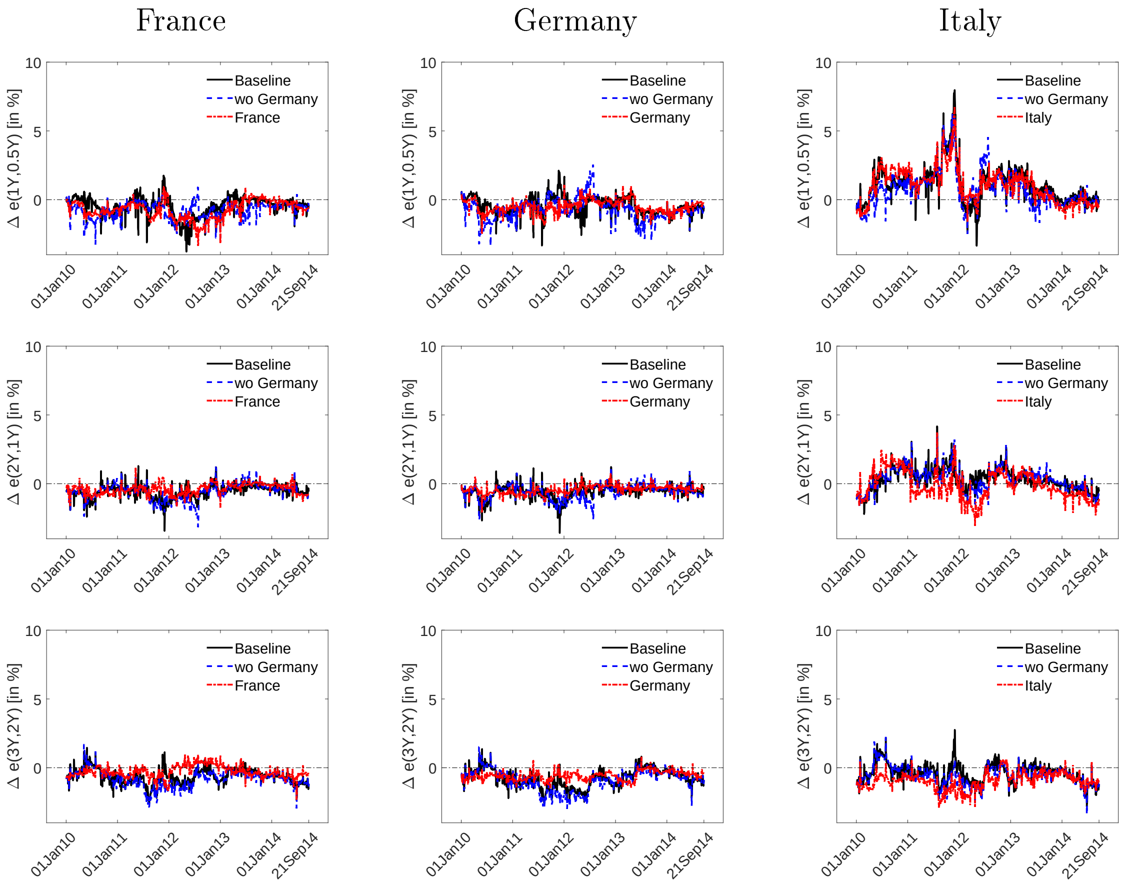
For redenomination risk, this means that our baseline estimate for Italy *underestimates* redenomination risk during this period. Nevertheless, the estimated time series for redenomination risk hardly changes; see Figure 2.12. Finally, we also show that the choice of the window around the ECB's intervention is not key for our results. We can alternatively use a 5-day window as well and obtain qualitatively the same results; see Figure 2.13. For completeness, Figure 2.14 displays all five estimates of the safe international corporate €-bond yield.

Figure 2.11: One-year yield on safe corporate €-bonds by country of issuer



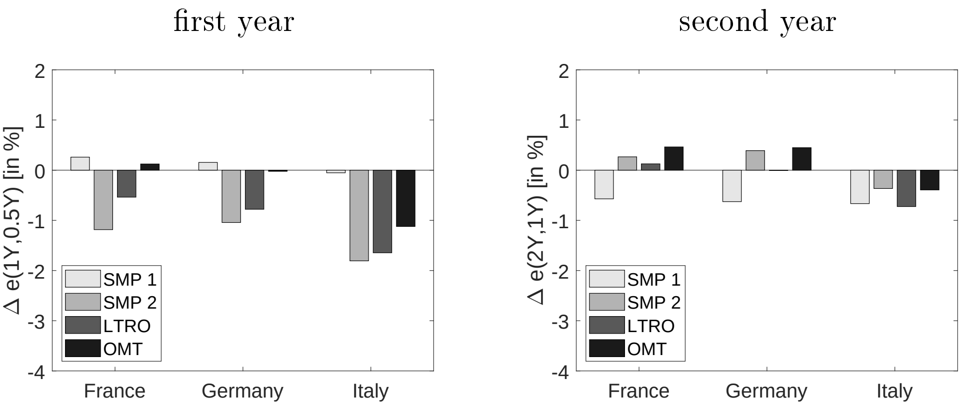
Notes: One-year yield estimates from CDS-insured corporate bonds issued by a corporation in one country under another country's jurisdiction. Bonds are grouped by the country in which the parent company is incorporated if the bond is issued by a subsidiary. All French and Italian bonds are issued under English law, and all German bonds are issued under German law by subsidiaries in the Netherlands.

Figure 2.12: Different specifications for redenomination risk measures



Notes: Expected changes in the exchange rate as implied by (2.6) for the estimated yield curves for CDS-insured sovereign bonds and CDS-insured international corporate €-bonds. The first row gives the annualized expected exchange rate movements between the 183th day after the trading day and the 365th day after. The second and third rows display the expected exchange rate movement between the 366th and the 730th day and between the 731st and the 1095th day after the trading day, respectively. In short, the rows display the expected exchange rate movement for the first, second and third year after the trading day. The black solid line is our baseline estimates, the blue dashed line replaces the €-bond yield with an estimate excluding German corporate bonds and the red dotted line uses local corporate €-bonds instead.

Figure 2.13: Expected change in exchange rate (5-day window)



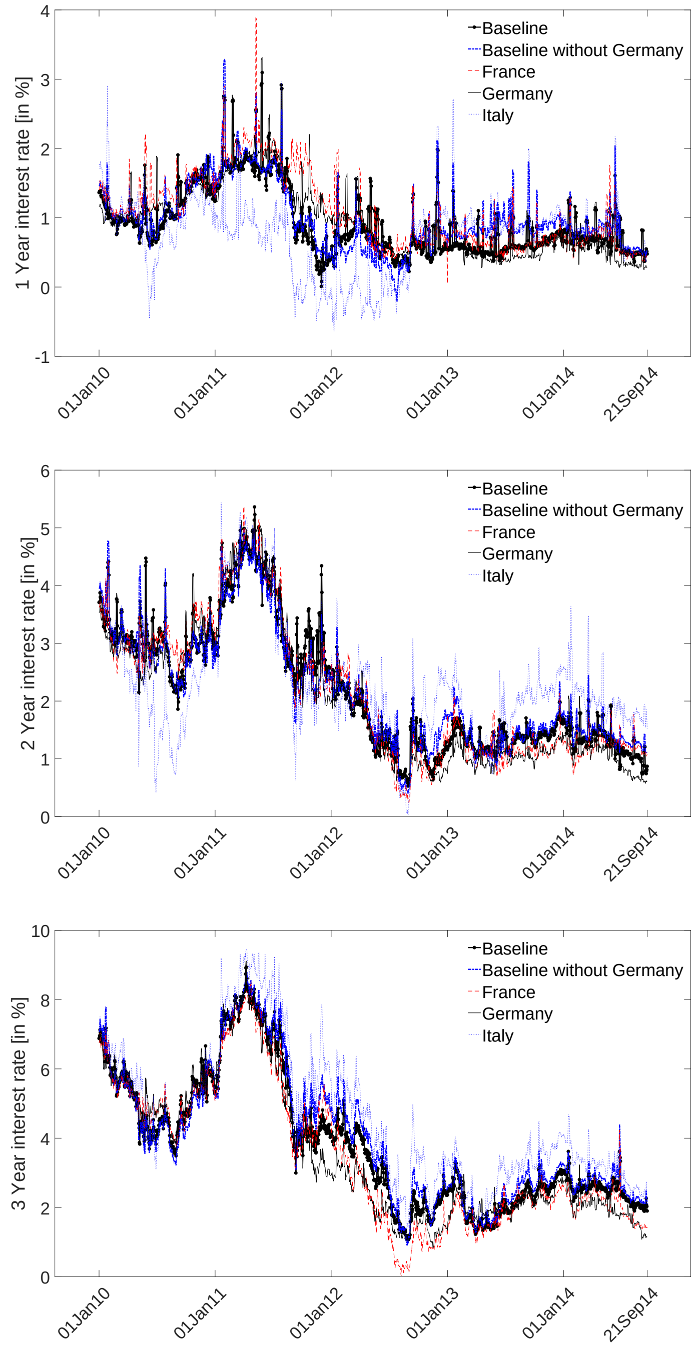
Notes: The graphs display the average change in redenomination risk after an ECB intervention as in Figure 2.3 but using an average over 5 trading days before and after the intervention. For bootstrapped standard errors and significance, see Table 2.6 in Appendix 2.D.

Figure 2.14: The riskless interest rates, five measures

1-Year

2-Year

3-Year



Notes: The figure shows the baseline and four alternative measures of the safe corporate €-bonds rate: Excluding German law bonds, WLS estimate, and the three rates implied by estimating separately by country.

2.D Significance of policy interventions

Table 2.4 reports the bootstrapped significance levels of the results displayed in Figure 2.3. We draw for each day 5000 bootstrap replications from the set of sovereign and corporate bonds (stratified by country) and estimate the corresponding yield curve. On that basis we calculate the average effect of the intervention for each bootstrap replication. Tables 2.5 to 2.6 report the bootstrapped standard deviations and significance levels for Figures 2.10 and 2.13. Table 2.7 report the results for Figure 2.7.

Table 2.4: Expected changes in the exchange rate around intervention dates

		183th day to 365th day		1st year to 2nd year	
Event	Country	Estimate	Std. Error	Estimate	Std. Error
SMP-1	France	−0.44*	0.23	−0.51***	0.20
	Germany	−0.82***	0.23	−0.71***	0.19
	Italy	−0.15	0.61	−0.01	0.44
SMP-2	France	−0.41	0.28	−0.57**	0.20
	Germany	0.21	0.29	−0.17	0.20
	Italy	−0.43	0.75	−0.91	0.51
LTRO	France	−0.36**	0.28	−0.24	0.24
	Germany	−0.19	0.29	−0.13	0.25
	Italy	−1.44***	1.22	−0.83**	0.88
OMT	France	0.30	0.25	0.54***	0.14
	Germany	0.15	0.29	0.56***	0.15
	Italy	−0.83*	0.60	−0.09	0.32

Notes: The table shows the expected changes in the exchange rate within a 20-day window around four ECB events. Confidence intervals are calculated with bootstrapping (5000 replications). One asterisk indicates significance at the 10% level, two asterisks indicate significance at the 5% level, and three asterisks indicate significance at the 1% level.

Table 2.5: Expected changes in the exchange rate around intervention dates (spread estimated with OLS)

		183th day to 365th day		1st year to 2nd year	
Event	Country	Estimate	Std. Error	Estimate	Std. Error
SMP-1	France	-0.42^{**}	0.20	-0.62^{***}	0.19
	Germany	-0.42^{**}	0.20	-0.78^{***}	0.19
	Italy	-0.26	0.20	-0.12	0.43
SMP-2	France	-0.25	0.27	-0.55^{***}	0.20
	Germany	-0.08	0.28	-0.27	0.20
	Italy	0.02	0.41	-0.68^{**}	0.53
LTRO	France	-0.54^{**}	0.26	-0.31	0.24
	Germany	-0.32	0.24	-0.24	0.24
	Italy	-1.25^{**}	0.73	-0.58	0.92
OMT	France	0.36	0.25	0.60^{***}	0.14
	Germany	0.17	0.28	0.61^{***}	0.15
	Italy	-0.28	0.53	0.29	0.32

Notes: The table shows the expected changes in the exchange rate within a 20-day window around four ECB events. Confidence intervals are calculated with bootstrapping (5000 replications). One asterisk indicates significance at the 10% level, two asterisks indicate significance at the 5% level, and three asterisks indicate significance at the 1% level.

Table 2.6: Expected changes in the exchange rate around intervention dates (5-day window)

		183th day to 365th day		1st year to 2nd year	
Event	Country	Estimate	Std. Error	Estimate	Std. Error
SMP-1	France	0.33	0.40	−0.48	0.38
	Germany	0.23	0.46	−0.54	0.39
	Italy	0.08	1.17	−0.54	0.47
SMP-2	France	−1.10**	0.47	0.19	0.32
	Germany	−0.90**	0.50	0.31	0.33
	Italy	−1.44	1.49	−0.31	0.40
LTRO	France	−0.50	0.56	0.00	0.55
	Germany	−0.73	0.64	−0.12	0.56
	Italy	−1.27	2.67	−0.65	1.08
OMT	France	0.18	0.45	0.46*	0.27
	Germany	0.05	0.53	0.45	0.29
	Italy	−1.22	1.173	−0.49	0.49

Notes: The table shows the expected changes in the exchange rate within a 5-day window around four ECB events. Confidence intervals are calculated with bootstrapping (5000 replications). One asterisk indicates significance at the 10% level, two asterisks indicate significance at the 5% level, and three asterisks indicate significance at the 1% level.

Table 2.7: Expected changes in the exchange rate around FOMC meetings

	183th day to 365th day		1st year to 2nd year	
Event	Estimate	Std. Error	Estimate	Std. Error
QE-2	-0.32^{**}	0.12	-0.1	0.09
Twist	-0.30^{**}	0.13	0.03	0.09
QE-3	-0.15^{*}	0.09	-0.02	0.07
Hike-1	0.23^{*}	0.10	0.05	0.06
Hike-2	-0.07	0.11	-0.08	0.07
Hike-3	0.09	0.09	0.01	0.06

Notes: The table shows the expected changes in the exchange rate within a 20-day window around six FOMC meetings. Confidence intervals are calculated with bootstrapping (5000 replications). One asterisk indicates significance at the 10% level, two asterisks indicate significance at the 5% level, and three asterisks indicate significance at the 1% level.

CHAPTER 3

The Effect of Monetary Policy on Stock Market Investment Decisions: The Role of Gender and Marital Status[1]

Caterina Forti Grazzini and Chi Hyun Kim

We use the Panel Study of Income Dynamics (PSID) household survey data from 2001-2017 to investigate whether monetary policy has heterogeneous effects on the financial portfolio decisions of different household groups that differ in their head's gender and marital status. On the one hand, we show that a monetary policy shock affects the stock market entry decisions of single female-headed households, while it does not impact single- nor married male-headed households. On the other hand, monetary policy has no significantly different effects on exit decisions nor stock market investment rebalancing choices. These results suggest that monetary policy does not have a heterogeneous impact on portfolio decisions along gender and marital status, when single female-headed households participate in the stock market.

Keywords: monetary policy, gender, stock market participation, portfolio choices
JEL classification: E58, J16, G11

[1]We thank Franciska Bremus, Marco del Negro, Alexander Kriwoluzky, Dieter Nautz, seminar participants at the 2019 and 2020 Time Series Workshop in Tornow, at the 20th IWH-CIREQ-GW Macroeconometric Workshop in Halle and three anonymous referees for helpful comments and suggestions on the earlier versions of the paper.

3.1 Introduction

The primary mandate of major central banks is to maintain price stability, which is the reason why central bankers have traditionally paid less attention to the distributional impact of their policy measures on inequality. However, in the aftermath of the Global Financial Crisis, economic inequality in industrialized countries increased drastically and the public raised concerns that the long-enduring low interest rate environment is exacerbating this problem, since low interest rates might only benefit certain groups of households (Bivens, 2015). One key aspect of this debate is to understand how financial and demographic characteristics of households interact with monetary policy. Several papers document monetary policy's heterogeneous effects along income, wealth, house ownership, and employment status of households (see, among others, Adam and Tzamourani, 2016; Ampudia et al., 2018; Wong, 2019).

In this chapter, we take a different perspective and evaluate how *gender and marital status* interact with monetary policy. By doing so, we specifically focus on single female-headed households in contrast to both single and married male-headed households. Insights of feminist economics show how traditional monetary policy, in combination with finance-dominated capitalism, may favor men at the expense of women (Bakker, 1994; Van Staveren, 2014a,b; Young, 2018). In particular, several studies analyze the role of gender and marital status on financial decisions and show that single female-headed households belong to one of the most fragile groups in society: Married women have a higher propensity to invest in risky assets than single women, while the marital status gap does not apply to men (Bertocchi et al., 2011); Single women are more risk averse in their financial decisions than single and married men (Sung and Hanna, 1996; Sunden and Surette, 1998; Jianakoplos and Bernasek, 1998); Barber and Odean (2001) find that differences in portfolio turnover and net return performance are larger between the accounts of single men and single women than between the accounts of married men and married women. All these characteristics of single women in combination with a prolonged period of low interest rates may lead to a distributional divergence between this household group and the rest of the population.

Using the Panel Study of Income Dynamics (PSID) household survey data, we empirically investigate the effect of monetary policy on households' stock market investment behavior, thereby focusing on (i) entry, (ii) exit, and (iii) the active investment choices of single female- and both married and single male-headed households in a separate manner. One empirical challenge lies on the identification of US monetary policy shocks. We first identify monetary policy shocks of the Federal Reserve (FED) at a daily frequency following the high frequency identification method proposed by Nakamura and Steinsson (2018). As a next step, we match the frequency of our monetary policy shock measures with those of household survey data by aggregating the daily monetary policy shocks into a series with biennial frequency. By doing so, we consider the month of the year in which each household answers the survey questions. This allows us to construct an idiosyncratic biennial monetary policy shock series for each household. Finally, we improve our identification by exploiting households' heterogeneity in financial wealth and gender, which influences their exposure to monetary policy shocks.

Results can be summarized as follows. First, only single women are affected by monetary policy with respect to stock market entry choices, even after we control for socio-economic characteristics that are correlated with gender, marital status, and financial wealth. After a contractionary monetary policy shock, single female-headed households are 11% less likely to enter the stock market, while male-headed households' entry decisions are not affected. To the contrary, exit decisions and active stock investment choices do not differ across the different household groups after a monetary policy shock. These results suggest that female- and male-headed households do not behave differently once they both participate in the stock market. These results are in line with the literature showing that (i) women, who do not participate in the stock market, have the highest risk aversion in society, and (ii) women generally react more strongly to economic events that negatively affect their wealth (Jianakoplos and Bernasek, 1998; Barber and Odean, 2001; Fisher and Yao, 2017).

In order to visualize the impact of monetary policy on single female-headed households, we conduct a static simulation exercise to calculate how much financial wealth they potentially missed out on (or gained) due to their entry choices driven by monetary policy. Our exercise suggests that between 2001 and 2017, single female-headed households lost more than $2 billion, suggesting that monetary policy has a sizeable impact. Indeed, US households' investment decisions are crucial not only for financial security during the working life, but also for retirement. The lower propensity of single women to invest in risky assets like stocks could translate into large differences in the accumulation of financial wealth for retirement. Combined with lower earnings, lower savings, longer life spans, and higher risk aversion, this implies that single female-headed households are more likely to be living in poverty (Cawthorne, 2008). According to *Statista* - a data portal of households - in 2018, there were about 15 million US single female-headed households between 25 and 65 years old, representing almost 12% of total US households. Nonetheless, the distribution of poverty across the US society is skewed towards single women, with 24.9% of them having a family income below the poverty line, compared to 12.7% of single male-headed households and only 4.7% of married couples (2018 data). Thus, given the prominence of monetary policy since the GFC, understanding how it influences single women's financial wealth accumulation is crucial.

Our study contributes to the growing literature that uses micro-level data on the composition of households' wealth and income to evaluate the heterogeneous effects of monetary policy and its impact on wealth inequality. Bivens (2015), Domanski et al. (2016), Lenza and Slacalek (2018), and Ampudia et al. (2018) focus on unconventional monetary policy tools and conduct empirical reduced-form simulation exercises. They quantify the distributional effects of monetary policy through the valuation of asset prices by examining households' financial portfolio structure. They show that unconventional monetary policy disproportionately benefits households at the top wealth distribution. The same result is reached by Adam and Tzamourani (2016) in the context of conventional monetary policy. In addition, a new strand of literature investigates the effect of interest rate changes on the active risk-taking behavior of private investors, finding that investors' risk appetite increases if monetary policy is loosened (Lian et al., 2018; Daniel et al., 2018; Forti Grazzini, 2020). Our analysis provides new insights to this literature by examining gender and marital status as an additional source of household heterogeneity that might interact with monetary policy. Young (2018) is the first to formulate potential mechanisms through which unconventional monetary policies can affect gender wealth inequality. However, the study only provides descriptive results. Our study complements her arguments by providing a structural analysis.

The remainder of the paper is structured as follows. Section 3.2 describes the data and the construction of our final data set. Section 3.3 discusses the identification of monetary policy, while section 3.4 outlines the empirical framework and the results. In section 3.5, we provide a simulation study to calculate the impact of our results on the capital gains/losses of women through monetary policy. Section 3.6 concludes.

3.2 Data

3.2.1 The Panel Study of Income Dynamics

We use the Panel Study of Income Dynamics (PSID) survey data, which is a nationally representative longitudinal study of US families and their offspring over time. In the PSID, the unit of observation is the *household*, which is defined as a group of people living together as a family. Besides a broad range of socio-economic variables - such as gender, age, marital status, number of children, etc. - the PSID also provides rich information on the households' financial wealth and portfolio composition.

With respect to the financial portfolio volume and composition, households are asked to report information on their holdings of three broad asset classes: (i) *stocks* (shares of stock in publicly held corporations, mutual funds, and investment trusts); (ii) *riskless assets* (checking and savings accounts, money market funds, certificates of deposits, savings bonds, treasury bills); and (iii) *other assets* (bond funds, cash value in a life insurance policy, a valuable collection for investment purposes, or rights in a trust or estate). While the PSID specifically inquiries about the purchases or sales of stocks by households, for the riskless asset class it does not. Finally, although provided by the PSID, we do not include any assets held in employer-based pensions or IRAs.[2]

The PSID survey is of biennial frequency and we include waves from 2001 to 2017. One important feature of the PSID data is that the interviews happen every second year (in odd years) between March and December, with the responses to questions regarding wealth referring to the month in which the interview takes place. For questions regarding income, however, the households are asked to report their annual income of the previous year. This implies that data on income and wealth are not perfectly aligned. Nevertheless, for our analysis, this does not constitute an issue, as we mainly focus on the wealth variables, which are all measured at the same point in time for any given household. Finally, to make magnitudes comparable over time, we deflate all income and wealth data by the consumer price index (CPI) into December 2007.

[2]We exclude investments in retirement accounts because there is little trading in these accounts (Ameriks and Zeldes, 2004). Furthermore, the liquidity and payoff properties of retirement accounts are different from direct stock ownership (Haliassos and Bertaut, 1995).

3.2.1.1 Definition of household groups

The focus of the paper is to understand how different household groups react to monetary policy. These groups are identified with a combination of their household head's gender (female/male) and marital status (single/married). The choice of keeping the household as the unit of observation for the analysis, instead of only the household head, is driven by the impossibility to identify the family member who is in charge of financial decisions in married families. In fact, in the PSID, by default the term head refers to the husband in a heterosexual married couple, irrelevant of whether it is the husband who makes financial decisions for the family unit or not. Thus, it is not trivial to identify who is in charge of financial decisions. Moreover, even if we were able to identify the financially responsible person in a married couple, we still cannot rule out the possibility that married couples tend to make joint investment decisions (Sunden and Surette, 1998; Barber and Odean, 2001; Agnew et al., 2003).

The groups included in the analysis are (i) *single female-headed households* (SFHHs), (ii) *male-headed households* (MHHs) - which include both single and married households -, and (iii) *married male-headed households* (MMHHs). Since we are interested in the behavior of single female-headed households, throughout the paper we compare single female- with male-headed households and married male-headed households, using the former as our baseline comparison.[3] Finally we identify *stock market participants* as the group of households investing at least $1 dollar in the stock market in two consecutive waves.

3.2.1.2 Construction of the relevant variables

In the first part of the analysis, we focus on households' dynamic stock market participation choices. We construct two binary variables, $Entry_{i,t}$ and $Exit_{i,t}$, which visualize stock market entry or exit decisions. $Entry_{i,t}$ equals one if household i has zero stock market investment in $t-1$, but positive investment in t, and zero if stock investment is null in both waves. $Exit_{i,t}$ is equal to one if household i owns stocks in $t-1$, but does not in t, and zero if the household owns stocks in both waves.

In the second part of the analysis, we focus on stock market participants to analyze their stock market portfolio choices. To do so, we need to decompose the change in the amount of held stocks held into an active investment/disinvestment component and a passive capital gains/losses component. In every wave, the PSID asks subjects to report on the amount of stocks and mutual funds bought and/or sold during the time since the previous wealth survey. We use this information to calculate the *stock active saving* as the sum of all stocks sold or/and purchased between $t-1$ and t.

In addition, we calculate riskless asset active saving. As the PSID does not include information on purchases or sales of this asset class, we proceed as follows. First, we approximate the capital gain/loss on this asset class between $t-1$ and t with the return

[3]Unfortunately, it is not possible to directly compare single female- vs single male-headed due to the limited number of observations of the latter group.

on the 1-year US Treasury.[4] Second, we subtract it from the change in the amount of riskless assets held between $t-1$ and t (provided by the PSID) to extract the *riskless asset active saving*, which is the amount of riskless assets sold or bought between $t-1$ and t.

3.2.1.3 Sample selection

We only include households that participate in the survey for at least three consecutive waves.[5] We exclude households where the age of the head is younger than 25 years or older than 65 years. Additionally, we only consider households, where the marital status of the head does not change throughout the sample. We also control for possible mismatches in the reported answers and eliminate households that do not report consistent data. In particular, we discard households that (i) declare not to have stocks, but then report a positive value of stock wealth; (ii) indicate a negative value of stock wealth; or (iii) declare a non-zero active saving, but at the same time report zero purchases or sales of assets. Moreover, we trim all wealth variables at the 1% level to mitigate the impact of outliers. Finally, we use sample weights provided by PSID when producing the summary statistics, but we do not weight observations in the regression analysis as it would be inefficient (Deaton, 1997).[6]

3.2.2 Summary statistics

Table 3.1 provides some household-level summary statistics of our final data set, pooling all waves. Panel A reports the statistics for all households that satisfy all minimum requirements to be included in the analysis (the full sample). Panel B shows the summary statistics of stock market participants. We compare single female-headed households versus households, where the head is a man (both single and married). Summary statistics are calculated using sample weights provided by the PSID.[7]

With regard to all households in our sample (Panel A), 29% of male-headed households participate in the stock market between 2001 and 2017. In contrast, only 16% of single female-headed households invest in the stock market. In addition, single female-headed households seem to display higher risk aversion, as their rate of stock market entry is lower than for male-headed households. On average, the value of male-headed family's financial portfolio sums up to more than \$100,000, which is three times higher than their female counterpart. Moreover, stock holding of male-headed households is about four times higher than woman's, implying that the former prefers riskier financial investments (with a 17% of their portfolio invested in risky assets, compared to the female 9%).

[4] As a robustness check, we also use the 2-year and the 10-year Treasury to approximate the return on the riskless asset capital gains.

[5] This choice is driven by the large number of households appearing in the PSID for only one wave. However, results do not change if we relax this constraint; see Table 3.9 in the appendix.

[6] There are two sets of explanations for the choice of not using sample weights in this context. First, as we control for outliers and trim the data set at the cross section, we reduce the representativeness of the weighted data. The second issue arises from the usage of panel data. When applying fixed effect estimation, it is not possible to assign different weights over time for the same family unit.

[7] For an overview of the unweighted summary statistics, see Table 3.7 in the appendix.

Table 3.1: Summary statistics

	Mean	SD	Mean	SD	Mean	SD
	All households		Single female-headed HHs		Male-headed HHs	
Panel A: Full sample						
Stock market participation	0.27	0.44	0.16	0.37	0.29	0.45
Stock market exit	0.30	0.46	0.34	0.47	0.30	0.46
Stock market entry	0.10	0.30	0.06	0.24	0.11	0.31
Stocks	55056.19	627113.15	14580.03	106326.21	63486.96	687198.74
Riskless assets	27205.67	90301.63	13009.60	30310.10	30128.64	98018.67
Stock active saving	235668.44	12375997.76	67.43	2235.77	284996.11	13609611.55
Riskless asset active saving	5882.27	88949.70	2403.31	29122.42	6595.02	96837.34
Liquid assets	96771.38	702582.00	35212.30	166566.79	109567.22	767798.25
Stock/liquid assets	0.15	0.30	0.09	0.24	0.17	0.31
Net worth	365631.10	1333478.63	218791.56	1550873.64	396103.14	1283039.44
Income	92487.03	123296.94	46927.72	39714.12	101881.37	132306.03
Home ownership	0.84	0.37	0.68	0.46	0.87	0.33
Observations	14807		2105		12673	
Panel B: Stock market participants						
Stocks	253891.93	1404329.77	113215.46	299449.61	269248.76	1474872.99
Riskless assets	57592.73	157731.34	36930.01	59835.86	59849.30	164785.85
Stock active saving	3387.74	30543.59	114.58	6042.69	3749.08	32103.14
Riskless asset active saving	12960.12	157328.94	7205.88	53615.60	13588.52	164736.41
Liquid assets	342547.85	1548857.51	175188.84	421582.63	360820.18	1624224.00
Stock/liquid assets	0.60	0.29	0.58	0.30	0.60	0.29
Net worth	878261.14	2234640.50	1090033.29	4460878.19	855101.66	1834792.98
Income	141347.15	203069.50	68378.32	84711.07	149304.05	210512.67
Home ownership	0.93	0.25	0.86	0.35	0.94	0.24
Observations	2389		162		2227	

Notes: The table shows the summary statistics of the relevant wealth and income variables included in the analysis. Panel A presents the results for the full sample; Panel B for the sub-sample of stock market participants (at least \$1 invested in stocks in both $t-1$ and t). The sample period is 2001-2017. Household's observations are weighted by the longitudinal weights provided by the PSID. *Source*: PSID and own calculations.

The picture is slightly different when considering stock market participants (Panel B). Although single women still invest a lower share of their net worth in risky assets, the financial gap partially closes, with the composition of the financial portfolio being the same across groups (roughly 60% invested in risky assets). When taking a closer look, however, we see that the net worth ratio between female-headed households in the participating and full sample groups is roughly 5, while that of male-headed households is around 2.3. This result suggests that the former might require more wealth to bear risk to invest in risky assets, in turn, more risk averse. The ratios between SFHHs and MHHs that hold stocks in Panel A and B (7.6 and 4.4, respectively) point in the same direction.

3.3 Identification of monetary policy shocks

The identification of the effects of monetary policy on the investment behavior of households poses several challenges. First, it is crucial to obtain exogenous monetary shocks. Second, it is necessary to combine the monetary policy shocks with biennial data on household financial and investment characteristics that we obtain from the PSID. Third, we have to overcome the issue that the identified monetary policy shocks are of small size and transitory nature, which can pose challenges on the estimated responses of our variable of interest. Fourth, we need to take into account the household head's gender/marital status.

3.3.1 High frequency identification method

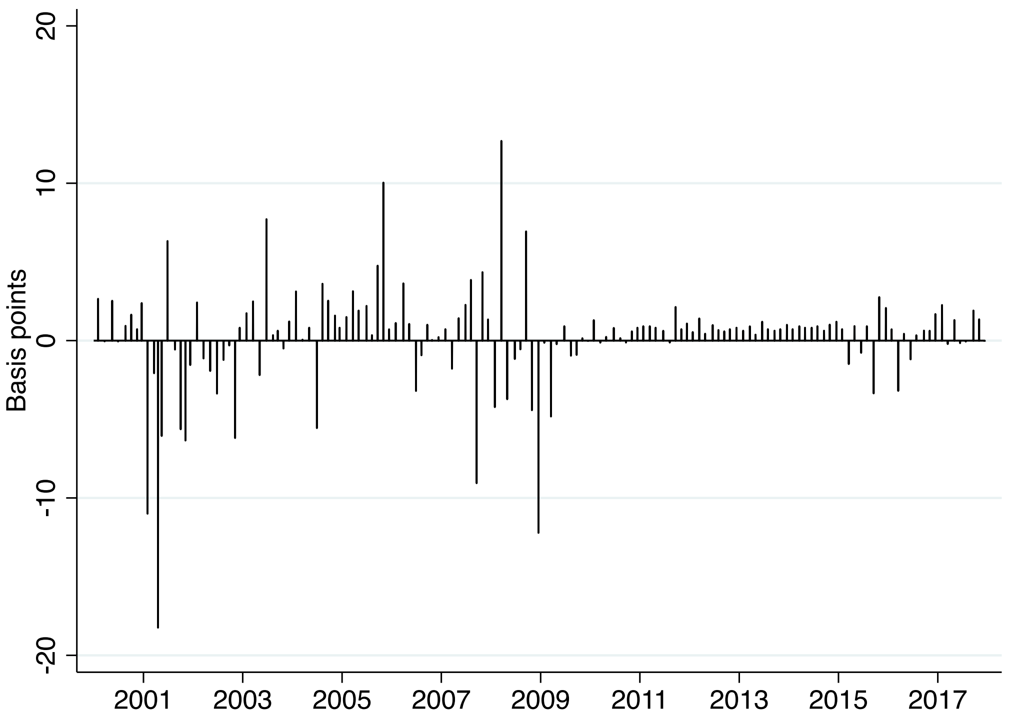
In order to measure monetary shocks, we use a high frequency identification technique (see, among others, Kuttner, 2001; Gürkaynak et al., 2005; Gertler and Karadi, 2015b; Nakamura and Steinsson, 2018). This method employs high frequency data on interest rate futures to identify the surprise component of monetary policy announcements. To derive this shock measure, changes in these futures are measured in a narrow time window around the FOMC meetings. If all publicly available information is already incorporated into the financial markets at the beginning of the time window, fluctuations in the interest rate futures around the FOMC announcement are only driven by the unexpected component of the monetary policy announcement. In order to ensure the exogeneity of the shock measure, it is crucial that the time span around the FOMC meeting is short enough, such that the only relevant shock during that time period (if any) is the monetary policy shock.

We adopt the method of Nakamura and Steinsson (2018) and construct monetary policy shocks as the first principle component of the daily change in five interest rate futures. These include federal funds futures (the current-month contract rate and the contract rates for each of the next three months) and Eurodollar futures (at two to four quarters in the future). We refer to the identified shocks as "high frequency monetary policy news shocks." For convenience, we scale the shocks such that their effect on the 1-year nominal Treasury is 100 basis points.[8] We use daily data from January 1, 2001, to December 31, 2017, and we include all FOMC scheduled meetings that happened throughout this period. Figure 3.1 depicts the time series of the policy news shock.[9]

[8] For a more detailed description of the method of Nakamura and Steinsson (2018), see Appendix 3.A.

[9] For a visual comparison of the high frequency monetary policy news shocks and the original daily Nakamura and Steinsson (2018)'s monetary shocks, see Table 3.6 in the appendix. Between 2001 and 2014 (the Nakamura and Steinsson (2018)'s monetary shock series is available only until 2014) The correlation between the two shock series is 0.88.

Figure 3.1: Monetary policy news shocks



Notes: The graph shows the monetary policy news shocks for the period 2001−2017 estimated at the daily frequency. The monetary policy shocks is scaled to have a 100 basis point impact on the 1-year US Treasury yield.

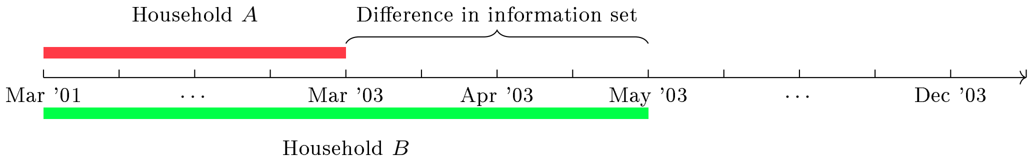
3.3.2 Biennial household-specific monetary policy shocks

After identifying daily monetary policy shocks, we need to aggregate them into biennial frequency to match the frequency of the household survey data. The simplest option is to aggregate the monetary policy shocks over 24 months (from January of wave $t-1$ to December of wave t). However, by doing so, we would neglect the fact that households are not interviewed in the same month and, thus, that their answers regarding the financial variables refer to different periods. Indeed, as shown in Figure 3.5 in the appendix, the interview dates are dispersed throughout all months of an interview year (with the exclusion of January and February).

Why should the difference in the interview dates matter for our analysis? Consider two households (A and B) that were interviewed in both 2001 and 2003. Suppose that in 2001 they are both interviewed in March, while in 2003, household A is interviewed in March but household B in May. Figure 3.2 provides a graphical presentation of the monetary policy shock information set that the two households experienced between the two surveys. We can clearly see that household A experienced fewer FOMC meetings than household B and, thus, is possibly exposed to fewer monetary policy shocks. Therefore, if we would aggregate the monetary policy news shocks from January 2011 to

December 2013, subsequently evaluating its effects on the investment behavior of both households, then we would obtain biased results. This can be especially problematic if, referring again to the example in Figure 3.2, there is a major monetary policy shock between March 2003 and May 2003.

Figure 3.2: PSID interview dates and FOMC meetings



Notes: The figure helps visualize how the PSID feature of staggered interviews in different months of the year implies that households A and B might be subject to different monetary policy shocks between waves $t-1$ (ending in March 2001 for both A and B) and t (ending in March 2003 for A and in May 2003 for B) if an FOMC meeting happens between March and May 2003.

Therefore, we construct a household-specific monetary policy news shock series that takes into account households' different information sets by summing up the monetary policy shocks for each household and taking into account their interview dates,

$$MP_{i,t} = \sum_{j=T_{i,t-1}}^{T_{i,t}} mps_j, \tag{3.1}$$

where $MP_{i,t}$ is the biennial monetary policy news shock series for household i in wave t; $T_{i,t-1}$ and $T_{i,t}$ are the day of household i's interview in wave $t-1$ and t, respectively; mps_j is the daily monetary policy news shock on day j. Figure 3.3 presents a boxplot of the household specific biennial shocks by wave. The green dots are the household specific biennial shocks data points, while the ends of the blue and red boxes are the lower and upper quartiles, respectively. The figure highlights the high level of data dispersion within each wave, confirming the importance of taking into account households' idiosyncratic exposure to monetary shocks when constructing the biennial monetary policy shock measure.[10]

3.3.3 Households' heterogeneous exposure to monetary policy

As a final step, we exploit the fact that households' responses to monetary policy can differ, depending on their socio-economic characteristics (see, among others, Wong, 2019; Forti Grazzini, 2020; Cumming and Hubert, 2020). This allows us to better understand the transmission mechanisms of monetary policy. We follow Forti Grazzini (2020) and consider financial wealth as the source of household heterogeneity. The intuition behind this choice is that the more financial wealth a household holds, the more it

[10] Table 3.8 in the appendix compares the summary statistics of the daily monetary policy news shock with the average biennial household-specific series over the empirical analysis period (2001-2017).

Figure 3.3: Biennial household-specific monetary policy news shock by wave

Notes: The graphs presents the boxplot of the biennial household-specific monetary policy news shock by wave. The green dots are the household-specific biennial shocks data points. They are obtained by summing up the monetary policy shock series at the daily frequency (Figure 3.1) within a two-year window that depends on the household's interview month to the PSID survey in each wave. The ends of the blue and red boxes are the lower and upper quartiles, respectively.

is affected by monetary shocks due to *valuation effects*: monetary policy impacts yields and prices of assets, thus affecting the value of households' stock of financial wealth. Consequently, we interact our biennial household-specific aggregated monetary policy shocks, $MP_{i,t}$, with the household's lagged financial wealth, $W_{i,t-1}$:

$$MP^*_{i,t} = MP_{i,t} \times W_{i,t-1}. \tag{3.2}$$

$W_{i,t-1}$ can be, depending on the empirical exercise we perform, either the lagged liquid assets or the lagged stock holding.

Finally, we also interact our monetary policy shock variable with a dummy variable that visualizes different households groups. If not stated otherwise, throughout the chapter the dummy variable $Head_i$ takes value one if the household's head is female and single, and zero if the head is male (baseline results). In this way, we are able to capture the possible gender/marital status-specific effects of monetary policy.

$$MP^*_{i,t} \times Head_i = MP_{i,t_i} \times W_{i,t-1} \times Head_i. \tag{3.3}$$

Furthermore, we also interact the monetary policy shock variable $MP^*_{i,t}$ with other dummy variables that distinguishes between (i) single female-headed households and married male-headed households; and (ii) married male-headed households and single male-headed households. We do this in order to confirm the robustness of our results.

3.4 Results

In this section, we present our econometric framework and results. First, we examine the effect of monetary policy on the change in stock market participation (entry and exit). Thereafter, we focus exclusively on stock market participants and analyze the effect of monetary policy on their trading activity. Throughout this section; we evaluate the impact of a contractionary monetary policy that is scaled to increase the 1-year Treasury yield by 100 basis points.

3.4.1 Monetary policy and stock market participation

We start by investigating how changes in the monetary policy stance affects the stock market participation status of households. We follow Brunnermeier and Nagel (2008) and employ the following probit model

$$\begin{aligned} y^*_{i,t} &= \delta_t + \delta_r + \alpha X_{i,t-1} + \beta_1 MP_{i,t} + \beta_2 MP_{i,t} \times Head_i + \beta_3 MP^*_{i,t} + \\ &\quad \beta_4 (MP^*_{i,t} \times Head_i) + \beta_5 (W_{i,t-1} \times Head_i) + \beta_6 W_{i,t-1} + \beta_7 Head_i + u_{i,t}, \\ y_{i,t} &= 1\left[y^*_{i,t} > 0\right] \end{aligned} \tag{3.4}$$

where $y_{i,t}$ can be either $Exit_{i,t}$ or $Entry_{i,t}$; $X_{i,t}$ is a vector of household-level controls that includes financial characteristics (lagged net worth and family income, change in net worth and family income, total inheritance, dummy for the first mortgage, dummy for the second mortgage) and demographic characteristics (the number of children, the age of the head, the head's age squared, marital status, completed college education, working in the finance industry, total number of family components, home ownership). We also include time- and region of residency- fixed effects (δ_t and δ_r). $Head_i$ is a dummy variable that, depending on the exercise we perform, allows us to compare different household groups. $W_{i,t-1}$ is the lagged liquid assets. The remaining terms in equation (3.4) are the triple interaction term constructed in section 3.3.3 and all other mean and double interaction effects that should be included when employing a three-way interaction term. Thus, with our empirical model, we are able to capture the mean effect of monetary policy (β_1) and how this effect varies across different household groups (β_2), across different values of the exposure variable (β_3), and across different household groups along different values of the exposure variable (β_4). $u_{i,t}$ is the error term. We estimate the model with maximum-likelihood on the 2001-2017 sample. Standard errors are clustered at the household level.

Table 3.2 presents the results. We present the marginal effects at the mean, i.e. evaluated at the sample mean of the explanatory variables. We only show the marginal effects of the parameters of interest. Columns (1) and (2) present the baseline results, where we compare single women with male-headed households (both married or single). Results in column (1) show no differences between SFHHs and MHHs stock market exit. On the contrary, column (2) suggests that there is a negative effect for single female headed households on the stock market entry decision, as they are 3.4% less likely to enter the stock market. This last finding is in line with the literature that documents a high female non-participation rate in the financial markets (Sunden and Surette, 1998; Barber and Odean, 2001; Dwyer et al., 2002; Agnew et al., 2003).

Now we turn to the effects of monetary policy. While a contractionary shock does not have a significant effect on households' exit decision, column (2) shows that it does have a negative and highly significant effect on SFHHs probability of stock market entry, as their likelihood of entering the stock market decreases by 11%. On the contrary, the entry probability of MHH is not significant. To ensure that the baseline results are not driven by either of the two male-headed subgroups, we repeat the analysis comparing single female-headed households with only married male-headed households. Marginal effects are reported in columns 3 and 4 of Table 3.2.[11] Results are very similar to the baseline, both in sign and magnitude: monetary policy does not have a significant impact on households' stock market exit decision and it only affects the entry choice of single women.

Taken together, these findings suggest that single female-headed households are the only group being significantly affected by monetary policy in their stock market participation decision. Moreover, we find that the monetary policy-driven difference between single female- and male-headed households in the entry decision is sizeable even when holding constant characteristics that are correlated with gender, marital status, and financial wealth, as position in the life cycle, education, and income. To the contrary, monetary policy does not affect different household groups' probability to exit the stock

[11] Here $Head_i$ equals one if the household head is single and female, it equals zero if it is married and male.

market, suggesting that female- and male-headed households do not behave differently once they participate in financial markets.[12]

Table 3.2: Monetary policy and stock market participation decision - marginal effects

	Single female-headed HH VS Male-headed HH		Single female-headed HH VS Married male-headed HH	
	Exit (1)	Entry (2)	Exit (3)	Entry (4)
Single female-headed HH	-0.017	-0.034***	-0.027	-0.045***
	(0.069)	(0.0076)	(0.070)	(0.008)
MP if single female-headed HH	0.211	-0.110**	0.248	-0.105**
	(0.273)	(0.053)	(0.275)	(0.050)
MP if male-headed HH	-0.068	-0.097		
	(0.164)	(0.065)		
MP if married male-headed HH			-0.047	-0.107
			(0.169)	(0.069)
Constant	yes	yes	yes	yes
Other inter. terms	yes	yes	yes	yes
Financial var., lag	yes	yes	yes	yes
Demographics, lag	yes	yes	yes	yes
Household FE	no	no	no	no
Time FE	yes	yes	yes	yes
Observations	3649	11,129	3,437	10,339
Pseudo R^2	0.07	0.07	0.08	0.07

Notes: This table presents the marginal effects of a contractionary monetary policy shock that increases the US 1-year Treasury by 100 basis points on stock market entry and exit decisions. We compare different households subsamples: single female-headed households and both married and single male-headed households in columns 1 and 2; single female-headed households and married male-headed households in columns 3 and 4. All models include data from 2001 to 2017. The variable $Exit_{i,t}$ is a dummy equal to 1 if the household exits the stock market in t and 0 if it stays in; the variable $Entry_{i,t}$ is a dummy equal to 1 if the household entries the stock market in t and zero if it does not. All marginal effects are obtained after estimating equation (3.4) and are evaluated at the sample average of the explanatory variables. Only the coefficients of interest are reported here. Standard errors (in parentheses) are clustered at the household level. *, **, and *** represent statistical significance at the 10%, 5%, and 1% levels, respectively.

[12]Robustness exercises are provided in Appendix 3.D.

3.4.2 Monetary policy and stock active saving

We now focus solely on stock market participants and examine how different groups of households adjust their stock investments following a monetary policy shock. Following Juster et al. (2006) and Calvet et al. (2009), we employ a fixed effect model

$$\begin{aligned}\mathrm{AS}_{i,t} =&\delta_i + \delta_t + \alpha X_{i,t-1} + \beta_1 MP_{i,t} + \beta_2 MP_{i,t} \times Head_i + \beta_3 MP^*_{i,t} + \\ &\beta_4(MP^*_{i,t} \times Head_i) + \beta_5(W_{i,t-1} \times Head_i) + \beta_6 W_{i,t-1} + \beta_7 Head_i + \varepsilon_{i,t},\end{aligned} \tag{3.5}$$

where $\mathrm{AS}_{i,t}$ is the net purchase amount of stocks of household i between $t-1$ and t; δ_i and δ_t are the individual- and time fixed effects, respectively; $X_{i,t-1}$ includes the same financial and demographic characteristics as the probit model described in section 3.4. Again, all remaining terms capture the three-way interaction effect and we are interested in the coefficients β_1-β_4. Since we consider households that participate in the stock market both in $t-1$ and t, we use the previous wave stock investment as the exposure variable for monetary policy ($W_{i,t-1}$). Standard errors are clustered at the household level.

Results are reported in Table 3.3. In our baseline specification, we compare single female-headed households with all male-headed households (column 1). Afterwards, we analyze households with single female and married male heads (column 2). Finally, in column 3, we contrast single male- and married male-headed households. Let us first concentrate on the baseline results in column 1. Monetary policy seems to play an important and significant role for the investment behavior of *all* households: after a contractionary monetary policy shock, households sell stocks (-0.003, MP^* coefficient). In economic terms, this means that investors sell, on average, \$762 of their stock investment after a contractionary monetary policy shock that increases 1-year Treasury yields by 100 basis points.[13]

However, the gender/marital status of the household head does not play a role, as the coefficients attached to any of the terms that include the dummy $Head_i$ are not significant. This result indicates that once single women participate in the stock market, their active saving decisions are not systematically different from those of households with a male head. This homogeneous response to monetary policy seems controversial to the literature that documents behavioral differences across gender in the financial markets. Nonetheless, it is important to highlight that our results do not reject the fact that different household groups invest heterogeneously, rather it provide evidence that, when participating in financial markets, single female- and male-headed households *react* to monetary policy in a homogeneous manner. Therefore, both household groups seem to understand the inverse relationship between the interest rates and asset prices.[14]

[13] The figures are calculated by multiplying the coefficients with the average stock holdings (\$253891.93).

[14] We perform two robustness checks. First, we use the value of households' financial portfolios of the previous wave as an alternative exposure variable. Second, we repeat the analysis including households that participate in the survey for two consecutive waves (in contrast to three). The results are contained in columns 1 and 2 of Table 3.9 in the appendix and they are very similar to the baseline results.

Table 3.3: Monetary policy and stock active saving

	Single female-headed HH VS Male-headed HH (1)	Single female-headed HH VS Married male-headed HH (2)
MP	-215.5	-136.1
	(7550.4)	(7859.8)
$MP \times Head$	8776.6	9222.2
	(7041.6)	(7565.8)
MP^*	-0.003**	-0.003**
	(0.001)	(0.001)
$MP^* \times Head$	0.046	0.053
	(0.036)	(0.042)
Constant	yes	yes
Other inter. terms	yes	yes
Financial var., lag	yes	yes
Demographics, lag	yes	yes
Household FE	yes	yes
Time FE	yes	yes
Observations	2,389	2,236
R^2	0.01	0.01

Notes: This table presents the results of the fixed effect model in equation (3.5) estimated using the sub-sample of households participating in the stock market between 2001 and 2017. The dependent variable is stock active saving. In column (1), the dummy $Head_i$ is equal to 1 if the household head is single and female, it is equal to 0 if the household head is male; in column (2) the dummy $Head_i$ is equal to 1 if the household head is single and female, it is equal to 0 if the household head is married and male. The variable $MP_{i,t}$ is the household-level biennial monetary policy shock series constructed in Section 3.3.2. The variable $MP^*_{i,t}$ is the interaction between $MP_{i,t}$ and the household's lagged stock holding, $W_{i,t-1}$. Standard errors (in parentheses) are clustered at the household level. *, **, and *** represent statistical significance at the 10%, 5%, and 1% levels, respectively.

3.4.2.1 Monetary policy and active saving at different levels of financial wealth

One might suspect that the absence of a gender-specific and marital status-specific response to monetary policy is driven by investment decisions of wealthier households. In fact, differences in risk aversion, inertia, and financial literacy across gender progressively decline for increasing values of financial investment. This, in turn, would increase the chance that single female- and male-headed household heads at the top of the financial wealth distribution react homogeneously to monetary shocks, therefore influencing the direction and magnitude of our estimates.

Table 3.4: Monetary policy and stock active saving - households at the top and bottom of their respective group's liquid asset distribution

	Single female-headed HH VS Male-headed HH	
	Top 50% of the liquid asset distribution (1)	Bottom 50% of the liquid asset distribution (2)
MP	7627.0	-8168.6
	(14297.5)	(7237.6)
$MP \times Head^{top}$	25251.4	
	(19787.8)	
$MP \times Head^{bottom}$		759.2
		(5662.1)
MP^*	-0.004**	-0.050
	(0.002)	(0.087)
$MP^* \times Head^{top}$	0.010	
	(0.077)	
$MP^* \times Head^{bottom}$		0.298
		(0.298)
Constant	yes	yes
Other inter. terms	yes	yes
Financial var., lag	yes	yes
Demographics, lag	yes	yes
Household FE	yes	yes
Time FE	yes	yes
Observations	1,194	1,195
R^2	0.01	0.01

Notes: In column (1) the dummy $Head^{top,i}$ is equal to 1 if the household head is single, female and in the top 50% of its household group liquid asset distribution; it is equal to zero 0 if the household head is male and in the top 50% of its household group liquid asset distribution. In column (2) the dummy $Head^{bottom,i}$ is equal to 1 if the household head is single, female, and in the bottom 50% of its household group liquid asset distribution; it is equal to zero 0 if the household head is male and in the bottom 50% of its household group liquid asset distribution. Standard errors (in parentheses) are clustered at the household level. *, **, and *** represent statistical significance at the 10%, 5%, and 1% levels, respectively.

To examine this hypothesis, we repeat our baseline analysis splitting our sample in two. Results are reported in Table 3.4. In column (1) we compare single female- and male-headed households from the top 50% of the liquid wealth distribution; in column (2) we analyze households of the bottom 50%. The dependent variable is stock active saving. For the top 50%, the coefficients are similar to our baseline estimates presented in Table 3.3 (column 1), although slightly bigger in absolute value. This confirms that wealthier households respond more heavily to monetary policy and that they display no heterogeneity in their response. Compared to this, the picture of the bottom 50% is quite different, as there is no evidence of a systematic response to monetary policy (column 2).

3.4.2.2 Monetary policy and riskless active saving

In this section, we additionally test whether there are any significant differences among different groups of households in the way they rebalance their riskless investment. According to both the rebalancing channel (Gagnon et al., 2010; Joyce et al., 2012) and the risk-taking channel of monetary policy (Lian et al., 2018; Daniel et al., 2018), contractionary monetary policy induces investors to rebalance their portfolio by selling risky assets (like stocks) to purchase safer options (like Treasury bonds). To do so, we re-estimate equation (3.5) using as the dependent variable the net purchase amount of riskless assets of household i between $t-1$ and t. Results are shown in Table 3.5. Additionally, in this case, we report the results of two comparisons, SFHHs vs MHHs (column 1) and SFHHs vs SMHHs (column 2).

On the one hand, column (1) shows that after a contractionary monetary shock, on average, households buy more riskless assets (+0.088, MP^* coefficient). This result is consistent with the rebalancing channel of monetary policy. On the other, hand we also find a negative and slightly significant coefficient attached to the interaction between monetary policy and the household head's gender/marital status, suggesting that the overall effect for female-headed households is negative. Nevertheless, the coefficient is weakly significant at the 10% level and robustness exercises show that this result does not hold for most of the other specifications (see Table 3.10 and 3.11 in the appendix). Further, there is also no significant difference when contrasting single female- with only married male-headed households (see column (2) of Table 3.5).

Table 3.5: Monetary policy and riskless asset active saving

	Single female-headed HH VS Male-headed HH (1)	Single female-headed HH VS Married male-headed HH (2)
MP	-1034.0	-14364.7
	(46378.3)	(48697.7)
$MP \times Head$	1949.8	-6510.8
	(41660.9)	(42237.1)
MP^{*}	0.088***	0.089***
	(0.010)	(0.010)
$MP^{*} \times Head$	-0.700*	-0.586
	(0.423)	(0.466)
Constant	yes	yes
Other inter. terms	yes	yes
Financial var., lag	yes	yes
Demographics, lag	yes	yes
Household FE	yes	yes
Time FE	yes	yes
Observations	2,389	2,236
R^2	0.10	0.11

Notes: In column (1), the dummy $Head_i$ is equal to 1 if the household head is single and female, it is equal to 0 if the household head is male; in column (2) the dummy $Head_i$ is equal to 1 if the household head is single and female, it is equal to 0 if the household head is married and male. The variable $MP_{i,t}$ is the household-level biennial monetary policy shock series constructed in Section 3.3.2. Standard errors (in parentheses) are clustered at the household level. *, **, and *** represent statistical significance at the 10%, 5%, and 1% levels, respectively.

3.5 A counterfactual analysis

Our empirical analysis shows that monetary policy has heterogeneous effects on different household groups with respect to gender and marital status, but only when it comes to their entry decisions. After a contractionary monetary policy shock that increases the 1-year US Treasury bond yields by 100 basis points, single women are 11% less likely to enter the stock market. However, how large is this effect in terms of economic numbers? In this section we conduct a static simulation exercise to visualize how the entry rate of single women is affected by monetary policy shocks during our sample period of 2001 to 2017. We then calculate how much financial wealth they potentially missed out or gained through monetary policy induced non-participation or entry in the stock market.

To construct the counterfactual of how many single women would have participated in the stock market, had monetary policy not affect their entry decisions, we proceed in the following way. First, we aggregate the daily monetary policy shocks into biennial frequency and quantify the average effect of monetary policy on single women's entry decisions. For this exercise, we do not use the household-specific aggregated shocks, but sum the shocks from January 1st of wave $t-1$ to December 31 of wave t (Table 3.6, column 1). Next, we use the stock market entry rate per wave of single female-headed households provided by the PSID (Table 3.6, column 2) to calculate how high the entry rate would be if their entry decisions were unaffected by monetary policy (same table, column 3). For example, the biennial monetary policy shock between 2001 and 2003 increases the 1-year US Treasury yield by 11.5 basis points. From Table 3.2, we know that this decreases the likelihood of single female-headed household entering the stock market by $0.11 \times 0.115 = 0.013$.[15] Thus, in 2003 single female-headed households are 1.3% less likely to enter the stock market. Finally, we can use this information together with the actual entry rate provided by the PSID data (that is 10.4% in 2003) to construct their entry rate without a monetary policy shock, $0.104 \times (1 + 0.013) = 0.105$.

Table 3.6: Entry rate with and without monetary policy shocks

	Biennial MP shock (1)	Entry rate with MP (2)	Entry rate without MP (3)	Δ entry rate (4)	S&P 500 biennial return (5)
2003	0.115	0.104	0.105	-0.001	-0.031
2005	0.317	0.082	0.085	-0.003	0.123
2007	0.061	0.078	0.079	-0.001	0.176
2009	-0.127	0.065	0.064	0.001	-0.241
2011	0.107	0.039	0.039	-0.001	0.128
2013	0.124	0.057	0.058	-0.001	0.470
2015	0.088	0.032	0.032	-0.000	0.106
2017	0.064	0.057	0.057	-0.000	0.308

Note: This table provides information on the biennial cumulated monetary policy shocks (column 1, see Section 3.5 for construction); the single female-headed household entry rate in the stock market taking into account the effect of monetary policy (column 2, provided by the PSID); the single female-headed household entry rate in the stock market excluding the effect of monetary policy (column 3, authors' own calculations); the difference between entry rate with and without monetary policy (column 4); the S&P 500 index biennial return (column 5, provided by Bloomberg).

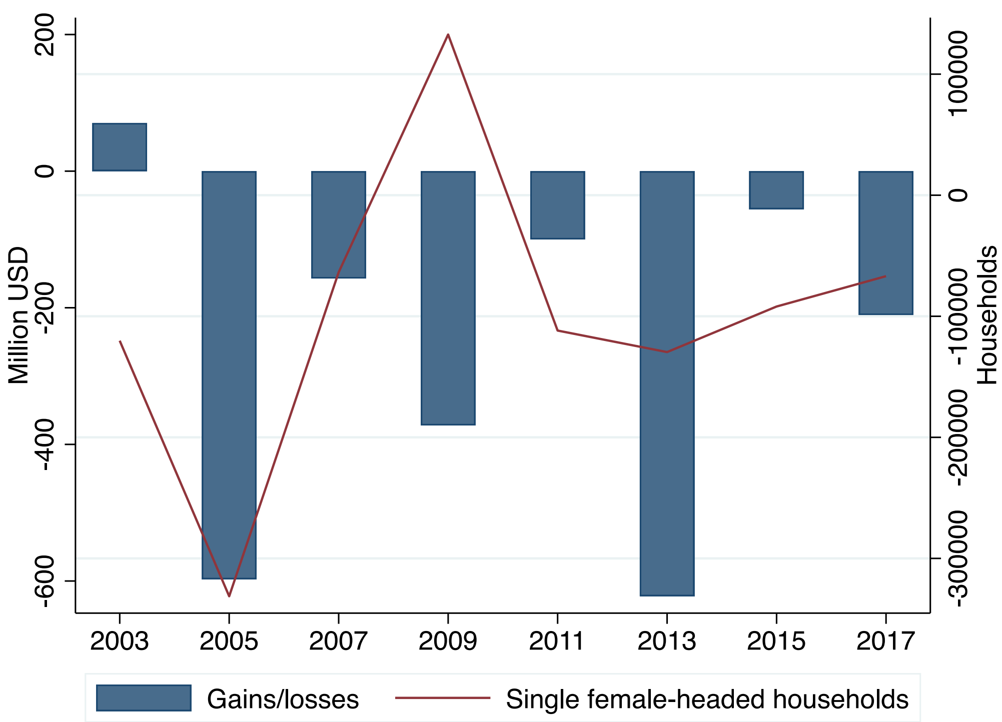
At first glance, Table 3.6 shows that monetary policy is mostly of a contractionary nature in our sample, implying that the entry rate of women has been rather negatively affected by monetary policy (with only the biennial aggregated shocks between 2007 and 2009 being negative, hence accommodative). Second, if we compare the entry rate

[15] The coefficient 0.11 in Table 3.2 is the marginal response to a monetary policy shock the increases the 1-year Treasury yields by 100 basis points. Thus, we need to scale it down for the actual magnitude of the shock, 11.5 basis point.

with and without monetary policy, we would be inclined to think that the difference is very marginal (column 4). However, how many single female-headed households are affected by monetary policy? How much capital gains do they gain (or miss) by entering or staying out of the stock market? According to *Statista* - a data portal of households - in 2018 there are about 15 million single female-headed households. If we assume that this figure is constant throughout our sample, we can calculate how many single women entered or did not enter the stock market due to monetary shocks (red line in Figure 3.4).

We calculate the capital gains (or losses) that single female-headed households experience due to monetary policy induced entry/non entry in the stock market. For this exercise, we use single women's average stock holding, which is $113215.46 (Table 3.1) and we assume that this number is constant throughout all waves. We proxy the average biennial stock market investment return using S&P 500 index return (Table 3.6, column 5). Thus, we multiply the biennial return (e.g., -3.1% in 2003) for the average stock investment. Finally, we multiply the obtained biennial capital gains with the number of single female-headed households that are affected by monetary policy (column 4). The blue bars in Figure 3.4 present the final numbers.

Figure 3.4: Capital gains (losses) of single female-headed households



Notes: The redline presents the number of single female-headed households that entered (or did not enter) the stock market due to a monetary policy shock. The blue bars show the corresponding capital gains/losses over the two waves.

Between 2003 and 2017, single female-headed households missed \$256 million of capital gain every two years, for a total of \$2046 million over the entire period. These financial losses are equally distributed before and after the beginning of the zero lower bound period (2009), with stronger monetary policy shock but lower (and also negative) stock market returns in the first half of the sample and smaller monetary policy shocks but higher returns in the post ZLB.[16] Therefore, our results show that, although the effect of monetary policy on stock market entry may seem small, the missed out capital gains can be significantly large. Due to the fact that single female-headed families are the poorest household group in the USA, live the longest, and often have fewer working years than other household groups, this can have severe effects on their financial wealth. This, in turn can negatively affect their already higher probability to be living in poverty during their retirement years.

3.6 Conclusion

This chapter investigates potential heterogeneous effects of monetary policy on stock market investment choices of single women compared to both single and married male-headed households. The empirical analysis shows that, on the one hand, contractionary monetary policy negatively affects single female-headed households' stock market participation status, decreasing their probability of stock market entry (while this is not the case for male headed-households). On the other hand, monetary policy does not have a heterogeneous impact across household groups with regards to their decision to exit the stock market or to rebalance their equity investment. Finally, we conduct a simulation study to quantify the missed out capital gains stemming from monetary policy-driven stock market non-participation.

We conclude that gender and marital status significantly affect household response to central banks' actions. Single female-headed households are more sensitive to monetary policy cycles than male headed-households, but only if they are not already participating in the stock market. This is true even controlling for a wide range of demographic and financial characteristics that could explain such differences. Therefore, we conclude by highlighting the importance of making single women resilient to monetary policy adjustments by educating them on saving and investment decisions. In the US, investment decisions are particularly important for the accumulation of wealth for retirement, something that is especially true for single female-headed households, as they are the most fragile family groups in society.

[16] We also perform additional and complementary counterfactual exercises. The results are presented in Figure 3.7 in the appendix.

3.A Monetary policy shock identification

The identification method of Nakamura and Steinsson (2018) employs high frequency data on interest rate futures to construct a monetary policy shock measure. It identifies the exogenous and unanticipated component of FED's announcements (the "shock") by extracting it directly from financial market responses. The identification strategy relies on measuring the change in the futures during a narrow time window around FOMC meetings. The idea is that right before any meeting, all public available information of the economy is already incorporated into the financial markets and reflected in their prices/yields. Thus, if the time span around the FOMC announcement is tight enough, any immediate change in the futures is dominated by the information about future monetary policy contained in the announcement itself. Moreover, by using a broad range of interest rate futures, the measure captures not only unanticipated changes in the federal funds rate, but also the effect of "forward guidance" and other unconventional monetary policies.

Nakamura and Steinsson (2018) construct the monetary policy news shock as the first principal component of the change in five interest rate futures. The first of these is the change in market expectations of federal funds rates during a narrow time window around FOMC meetings. In general, the payout of the federal funds futures is calculated as the average effective federal funds rate that prevails over the calendar month specified in the contract. Therefore, immediately before an FOMC meeting at time $t-\Delta t$, the current-month federal funds future contract can be written as the weighted average of (i) the federal funds rate of the month r_0 (before the FOMC meeting) and (ii) the rate that is expected to prevail for the remainder of the month r_1,

$$ff1_{t-\Delta t} = \frac{d1}{D1}r_0 + \frac{D1-d1}{D1}E_{t-\Delta t}(r_1), \tag{3.6}$$

where $d1$ denotes the day of the FOMC meeting, $D1$ is the number of days in the month.[17] Accordingly, the current-month federal funds rate contract right after the FOMC meeting is,

$$ff1_t = \frac{d1}{D1}r_0 + \frac{D1-d1}{D1}E_t(r_1). \tag{3.7}$$

Thus, the change in expectations before and after the FOMC meeting can be calculated as

$$mp1_t \equiv E_t(r_1) - E_{t-\Delta t}(r_1) = (ff1_t - ff1_{t-\Delta t})\frac{D1}{D1-d1}. \tag{3.8}$$

The second future we use is the change in expectations regarding the federal funds rate target after the *second* FOMC meeting. We apply the same method:

[17]Note that Gürkaynak et al. (2005) introduce a risk premium term in the equation. For simplicity, we exclude this term.

$$ff2_{t-\Delta t} = \frac{d2}{D2}E_{t-\Delta t}(r1) + \frac{D2-d2}{D2}E_{t-\Delta t}(r2), \tag{3.9}$$

where $d2$ and $D2$ are the day of that FOMC meeting and the number of days in the month containing that FOMC meeting, respectively.[18] Again, using the same calculations as above, we are able to calculate the change in expectations at the time of the next scheduled FOMC meeting,

$$mp2_t \equiv E_t(r_2) - E_{t-\Delta t}(r_2) = \left[(ff2_t - ff2_{t-\Delta t} - \frac{d2}{D2}mp1_t)\right]\frac{D2}{D2-d2}. \tag{3.10}$$

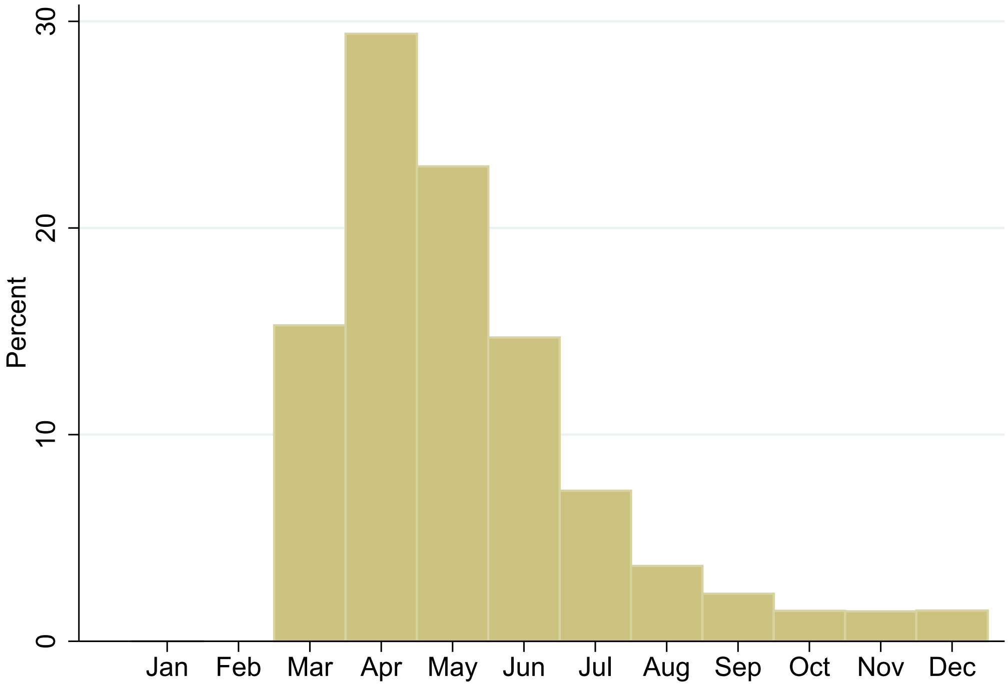
The last set of interest rate futures we use are the change in the price of three Eurodollar futures at the time of the FOMC meetings. Following Nakamura and Steinsson (2018), we use the Eurodollar futures at horizons of two, three, and four quarters in the future.

Since only daily data is available to us, we are not able to construct the monetary policy news shock within a 30-minute window. This can affect the exogeneity of our measure due to a wider time span around the FOMC meetings. Nevertheless, Piazzesi and Swanson (2008) show that a daily window is sufficient to identify exogenous components of monetary policy announcements. As a robustness check, we download the publicly available monetary policy shock series of Nakamura and Steinsson (2018), which stops in 2014 (our sample is until 2017) and apply them to our empirical analysis. The results remain robust.

[18]The next scheduled FOMC meeting can occur between the next month up to three months after the current meeting.

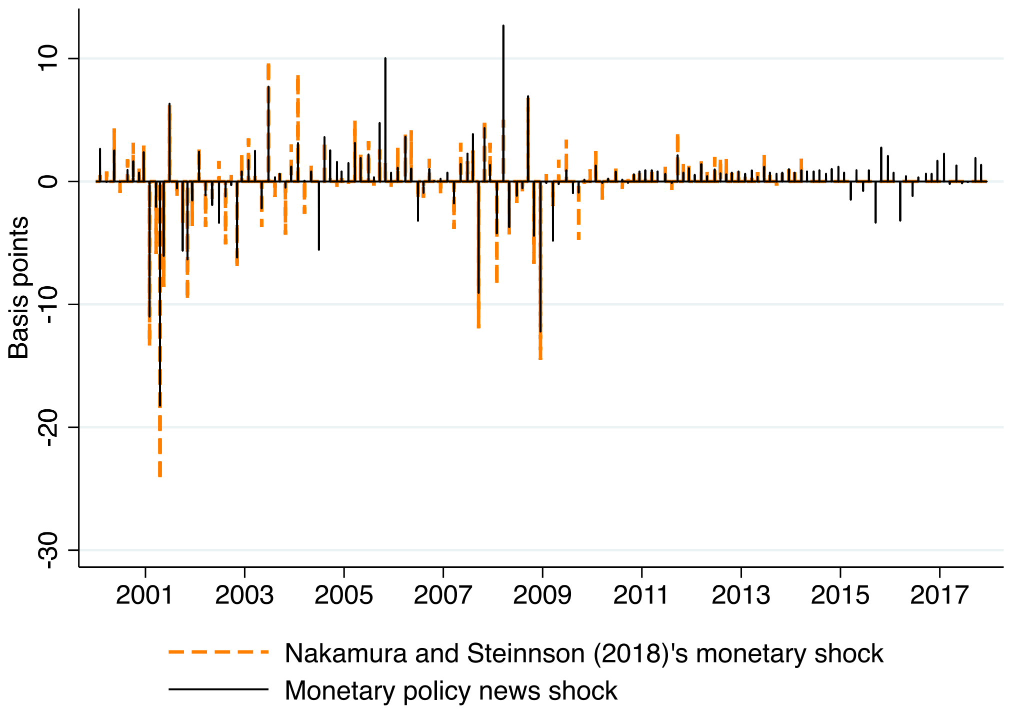
3.B Figures

Figure 3.5: Distribution of PSID interview months, 2003-2017



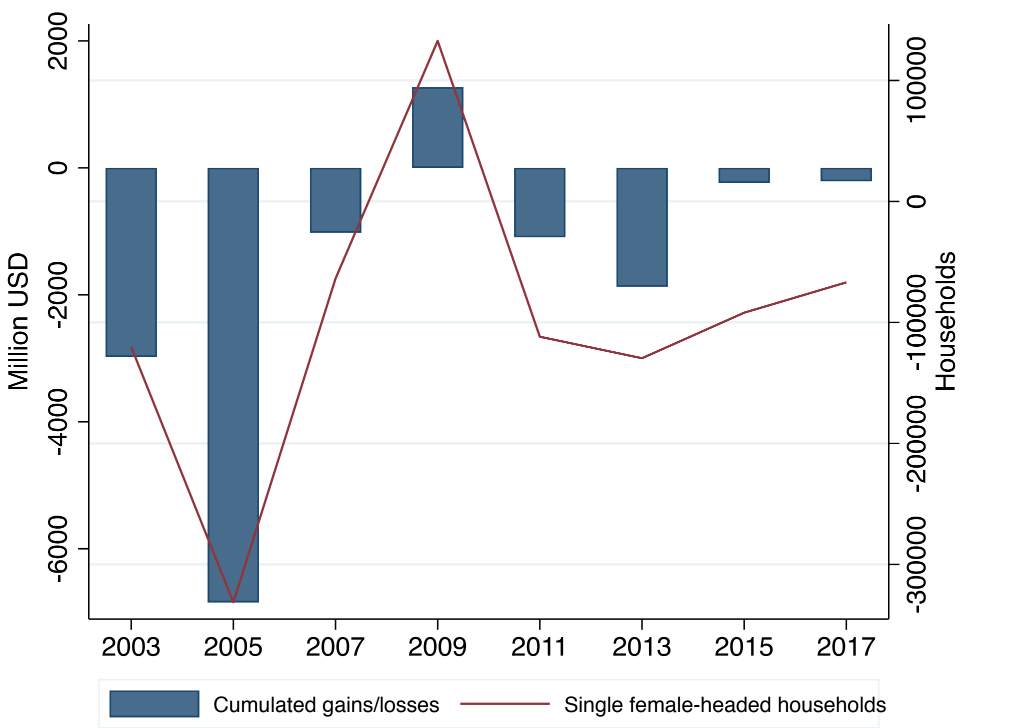
Notes: This figure shows the cross sectional distribution of PSID interviews over the year 2003-2017. *Source*: PSID

Figure 3.6: Nakamura and Steinsson (2018)'s original shocks and our monetary policy news shocks



Notes: The original Nakamura and Steinsson (2018)'s monetary shock series (dashed orange line) is constructed as the first principal component of the daily change in five interest rate futures around FED's FOMC meetings and is estimated using intra-daily data and is available until end-2014. The monetary policy news shocks (solid black line) is constructed as the first principal component of the daily change in five interest rate futures around FED's FOMC meetings using daily data. In both cases the included futures are the federal funds futures (the current-month contract rate and the contract rates for each of the next three months) and Eurodollar futures (at two to four quarters in the future). Both shocks are scaled to have a 100 basis point impact on the 1-year US Treasury yield.

Figure 3.7: Cumulated capital gains/losses of single women due to monetary policy



Notes: The graphs depicts: the number of single female-headed households that enter or do not enter the stock market due to monetary shocks (red line); the single female-headed households' capital gains/losses due to the monetary policy-driven stock market participation or non participation decision cumulated from the first wave they appear in the PSID (on the x-axis) to the end of the sample, 2017 (blue bars).

3.C Tables

Table 3.7: Summary statistics - unweighted sample

	Mean	SD	Mean	SD	Mean	SD
	All households		Single female-headed HHs		Male-headed HHs	
Panel A: Full sample						
Stock market participation	0.23	0.42	0.12	0.33	0.25	0.43
Stock market exit	0.33	0.47	0.39	0.49	0.32	0.47
Stock market entry	0.09	0.29	0.06	0.23	0.10	0.30
Stock holding	41225.61	545474.93	9226.45	75603.48	46975.97	591304.09
Riskless asset holding	22894.65	78762.07	10111.35	27983.65	25163.57	84503.71
Stock active saving	240986.52	12062381.55	205432.69	10481270.23	247823.64	12335552.57
Riskless asset active saving	4980.50	79403.64	1052.36	26494.25	5673.18	85480.07
Liquid asset holding	77781.49	613284.96	28763.40	162369.97	86579.55	662079.63
Stocks/liquid assets	0.13	0.28	0.07	0.22	0.14	0.29
Net worth	300973.88	1180274.72	140516.41	1030338.20	329636.15	1203852.51
Income	88062.41	111588.71	44404.58	34914.10	95799.12	118539.65
Home ownership	0.82	0.39	0.62	0.49	0.85	0.36
Observations	17496		2628		14837	
Panel B: Stock market participants						
Stocks	225329.34	1347519.14	94468.94	247749.63	235573.54	1396764.47
Riskless assets	53373.19	146875.34	31682.75	53956.23	55076.49	151616.12
Stock active saving	3421.39	31033.94	-207.80	5844.05	3705.44	32168.53
Riskless asset active saving	12196.13	148708.05	4923.89	51471.68	12765.47	153712.16
Liquid assets	308648.77	1485082.53	150920.92	368328.37	321018.26	1537715.64
Stocks/liquid assets	0.59	0.30	0.59	0.31	0.59	0.29
Net worth	808553.83	2160744.22	792217.26	3523859.95	810332.24	2018914.54
Income	139634.08	193253.61	63216.63	70290.99	145537.43	198467.16
Home ownership	0.93	0.26	0.86	0.35	0.93	0.25
Observations	2389		162		2227	

Notes: The table shows the summary statistics of the relevant wealth and income variables included in the analysis. Panel A presents figures for the full sample; Panel B for the sub-sample of stock market participants (at least \$1 invested in stock in both $t-1$ and t). The sample period is 2001-2017. *Source*: PSID and own calculations.

Table 3.8: Summary statistics of the monetary policy news shocks, 2001-2017

	High-frequency monetary policy news shocks	Biennial household-specific monetary policy news shocks
Mean	0.001	0.039
Median	0.007	0.080
Std. Dev.	0.034	0.177
Min	-0.182	-0.433
Max	0.127	0.306

Notes: Summary statistics of monetary shocks over the period 2001-2017. The shocks are all scaled to have a 100 basis point impact on the 1-year Treasury bond yields. The high frequency monetary policy news shocks are constructed as the first principal component of the daily change in five interest rate futures around FED's FOMC meetings. The included futures are the federal funds futures (the current-month contract rate and the contract rates for each of the next three months) and Eurodollar futures (at two to four quarters in the future). The biennial household-specific monetary news shocks are obtained by summing up the monetary policy shock series at the daily frequency (Figure 3.1) within a two-year window that depends on the household's interview month to the PSID survey in each wave.

Table 3.9: Different specifications for stock active savings

	Single female-headed HH VS Male-headed HH	
	Liquid assets (1)	Two waves (2)
MP	225.5	-1515.5
	(7495.0)	(6344.5)
$MP \times$ Head	4689.4	6775.6
	(6355.6)	(6027.1)
MP^*	-0.002*	-0.002*
	(0.001)	(0.001)
$MP^* \times$ Head	0.065	0.058
	(0.129)	(0.036)
Constant	yes	yes
Other inter. terms	yes	yes
Financial var., lag	yes	yes
Demographics, lag	yes	yes
Household FE	yes	yes
Time FE	yes	yes
Observations	2,389	2,660
R^2	0.01	0.01

Notes: This table presents the results of the fixed effect model in equation (3.5) estimated using the sub-sample of households participating in the stock market over the years 2001-2017. The dependent variable is the stock active saving. In Column (1) the monetary policy exposure variable is the household's lagged value of the liquid assets (instead of the lagged value of the stock investment). Column 2 includes households that participate in the PSID survey for at least two consecutive waves (insted of at least three consecutive waves). The dummy $Head_i$ is equal to 1 if the household head is single and female, it is equal to 0 if the household head is male. The variable $MP_{i,t}$ is the household-level biennial monetary policy shock series constructed in Section 3.3. The variable $MP^*_{i,t}$ is the interaction between $MP_{i,t}$ and $W_{i,t-1}$, the monetary policy exposure variable. Standard errors (in parentheses) are clustered at the household level. *, **, and *** represent statistical significance at the 10%, 5%, and 1% levels.

Table 3.10: Monetary policy and riskless asset active saving - alternative US Treasury maturities to calculate riskless active saving

	1-year US Treasury	2-year US Treasury	10-year US Treasury
	Riskless Act. Sav	Riskless Act. Sav.	Riskless Act. Sav
	(1)	(2)	(3)
MP	-1034.0	-745.6	-1076.6
	(46378.2)	(46389.6)	(46547.5)
MP × Head	1949.8	1384.3	-92.3
	(41660.9)	(41681.6)	(41549.6)
MP^{*}	0.088***	0.088***	0.089***
	(0.010)	(0.010)	(0.010)
MP^{*} × Head	-0.700*	-0.696	-0.686
	(0.423)	(0.424)	(0.426)
Constant	yes	yes	yes
Other inter. terms	yes	yes	yes
Financial var., lag	yes	yes	yes
Demographics, lag	yes	yes	yes
Household FE	yes	yes	yes
Time FE	yes	yes	yes
Observations	2,389	2,389	2,389
R^2	0.10	0.10	0.10

Notes: This table the presents results of the fixed effect model in equation (3.5) estimated using the sub-sample of households participating in the stock market over the years 2001-2017. In all columns the dependent variable is the riskless asset active saving, calculated using the US Treasury maturity that appears in the first row (for a detailed explanation of the methodology used to calculate the riskless asset active saving, please refer to Section 3.2.1.2). The dummy $Head_i$ is equal to 1 if the household head is single and female, it is equal to 0 if the household head is male. The variable $MP_{i,t}$ is the household-level biennial monetary policy shock series constructed in Section 3.3. The variable $MP^{*}_{i,t}$ is the interaction between $MP_{i,t}$ and the household's lagged value of the liquid assets, $W_{i,t-1}$. Standard errors (in parentheses) are clustered at the household level. *, **, and *** represent statistical significance at the 10%, 5%, and 1% levels.

Table 3.11: Monetary policy and riskless active saving - households that participate in the PSID survey for at least two consecutive waves

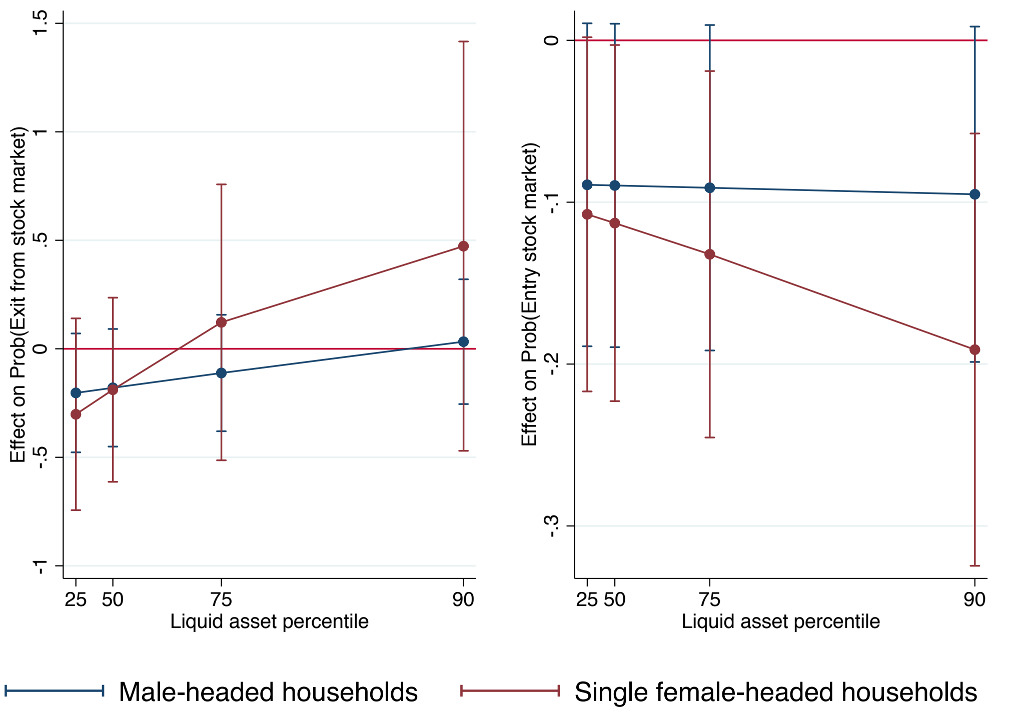
	Single female-headed HH VS Male-headed HH	
	Riskless Act. Sav. (1)	Riskless Act. Sav. (2)
MP	-5225.3	-27388.2
	(45364.9)	(56026.6)
$MP \times$ Head	-5018.0	19281.7
	(52797.8)	39957.9
MP^{*}	0.081***	0.091**
	(0.010)	(0.012)
$MP^{*} \times$ Head	0.306	0.653
	(0.518)	(0.440)
Constant	yes	yes
Other inter. terms	yes	yes
Financial var., lag	yes	yes
Demographics, lag	yes	yes
Household FE	yes	yes
Time FE	yes	yes
Observations	2,389	2,660
R^2	0.03	0.00

Notes: This table presents the results of the fixed effect model in equation (3.5) estimated using the sub-sample of households participating in the stock market over the years 2001-2017. The dependent variable is the riskless asset active saving. In Column (1) the monetary policy exposure variable is the household's lagged value of the liquid assets (instead of the lagged value of the stock investment). Column (2) includes household that participate in the PSID survey for at least two consecutive waves (insted of at least three consecutive waves). The dummy $Head_i$ is equal to 1 if the household head is single and female, it is equal to 0 if the household head is male. The variable $MP_{i,t}$ is the household-level biennial monetary policy shock series constructed in Section 3.3. The variable $MP^{*}_{i,t}$ is the interaction between $MP_{i,t}$ and $W_{i,t-1}$, the monetary policy exposure variable. Standard errors (in parentheses) are clustered at the household level. *, **, and *** represent statistical significance at the 10%, 5%, and 1% levels.

3.D Robustness exercises for section 3.4.1

3.D.1 Monetary policy and stock market participation at different levels of liquid assets

Figure 3.8: Monetary policy and stock market participation decision at different levels of liquid asset distribution - marginal effects



Notes: The figure plots the marginal effect of a contractionary monetary policy shock that increases the US 1-year Treasury by 100 basis points on single female-headed households' (red line) and male-headed households' (blue line) probability on exit the stock market (left panel) and entry the stock market (right panel). The marginal effects are calculated after estimating equation (3.4) and evaluated at different percentile of the liquid assets distribution (25%, 50%, 75%, 90%). All other explanatory variables are evaluated at their sample average. 90% confidence intervals.

As a robustness exercise, we refine the analysis of section 3.4.1 and examine how monetary policy affects stock market participation decisions of single women and male-headed households for different levels of liquid assets.[19] Figure 3.8 presents the effect of a contractionary monetary policy shock on the probability of stock market exit (left panel) and stock market entry (right panel) for single women (red line) and male-headed

[19]In the main analysis, we investigate the marginal effect of monetary policy on exit and entry decisions for both single women and male-headed households considering the average holding of lagged liquid assets.

household (blue line) at different percentiles of the lagged liquid assets distribution.[20] The reported coefficients are the marginal effects (with the 90% confidence bands) evaluated at the sample mean. Figure 3.8 confirms previous section findings: irrespective of the level of liquid wealth, monetary policy does not have a significant effect on the probability to exit the stocl market for both household groups (left panel), and on the probability to entry the stock market for male-headed households (right panel). On the contrary, for single female-headed households the impact of monetary policy on entry decision is increasing in liquid assets, as described in section 3.3.3: the more financial wealth a household holds, the more it is affected by monetary policy due to valuation effects. In fact, the impact of monetary shocks moves from being insignificant for female households in the 25th liquid assets percentile to larger (in absolute value) significant coefficients, up to a -19% at a p-value of 0.02 for households in the 90th percentile.

3.D.2 Monetary policy and stock market participation when allowing for more heterogeneity across household groups

The baseline results of section 3.4.1 are marginal effects evaluated at the sample average of the explanatory variables, obtained by pooling together data for both female- and male-headed households. On the positive side, this choice allows us to compare the effect of monetary policy on the two household groups having the same characteristics. On the negative side, the summary statistics in section 3.2.2 show that the explanatory variable averages for single female- and male-headed households are very different, implying that the sample averages used to calculate the marginal effects in the baseline results are not representative for the average single female- nor male-headed household. Thus, we estimate a different probit model that allows us to calculate marginal effects at the household group-specific subsample average. Therefore, all regressors are interacted with the dummy $Head_i$,

$$\begin{aligned} y^*_{i,t} &= \delta_t + \delta_r + Head_i \times (\alpha X_{i,t-1} + \beta_1 MP_{i,t} + \beta_2 W_{i,t-1} + \beta_3 MP_{i,t} \times W_{i,t-1}) + u_{i,t}, \\ y_{i,t} &= 1\left[y^*_{i,t} > 0\right]. \end{aligned} \tag{3.11}$$

$y_{i,t}$ can be either $Exit_{i,t}$ or $Entry_{i,t}$. $X_{i,t}$ is a vector of control variables that includes household financial characteristics (lagged net worth and family income, change in net worth and family income, total inheritance, dummy for the first mortgage, dummy for the second mortgage), and demographic characteristics (the number of children, the age of the head, the head's age squared, marital status, completed college education, working in the finance industry, total number of family components, home ownership). We also include time- and region of residency- fixed effects (δ_t and δ_r). The dummy $Head_i$ is equal to 1 if the household head is single and female, 0 if it is male. The

[20] We consider the 25th, 50th, 75th and 90th percentiles. They correspond to \$26,147 \$68,800 \$194,035 and \$473,445, respectively, in the sample used to estimate the impact of monetary policy on exit decision. They correspond to \$1,749, \$5,433, \$18,127, \$53,557, respectively, in the sample used to estimate the impact of monetary policy on entry decision.

household-specific monetary policy variable is denoted by $MP_{i,t}$ and $W_{i,t-1}$ is the lagged liquid assets.

Table 3.12: Monetary policy and stock market participation decision - marginal effects at the group-specific sample average

	Exit (1)	Entry (2)
MP if single female-headed HH	-0.058	-0.086**
	(0.185)	(0.038)
MP if male-headed HH	-0.075	-0.097
	(0.166)	(0.067)
Observations	3,649	1,1129
R^2	0.08	0.08
Constant	yes	yes
Other inter. terms	yes	yes
Financial var., lag	yes	yes
Demographics, lag	yes	yes
Household FE	no	no
Time FE	yes	yes

Notes: This table presents the marginal effects of a contractionary monetary policy shock that increases the US 1-year Treasury by 100 basis point on stock market entry and exit decisions of two groups of households, single female- and male-headed households. The variable $Exit_{i,t}$ is a dummy equal to 1 if the household exits the stock market in t and 0 if it stays in; the variable $Entry_{i,t}$ is a dummy equal to 1 if the household entries the stock market in t and zero if it does not. The marginal effects are calculated after estimating the probit regression in equation (3.12) with data from 2001 to 2017 and are evaluated at the group-specific sample average of the explanatory variables. Standard errors (in parentheses) are clustered at the household level. *, **, and *** represent statistical significance at the 10%, 5%, and 1% levels, respectively.

Results are reported in Table 3.12. A contractionary monetary policy shock that increases 1-year Treasury bond by 100 basis point has no impact on the probability of exit the stock market (Column 1), but it does affect the likelihood of single female households to entry (Column 2), decreasing it by 8.6%. This figure is comparable with the baseline result (Table 3.2, Column 2). All in all, this last set of results confirm the baseline findings that only SFHH's entry decisions are affected by monetary policy.

CHAPTER 4

Optimism Gone Bad? Persistent Effects of Traumatic Experiences on Investment Decisions of Households[1]

Chi Hyun Kim

Do memories of highly emotional stock market crashes permanently affect the investment decisions of households? The Initial Public Offerings (IPOs) of Deutsche Telekom during 1996-2000 provide an optimal base to address this question, as the dramatic crash of Telekom share prices shortly after the IPOs, followed by revelation of corruption scandals, stigmatized the German public and is reputedly "the last time Germans invested in stocks" (Handelsblatt, 2016). Using Socio-Economic Panel (SOEP) household survey data, I show that having experienced this event leads to persistently lower stock market participation in the future. In addition, this effect is greater for households that had directly invested in Telekom shares, those being more likely to have high emotional experiences. Finally, I also show that such traumatic experiences on investment decisions have intergenerational consequences, significantly affecting how the next generation invests in the financial market.

Keywords: stock market crashes, emotional tagging, stock market participation
JEL classification: D14, G01, G11, E21

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4.1 Introduction

Although stock investments yield high equity premia, most households do not invest in stocks – a phenomenon known as the *stock market participation puzzle* (Haliassos and Bertaut, 1995; Guiso et al., 2003). Much effort has been put forth to understand the drivers of such behavior, as inefficient investment decisions by households can have negative consequences for their long-term financial well-being. One important finding is that personal experiences are important for investment decisions (Choi et al., 2009; Malmendier and Nagel, 2011). In particular, negative experiences, such as stock market crashes, play a crucial role: after a crash, households are very likely to avoid the stock market (Bucher-Koenen and Ziegelmeyer, 2014; Guiso et al., 2018). What is not yet investigated is whether households recover from such negative events, as they gather new positive experiences while memories of the past crash fade away, or whether they permanently stay away from the stock market.

In this paper, I address these questions by examining a unique stock market crash event in Germany that has the reputation of being "the last time when German households invested in stocks" (Süddeutsche Zeitung, 2014). This event is the Initial Public Offerings (IPOs) of Deutsche Telekom, the German telecommunication giant, during 1996-2000 (which I label as the "Telekom event" throughout the paper). A unique feature of this event is its high media exposure, including an unprecedented advertisement campaign of Deutsche Telekom, which resulted in exceptionally high participation rate of retail investors. However, this high popularity had a downside, as the crash of the Telekom shares over the 2000 to 2002 period provoked strong negative emotions among the German public. I argue that this high emotional turbulence increased the salience among German households (in combination with the negative emotion that was generated along the way) such that German households are still reluctant to invest in stocks; a phenomenon called *emotional tagging* (Laudenbach et al., 2019a).

Using Socio-Economic Panel (SOEP) household survey data, I empirically examine whether the Telekom event persistently affects the stock market investment behavior of German households today. Within a three-part analysis, I investigate the impact of direct experience, the effect of emotional tagging, and the possibility of an intergenerational transfer. To investigate the first aspect, I assume that household heads who were older than 20 during the crash in 2000 are old enough to acknowledge the Telekom event and directly experienced it. After controlling for all relevant socio-economic characteristics, the analysis shows that these households are not only significantly less likely to participate in the stock market in 2016, but they also enter and exit the stock market less than those household heads who were too young to directly experience the event. This finding is striking, as stock market participation of the former group is much higher in levels than the latter.

Also, I examine whether the degree of emotional attachment amplifies the effect of such experiences. Households that had directly invested in Telekom shares (and resulted in real money losses after the crash) have higher emotional attachment to this event and, thus, are less likely to invest in stocks than households that only experienced the Telekom event through the media. As exact information on the Telekom investment of households is not available in the data set, I develop a proxy for being a Telekom investor. Additionally, I construct a control group, which I identify with matching methods, comprising households that have similar characteristics as Telekom investors, but did not enter the stock market during the Telekom event. My results confirm the importance of such emotional attachment on long-term investment decisions. Finally, my analysis shows that emotionally-driven experiences can be passed on to the next generation. I confirm this by analyzing investment decisions of children of Telekom investors.

Literature review

According to the seminal paper of Malmendier and Nagel (2011), households anticipate their life-time experiences of asset returns when they make investment decisions today. One key feature of their analysis is that households have a *fading memory*: far realized returns have less impact on investment decisions than more current ones. However, other studies prove how certain memories have the potential to stick in our memories, regardless of how far this event lies in the past. Ampudia and Ehrmann (2017), for instance, show that experiences of financial crises persistently affect investment decisions of euro area households; Kozlowski et al. (2019) show how crises can permanently shift the way how economic agents assess risk in the future. Within a similar context, Laudenbach et al. (2019b) confirm long-lasting effects of experiences with communism on risk-taking behavior, even after people live in a capitalistic system.

A study by Laudenbach et al. (2019a) proposes emotional tagging to explain why certain experiences tend to stick to our memories, thereby building a base for our economic decision makings. The mechanism that explains this phenomenon is based on the theory of memory formation in neuroscience, which shows how emotions can help the brain to memorize an event or experience (Talarico et al., 2004; LaBar and Cabeza, 2006). If we experience a strong emotional arousal during an event (either positive or negative), then we are likely to remember this emotion when we are exposed to a similar situation in the future. This emotion, in turn, influences the way we respond to the future event.

In the stock market, emotions play a key role: many studies confirm how human instincts and emotions that seemingly affect human behavior in economic decision makings, such as confidence, optimism (pessimism), or trust, can explain a substantial part of the dynamics in the financial markets (Keynes, 1936; Shiller, 2000; Brunnermeier and Nagel, 2004). Therefore, emotional tagging has the potential to explain how investors behave in the stock market, depending on their previous experiences. My study contributes to this literature by showing how an emotionally-driven stock market event can function as a single "rare disaster" that can lead to persistent low stock market participation of households (Haliassos and Bertaut, 1995; Campbell, 2006; Barro, 2006, 2009; Alan, 2012).

The remainder of this paper is structured as follows. Section 4.2 introduces the Telekom event and explains why this event is attached with high emotional turbulence. Based on these insights, I derive testable hypotheses. Thereafter, in section 4.3, I use SOEP household survey data to present descriptive statistics on German households' stock market investment behavior during 1990 and 2016. Finally, section 4.4 introduces the empirical framework, identification strategy of the Telekom event, and results. Section 4.5 concludes.

4.2 The Telekom event

In this section, I provide a detailed description of the Initial Public Offerings (IPOs) of Deutsche Telekom. First, I introduce the timeline of the rise and fall of the Telekom shares (T-shares) between 1996 and 2002. Then, I discuss the reasons why the Telekom event has the potential to be tagged with high emotions for the German public, even twenty years later. Afterwards, in order to visualize this emotional turbulence, I conduct a sentiment analysis on newspaper articles and analyze its development along with the timeline of the Telekom shares after the crash. Based on these insights, I develop two research hypotheses.

4.2.1 Timeline of the Telekom event

4.2.1.1 Telekom's advertisement spectacle

Deutsche Telekom AG, a former entity of the German federal government's postal administration (Deutsche Bundespost), was privatized in 1995 and made its IPO in November 1996. However, at the beginning of 1996, the company struggled with its very negative reputation amongst the general public due to its new tariff structure and high local call prices (König, 1997). Deutsche Telekom acknowledged that their IPO could only be successful if they manage to re-polish their image. Based on these insights, Deutsche Telekom hired *Dewe Rogerson*, a British financial and corporate communications consultancy, and invested around 200 million Deutsche Mark (around 98 million euro) in an advertising campaign. In particular, the company conducted a very aggressive strategy to capture the attention of potential retail investors from the general public. Their aim was to make the T-shares the "people's share." As a result, the IPO of November 1996 was a great success with around 1.9 million retail investors investing in 285 million Telekom shares.

Three main factors explain the great success of the T-shares. First, Deutsche Telekom did not shy away from directly contacting the public. The company sent personal invitations to all Telekom customers and asked them to register to the Deutsche Telekom's "Aktien-Informations-Forum" (AIF).[2] The AIF sent company booklets and information sheets for "stock investment for beginners." Within one month, the number of registrations reached the million threshold and, by November 1996, almost three million people were registered with the AIF. Additionally, Deutsche Telekom offered discounts to retail investors to facilitate the stock market entry of people who had never invested in stocks before. After the first IPO, Deutsche Telekom stayed in close contact with their retail investors by sending them letters, personally written by the director Ron Sommer.

Second, Deutsche Telekom used media presence to improve their brand image. The company placed television commercials to awake positive emotions of the audience:

[2]Translated: Stock-Information-Forum.

they streamed "High-Tech and High-Touch spots" on television to show off their advanced technical capability and, starting from August 1996, an advertisement for the T-shares, where Manfred Krug – a very popular actor at that time – introduces himself as a stock investor and announces that he will definitely invest in the T-shares.[3] Further, Deutsche Telekom also utilized newspapers and magazines to attract the attention of the public with information about the operational details of the company. Deutsche Telekom highlighted its "customer-oriented" operation design and great growth potential.

Last, but not least, director Ron Sommer also played a crucial role in gaining public trust. As the new head of the Deutsche Telekom after the 1995 privatization, he represented freshness, innovation, and a successful future for Deutsche Telekom as an upcoming "Global Player" in telecommunications. His progressive investment plans and confidence initiated an optimistic sentiment among retail investors. With confidence, he assured that "the T-shares are secure as an inheritable retirement pension supplement." Very quickly, Ron Sommer became the face of Telekom shares.

4.2.1.2 Dotcom bubble crash

Following the first IPO in November 1996, the price of T-shares continuously increased, as shown in Figure 4.1. The company launched two further IPOs in June 1999 and June 2000, where the participation rate of retail investors was remarkably high, with an increasing trend over the three IPOs (see Table 4.1). However, the steep rise of the Telekom shares did not last long. After peaking at 103.90 Euro in March 2000, both global factors and internal company problems led to a very fast crash of the Telekom share prices. In June 2002, Telekom shares were worth only 9 Euro per share.[4]

Table 4.1: Allocation of the IPOs of the Telekom shares

	DT1	DT2	DT3
Date	18.11.1996	28.6.1999	19.6.2000
Issue amount (€)	713.7 mio	280.9 mio	200 mio
Retail	43%	54%	70%
Institutions	57%	46%	30%

Notes: DT1, DT2, and DT3 present the first, second and third IPO of the Deutsche Telekom, respectively. *Source*: Deutsche Telekom Investor Relations

However, the crash of the T-shares was not an individual case, rather it was part of the overall situation in the global stock market. Two major global events happened during the three IPOs of Deutsche Telekom, one being the Asian financial crisis starting in July 1997 and the other being the Dotcom bubble burst in the early 2000s. While Deutsche Telekom was not severely hit by the Asian financial crisis, the bust of the

[3]Deutsche Telekom also increased its media presence by investing in sports sponsorships (cycling, soccer) and tech-exhibitions (such as the CeBit-exhibition).

[4]Since then, the price of the Telekom shares stayed stable around the 15-20 Euro mark (status as of 2021).

Dotcom bubble hit the German stock market very hard as a whole. In particular, the market segment “Neuer Markt” (“the new market”) of the German stock market, which was introduced in 1997 with a focus on tech companies, had to deal with serious corrections of their overrated market performances – Deutsche Telekom being part of this market segment.

Figure 4.1: Price of the Telekom sharess

Notes: *Source*: Datastream.

4.2.1.3 Internal problems

Additionally, rumors were spreading more and more with regard to Deutsche Telekom’s balance sheet. It started with an article in *Der Spiegel* in August 1998, which states that Deutsche Telekom over-valued its real estate holdings by anticipating price increases during the real estate boom. Calculations predicted the value adjustment according to this mispricing to be between 3.5 and 4.2 billion Deutsche Mark (Der Spiegel, 1998). After these rumors were confirmed, on July 16, 2002, Ron Sommer was forced to resign from his position as the directorate of Deutsche Telekom, which was a clear signal of the end of the T-shares era. Afterwards, the company made changes in their operational strategy with strict austerity policies.

4.2.2 Sentiment analysis and hypotheses

Although the entire German stock market suffered a crash during the Dotcom bubble burst, the Telekom event is probably the most prominent event, resulting in extensive media coverage and, consequently, high interest of the German public. The very fact that 2 million retail investors participated in the first IPO (König, 1997) reflects the scale of this event, demonstrating how many people were affected by the crash in 2000. Further, the company's corruption scandals provoked strong negative emotions among the German public, as German households had high faith in the business model of the former public cooperation.

Since the Telekom event was quite prominent in the media, a sentiment analysis of newspaper articles is well suited to visualize the public's sentiment. I use digitalized newspaper articles of the *Handelsblatt*, which is Germany's leading financial newspaper. As a first step, I identify all articles between 1998 and 2002 that report information on Telekom shares.[5] Next, I clean the text to minimize noise that can distort the true sentiment of the content. After the data-cleaning process, I apply a lexicon-based sentiment analysis using a publicly available German-language lexicon source called *SentiWS*. This lexicon lists positive and negative sentiment bearing words weighted within the interval of -1 and 1, with 1 being absolute positive and -1 being absolutely negative. For my purpose, I first calculate the average tone of all newspaper articles in a given year. Then, I calculate the yearly growth rate of this sentiment measure, as I am interested in how sentiment has evolved since the crash.[6]

Figure 4.2 presents the results. Besides the change in sentiment, I also plot the change in Telekom share prices, which serves as an "objective" benchmark of the evolution of the Telekom event. Right after the crash in 2000, both price and media sentiment experience a steep drop. However, while the drop in prices seem to stabilize thereafter, sentiment continues to drop in an even more substantial manner in 2001. This fits very well with the timeline of the Telekom event as Deutsche Telekom experienced massive criticism due to poor management and corrupt behavior by the executive board, both of which may have worsened the sentiment of the German public. Additionally, the fact that sentiment decreases in a much stronger manner than the price itself further supports the argument that emotional turbulence was high.[7]

Based on these insights, I hypothesize that the Telekom experience is highly attached with emotions that enhanced the emotional tagging of German households.

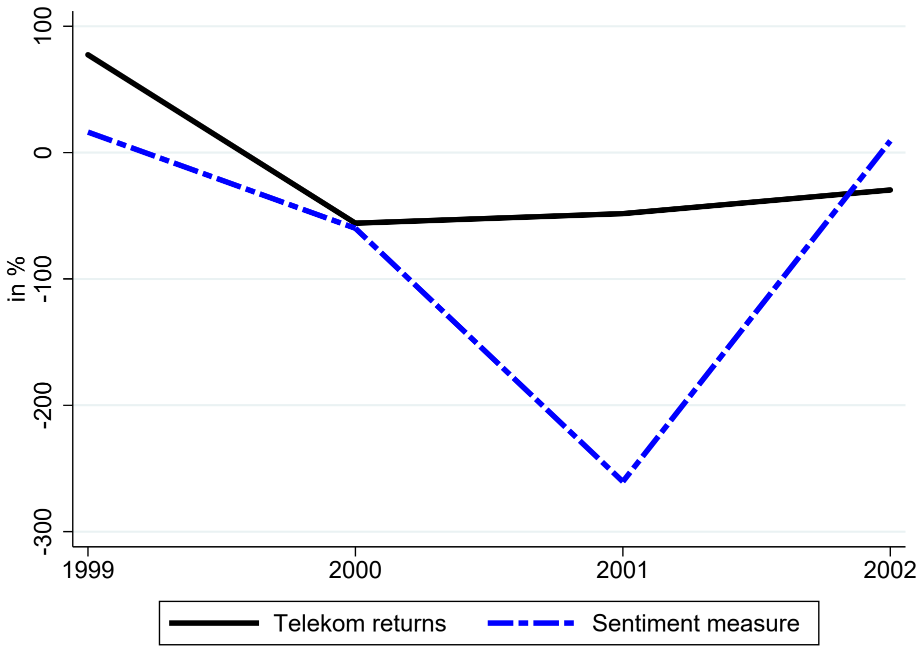
Hypothesis 1: The Telekom event has persistently affected investment decisions of German households.

[5] I concentrate on the crash period, as the negative emotions during this period dramatically enhanced the emotional tagging of German households.

[6] Detailed description of the method is provided in Appendix 4.A.

[7] As an extension, it would be also interesting to include additional newspapers and even television coverage of Deutsche Telekom, as *Handelsblatt* may be somewhat neutral in their reports compared to other populist media sources like the *Bild newspaper*. Nevertheless, to some degree, this exercise shows the lower bound of media sentiment and including other sources may amplify my results.

Figure 4.2: Change in media sentiment vs. price



Source: Own calculations using newspaper articles of the *Handelsblatt*.

After the crash, German households that had invested in Telekom shares ended up filing mass lawsuits against the company (Der Spiegel, 2001).[8] This implies that active investment in Telekom shares, which led to real money losses for many households, may have triggered stronger emotional turbulence than for households that only experienced the Telekom event through the media. Therefore, as a complementary hypothesis,

Hypothesis 2: Households that had actively invested in Telekom shares have higher emotional attachment to the event and, thus, are less likely to participate in the stock market today than households that did not invest.

4.3 Descriptive statistics

In this section, I provide descriptive statistics of stock market participation behavior of German households during 1990 and 2016 and connect them with the insights of section 4.2. To do so, I use Socio-Economic Panel (SOEP) data provided by DIW Berlin. After a description of the data, I present the development of stock ownership between 1990 and 2000, which includes the three IPOs of the Telekom shares. Afterwards, I analyze stock market participation behavior of German households after the Telekom event.

[8]Interestingly, they even accused the actor Manfred Krug, who was the main figure of the TV-commercial, for lying to the public. Manfred Krug, even though he had nothing to do with the crash itself, ended up giving an official apology to the public during a newspaper interview several years after the crash, noting that his appearance in the commercial was "the biggest failure of his career" (FAZ, 2007).

4.3.1 The Socio-Economic Panel

The SOEP is a nationally representative household survey data that contains not only information on households' socio-economic characteristics, such as family composition, employment, home ownership, and income, but also on their holdings of financial assets. Overall, in each wave, the SOEP provides information on ownership of six financial asset categories of the previous year: saving accounts, home ownership saving contracts ("Bausparvertrag"), life insurance policies, fixed-interest securities (including saving bonds issued by banks, mortgage-backed bonds, and government bonds), stocks held directly or through mutual funds, and ownership (or shares) of non-listed firms.[9] In addition, the SOEP also provides a wealth module since 2002, where they ask respondents to report their holding amounts in different asset classes. This information is gathered every five years.

The SOEP is the only survey that includes information of German households' investment behavior for the pre-2000 period, when the three IPOs of the Telekom shares took place.[10] Additionally, the SOEP has a panel dimension that allows me to track households over time. This is very helpful for identifying households during the current waves that were exposed to the Telekom event in the past. Therefore, SOEP data are best suited for my data analysis.

However, despite these advantages, there are also some drawbacks. Before 2001, households were asked to report their ownership in fixed-interest securities (including saving bonds issued by banks, mortgage-backed bonds, and government bonds) together with their stock holdings. Only after 2001 did the survey distinguish between these two asset categories. Nevertheless, ownership of these two asset categories are highly correlated and, thus, can still serve as a reliable proxy for stock ownership.[11]

4.3.2 Stock market participation during the Telekom event

Serving as a proxy for stock ownership, Figure 4.3 presents the development of security papers ownership of German households between 1990 and 2000.[12]

On average, 27% of German households reported owning stocks between 1990 and 2000.[13] Interestingly, stock ownership doubled over these ten years. In 1990, the stock ownership rate was 21%; in 1996 it started increasing exponentially, then peaking at

[9]In Appendix 4.B, I summarize asset ownership of German households between 1990 and 2016.

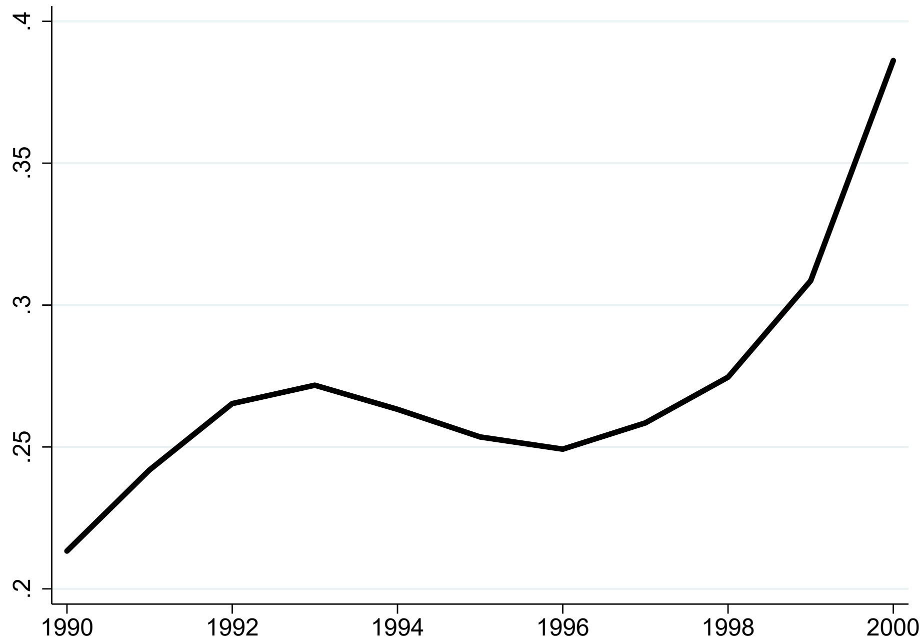
[10]Besides the SOEP, there are two surveys that capture investment behavior of German households. The SAVE study is a cross-sectional study that was initiated by the Munich (formerly Mannheim) Center for the Economics of Aging (MEA) to explicitly analyze investment and saving behavior of German households. This study started in 2001 and was terminated in 2011. The Deutsche Bundesbank also started its own survey, the Panel of Household Finances (PHF), which is conducted every third year since 2011. This survey provides the most detailed information on households' portfolio composition, along with their socio-economic characteristics.

[11]For the comparison of these ownership of these two asset classes, see Figure 4.9 in Appendix 4.C.

[12]Note again that this indicator includes both ownership of stocks and fixed-interest securities, as separate information is not available prior to 2001. Nevertheless, for simplicity I define this proxy as stock ownership.

[13]Note that this indicator includes both direct and indirect holdings of stocks.

Figure 4.3: Stock market participation rate (proxy): 1990-2000



Source: Own calculations using SOEP data.

41% in 2000. By comparing summary statistics of stock owners of 1990 and 2000, Table 4.2 shows that the overall picture of stock investors changed as well. In 2000, stock investors (i) are more likely to be a woman, (ii) are slightly younger, (iii) are more likely to come from the former GDR, and (vi) have lower total family income compared to stock investors in 1990.

To better understand who entered the stock market between 1996 and 2000 (thus having a positive probability of having invested in Telekom shares), I analyze stock market participation of different household groups. As a first step, I distinguish between three income groups: the bottom 50%, 50% to 90%, and the top 10% of the total income distribution. This distinction is useful in understanding whether households with higher income were more likely to invest in Telekom shares, as income positively correlates with household stock market participation. Figure 4.4 presents the results.

From 1990 to 1995, the development of stock ownership is quite stable for the bottom 90% of the income distribution, while for the top 10%, stock ownership decreased slightly in 1994. Thereafter, stock ownership increases across all income groups, but differently in timing. While households in the top 10% of the income distribution started to increase their ownership in 1994 – before the IPOs of T-shares – stock ownership among the bottom 90% increases after 1999, so rather during the second and third IPOs.

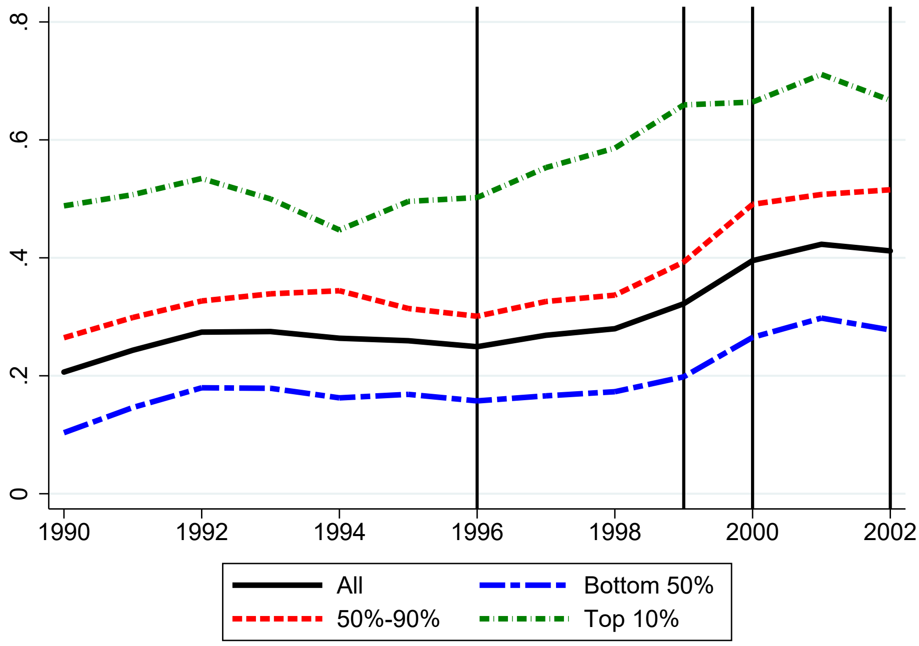
Additionally, I also examine whether stock market participation differs across birth cohorts, as age (along with its linked economic position) correlates with stock invest-

Table 4.2: Summary statistics

	1990		2000	
	mean	sd	mean	sd
Demographics				
Sex	0.73	0.44	0.68	0.47
Age	46.51	13.53	46.32	12.70
Married	0.65	0.48	0.58	0.49
Number of children	0.48	0.84	0.45	0.81
Household size	2.52	1.23	2.32	1.19
Highschool degree (*Abitur*)	0.24	0.43	0.29	0.45
University degree	0.26	0.44	0.31	0.46
Employment	0.73	0.44	0.73	0.45
Former FRG/GDR	0.94	0.24	0.82	0.39
Total family income	44295.96	28721.20	43193.14	23967.40
Financial decisions				
Riskl. asset ownership	0.98	0.15	0.97	0.17
Home ownership	0.51	0.50	0.52	0.50
Asset income	2352.86	6559.14	2670.70	8002.18
Risk aversion	4.60	1.86	4.70	1.77
Fin. risk taking	2.80	2.09	2.99	2.06
Observations	1038		3879	

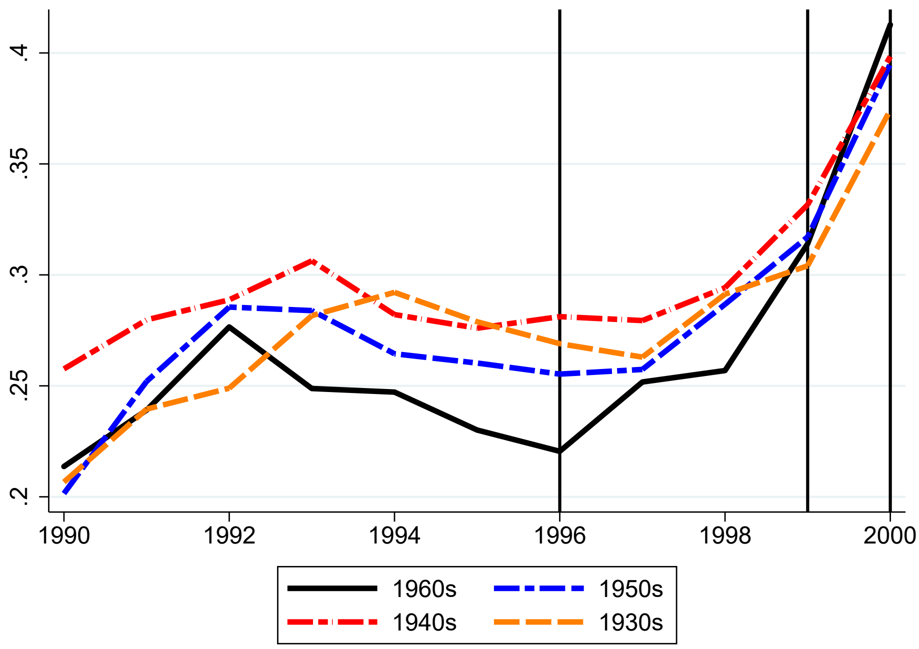
Notes: This table shows the summary statistics of German stock owners' socio-economic characteristics in 1990 and 2000. *Source*: Own calculations based on SOEP data.

Figure 4.4: Stock ownership by income group



Source: Own calculations using SOEP data.

Figure 4.5: Stock ownership by cohort

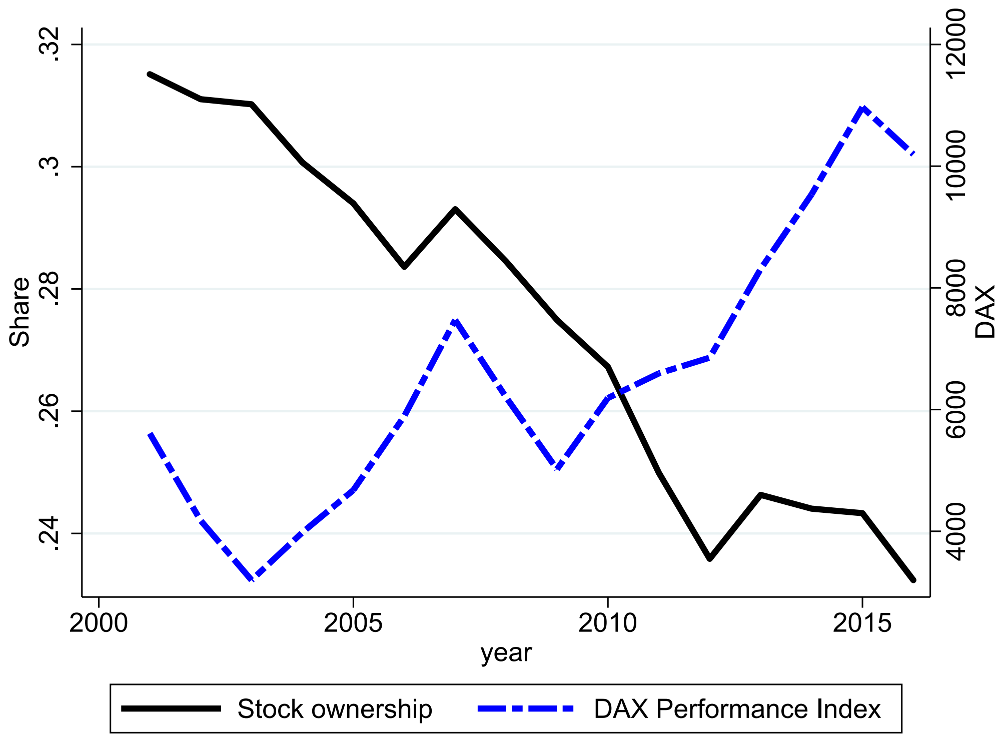


Source: Own calculations using SOEP data.

ments. I do not consider household heads who are born after 1970, as stock market participation among this group is very low in the 1990s due to their age. Results for the different cohorts are presented in Figure 4.5. With regard to the different cohorts, I do not observe distinct differences. Stock market participation increases rapidly during the period of the three Telekom IPOs in 1996-2000 for all groups. In particular, the 1960s cohort, which had the lowest stock ownership, experiences the largest increase and outperforms the other cohorts. Overall, my descriptive results indicate that the Telekom event was not just popular for specific income classes or cohorts but rather was a universal phenomenon, thus supporting the mass nature of the event.

4.3.3 Stock investment of German households after the Telekom event

Figure 4.6: Stock ownership: 2001-2016



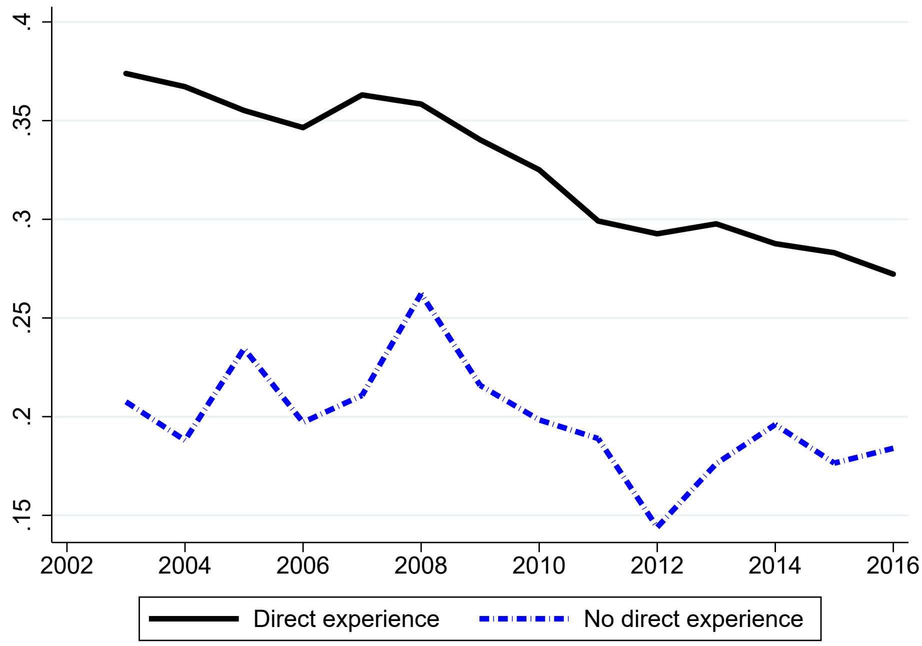
Notes: This figure presents the share of German households that report to have positive holdings of stocks (direct- or indirect holdings). Note that starting from 2001, the SOEP provides information of stock holdings separately from fixed-income security ownership. *Source*: Own calculations based on SOEP data.

Since the crash in 2000, stock ownership continuously decreased without any remarkable recovery through 2016 (Figure 4.6). In 2007 and 2013, stock ownership seems to slightly increase, however the magnitude is rather small and not long-lasting. This is remarkable, given the fact that stock prices have steadily increased, despite crash periods during the Global Financial Crisis (see blue dashed line).[14]

[14]Daily time series of the DAX Performance Index is presented in Figure 4.11 in Appendix 4.C.

To shed light on whether this persistent decrease in stock market participation relates to the Telekom experience, I distinguish between households that were at least twenty years old during the crash (solid black line), as they are more likely to have directly experienced the crash than household heads younger than twenty (dashed blue line). Results are presented in Figure 4.7. Since 2000, stock market participation has gradually decreased for those with direct experience, showing no recovery. Compared to this, households with young heads seem to be rather unaffected by the Telekom event, as stock market participation increases until 2008. For this younger generation, the Global Financial Crisis seems to have a larger impact on their stock market participation decisions, while the older generation does not seem to exhibit any noticeable adjustments during this event.

Figure 4.7: Stock ownership by Telekom experience



Source: Own calculations using SOEP data.

An additional feature is that stock market participation did not decrease drastically in the immediate aftermath of the Telekom event; rather it fell slowly. This rather slow-moving stock market participation rate of the experienced group may be due to the fact that many retail investors did not immediately sell their T-shares, but rather kept them in their portfolio (even though they did not actively invest in stocks again) or sell them at a later point of time.[15] Since the T-shares were only worth 9 EUR each, selling them would have required German households to bear all the losses. Also, as Deutsche Telekom may be "too big to fail," investors may have expected a recovery in the mid- to long-run. However, this never happened, which may have entrenched the

[15]Figure 4.10 in Appendix 4.C presents entry- and exit-rates of German households after the crash in 2000. I also observe how household heads who did not directly experience the Telekom event invest in stocks in a more active manner than household heads with Telekom experience.

negative emotions of Telekom investors. They can still track the price development of their T-shares, which, in 2021, were worth about 15 EUR per share.

4.4 Empirical analysis

In this section, I empirically examine the effect of the Telekom event on the investment decisions of German households in 2016. I first introduce the baseline model and then explore two possible channels through which the Telekom event can affect long-term investment decisions. First, I test whether households that directly experienced the Telekom event invest differently than households that did not. Second, I analyze whether emotional attachment matters for the Telekom event. Finally, I investigate the possibility of an intergenerational transfer of these effects.

4.4.1 Baseline model

My baseline empirical model is a probit. Since I am interested in investment behavior of German households in the current period, I restrict my analysis to the year 2016, which is the latest period available to me.[16] The model is

$$y_i = \alpha + \beta_1 TE_i + \beta_2 A_i(\lambda) + \gamma' x_i + \delta_r + \varepsilon_i, \tag{4.1}$$

where y_i is an indicator for the stock market investment behavior of a German household i in 2016. I use data from SOEP that indicates households' ownership of stocks held directly or through mutual funds and calculate three indicators that visualize (i) stock market participation SMP_i, (ii) stock market entry $Entry_i$, and (iii) stock market exit $Exit_i$. SMP_i equals one if household i reports to directly or indirectly own stocks and zero if not. $Entry_i$ equals one if household i does not participate in the stock market in 2015, but participate in 2016, and zero if the household does not participate in the stock market in both waves . Note that households that participate in the stock market both in 2015 and 2016 are not included in this variable. $Exit_{it}$ is equal to one if household i owns stocks in 2015, but not in 2016, and zero if the household participates in the stock market in both consecutive years. In this variable, households that participate in the stock market in neither year are not considered.

On the right side of the equation, I have four sets of variables. First, x_i contains a set of socio-economic control variables, such as age of the household head, their gender, and their education degree, as well as household-level information, such as household size, income, and wealth. In addition, I include households' subjective evaluation of their willingness to take financial risk.

Second, in the spirit of Malmendier and Nagel (2011), I also include a variable that incorporates the whole information set of past returns of household heads since their birth year with a weighting function. The inclusion of this variable is important to

[16]Note that survey respondents report their stock ownership for the previous year. Therefore, information on stock ownership is obtained by the SOEP 2017 wave. Data on households' socio-economic characteristics that are used as control variables in the regression analysis are obtained by the SOEP 2016 wave.

cleanly isolate the Telekom effect from the general stock market situation. As already explained, the Telekom event happened in parallel with the Dotcom bubble burst. Therefore, I need to be sure that my identification of the Telekom event is not driven by other factors in the stock market.

The formula to calculate this variable A_i is

$$A_i(\lambda) = \sum_{k=1}^{age_i - 1} \omega_i(k, \lambda) \ R_{2016-k}, \text{where} \tag{4.2}$$

$$\omega_i(k, \lambda) = \frac{(age_i - k)^\lambda}{\sum_{k=1}^{age_i - 1} (age_i - k)^\lambda}. \tag{4.3}$$

The weighting function $\omega_i(k, \lambda)$ determines how households anticipate past experiences of asset returns for their current investment decisions. Therefore, it depends on how far an experienced return (growth) R_{2016-k} is for household i in 2016 (using $age_i - k$ to calculate the distance) and a shaping parameter λ. If λ is positive, nearer experienced returns are more strongly weighted than returns that are far away, and vice versa. For my purpose, I adopt the conventional results in the literature and assume $\lambda = 1$.[17] Since I am interested in stock market investment decisions of German households, I use the yearly growth rates of the DAX index for R_{2016-k}.

Finally, I introduce TE_i, which visualizes the exposure of German households to the Telekom event. For the identification of TE_i, I exploit the panel dimension of the SOEP and track households that are available in the 2017 survey wave that experienced the Telekom event during 1996-2000. By doing so, I consider two dimensions: the effect of direct experiences (*Hypothesis 1*) as well as the effect of emotional attachment (*Hypothesis 2*). A detailed description of the identification strategy of these two channels is presented in the following subsections and also summarized in Table 4.8 in Appendix 4.D.

4.4.2 The effect of direct experiences

Households that experienced the Telekom event in an active manner are more likely to be affected than households that only experienced it indirectly. As a first identification strategy, I assume that household heads, who were at least twenty years old in the year of the crash, had a more direct experience as they may have not only noticed the media presence of the T-shares, but were also more likely to have actively invested. In turn, household heads younger than the age of twenty were too young to actively notice the severeness of the situation at that time.[18]

[17] Robustness exercises with different values of λ are presented in Table 4.9 in Appendix 4.D.

[18] This threshold is also chosen based on the fact that data on household heads younger than twenty is scarce and, thus, choosing a smaller threshold may lead to unreasonable results. Robustness exercises with a different threshold are provided in Table 4.10 in the appendix. Section 4.4.4 addresses the issue that children may also have experienced the Telekom event through their parents.

TE_i is thus defined as:

$$TE_i = \begin{cases} 1 & \text{if age} \geq 20 \quad \text{in} \quad 2000 \quad \text{and} \\ 0, & \text{otherwise.} \end{cases} \tag{4.4}$$

In Table 4.3, I present the marginal effects at the mean of TE_i estimated with equation (4.1) for the three different dependent variables introduced in subsection 4.4.1. Besides the general stock market participation indicator, entry- and exit-behavior sheds additional light on whether households with direct Telekom experience also adjusted their *active* investment decisions in the stock market. Again, note that I estimate the model for 2016.

Table 4.3: The effect of direct experience: marginal effects at the mean

	SMP (1)	Entry (2)	Exit (3)
Direct experience (β_1)	-0.59^{***}	-0.92^{***}	-0.89^{***}
Stock returns (β_2)	-2.87	30.04^{*}	21.47
Controls	yes	yes	yes
Constant	yes	yes	yes
Region FE	yes	yes	yes
Observations	4466	2577	1889
Pseudo R^2	0.19	0.11	0.16

Notes: This table presents the marginal effects of TE at the mean, with TE capturing direct vs. no direct experience of the Telekom event. *, **, and *** represent statistical significance at the 10%, 5%, and 1% levels, respectively.

For all three specifications, the Telekom dummy is highly significant. First, households with direct experience invest significantly less in stocks than households that are not. This result confirms my first hypothesis that the Telekom event has significantly affected the stock market investment decisions of German households. This is remarkable as unconditional stock market participation is substantially higher for older household heads (see Figure 4.7). In addition, the directly experienced group is also rigid in their *active* investment behavior, i.e. they are much less likely to enter or exit the stock market.[19]

It is also worth mentioning the effect of the past experienced DAX returns (β_2). My results show that past stock returns typically do not have a significant effect on investment decisions of German households. It is only for entry decisions that it has

[19] One might ask why exit rates are lower for households with direct experience, as one may think that they are more likely to exit the stock market. However, the results are for 2016 and not directly after the crash. Further, lower entry- and exit rates can be interpreted as generally more rigid behavior in stock investment.

a weak significant impact. At first glance, this result seems to contradict the message of Malmendier and Nagel (2011) and Ampudia and Ehrmann (2017), who show that households anticipate their life-time experiences of returns when making investment decisions today. However, they also discuss in their paper that a parametric approach to determine the shape of the weighting function may be problematic, as it does not account for extreme events. Ampudia and Ehrmann (2017) confirm this concern by showing that extreme events, like the GFC, can have persistent effects on the investment decisions of euro area households. Therefore, the long-lasting effect of the Telekom event serves as an indication that German households seem to weight this event so strongly that they do not anticipate fluctuations in the stock market for their stock market participation decisions.

4.4.3 The effect of emotional attachment

As a next step, I address my second hypothesis and focus on the group of households that directly experienced the Telekom event. Within this group, I test whether the degree of emotional attachment amplifies its persistent effect on long-term investment behavior.

For identification, I assume households that actively invested in the T-shares are likely to be more emotionally attached to the event, as they have experienced losses to their wealth. This group of households is defined as the "Telekom investors." In order to cleanly identify this effect, I use synthetic control methods (Abadie et al., 2010) and contrast the Telekom investors (which is the treatment group, with emotional attachment being the treatment) with a control group of households that (i) experienced the Telekom event (as defined in section 4.4.2), (ii) exhibit similar socio-economic characteristics as the Telekom investors, but (iii) did not invest in the T-shares.

Unfortunately, the SOEP does not provide information on the detailed portfolio composition of households, meaning that exact information on T-shares investment is not available. Therefore, I must make assumptions that provides a proxy for T-shares investment. I utilize information from the SOEP and define that households have a high probability to have invested in the T-shares if they first entered the stock market during the first three IPOs. This group of households is defined as the "Telekom investors." Next, I use their characteristics in the year 2000, apply propensity score matching on the control pool (households that have experienced the Telekom event without direct investment), and search for the treatment group's nearest neighbor.[20] Finally, I track these two groups to the year 2016. The Telekom indicator TE_i is thus

$$TE_i = \begin{cases} 1 & \text{if direct experience } \textit{and} \text{ investment in T-shares,} \\ 0 & \text{if direct experience but } \textit{no} \text{ investment.} \end{cases} \tag{4.5}$$

[20] I use Mahalanobis propensity matching setting a caliper to 0.1 of the standard deviation. Results remain robust when taking other values.

Table 4.4: The effect of emotional attachment: marginal effects at the mean

	SMP (1)	Entry (2)	Exit (3)
Emotional experience (β_1)	-0.12^{**}	-0.18^{***}	-0.02
Stock returns (β_2)	-16.82	36.50^{*}	17.03
Controls	yes	yes	yes
Constant	yes	yes	yes
Region FE	yes	yes	yes
Time FE	yes	yes	yes
Observations	468	227	183
Pseudo R^2	0.27	0.36	0.33

Notes: This table presents the marginal effects of TE at the mean, with TE capturing the emotional attachment to the Telekom event. SMP stands for stock market participation. *, **, and *** represent statistical significance at the 10%, 5%, and 1% levels, respectively.

Table 4.4 presents the results for (1) stock market participation, (2) stock market entry, and (3) stock market exit. In addition to direct experience, the results indicate that emotional attachment to the Telekom event matters as well. Households that have the highest potential to have actively invested in T-shares are 12% less likely to invest in stocks and are 18% less likely to enter the stock market. However, households with high emotional attachment to the Telekom event do not significantly differ from other households when they make exit decisions.

4.4.4 The effect of indirect experiences: Intergenerational transfer

The previous exercises show how (i) households with *direct* experience of the Telekom event invest significantly less in stocks compared to households that do not, and (ii) households that have higher emotional attachment to the event due to active investment in T-shares invest differently compared to households that have just witnessed the event. Finally, I also address an additional hypothesis on whether experiences of parents can transfer to children, thereby indirectly influencing their investment decisions. In the literature, studies confirm the relevance of intergenerational transmission with respect to trauma (Dekel and Goldblatt, 2008; Lev-Wiesel, 2007), risk- and trust attitudes (Charles and Hurst, 2003; Dohmen et al., 2012), and poverty (Bird, 2013). The SOEP provides a good base for this kind of intergenerational exercise, as the data tracks children of survey respondents over time.

For this exercise, I distinguish between two groups of children, differing with respect to the nature of the parental experience of the Telekom event. The first group comprises children of Telekom investors, where I assume that the negative emotional experience of their parents is transmitted to the children, thus with the potential to indirectly affect their investment decisions. The second group includes children of the so-called "experienced investors," who I define as household heads who regularly invest in the stock market. I approximate this based on how frequently they report owning stocks: if they report investments in more than half of survey waves, then I assume that they regularly invest in stocks. My assumption is that this group has the least emotional attachment to the Telekom event, as they are more experienced investors and, thus, are less likely to be affected by a single stock market event.[21] Therefore, experienced investors can be Telekom investors at the same time, but they are less likely to be emotionally exposed to the event.

Then, I follow these two groups of children to the 2017 survey wave and analyze their stock market participation, once they become household heads themselves.[22] I use the same empirical model of equation (4.1), where TE_i is defined as follows:

$$TE_i = \begin{cases} 1 & \text{if children of Telekom investors,} \\ 0 & \text{if children of experienced investors.} \end{cases} \tag{4.6}$$

Table 4.5 presents the socio-economic characteristics of these two groups. Children of Telekom investors and of experienced investors differ greatly in their characteristics, with children of experienced investors having much higher education degree, as well as higher income and wealth. Particularly interesting is the difference in their stock ownership: while almost 60% of the children of experienced investors own stocks, only 18% of the Telekom investor children do. This difference is substantial compared to other asset ownership, such as riskless assets and home ownership. Finally, their willingness to take financial risk differs greatly as well.

Finally, Table 4.6 presents the results of the regression analysis for stock market participation using equation (4.1).[23] The regression analysis shows that children of Telekom investors are 53% less likely to participate in the stock market, after controlling for all relevant control variables. This difference is statistically significant, thus confirming the effect of indirect experience of such emotional events on investment decisions.

[21] Table 4.11 in the appendix presents summary statistics of experienced investors' socio-economic characteristics in comparison to (i) the German average and (ii) Telekom investors.

[22] I restrict the children, who are household heads of a new household, to be younger than 45. This decision is based on the fact that I want to ensure that children were sufficiently young during the Telekom event that they were not actively involved in investment decisions of the household. If I do not restrict the sample, the results remain robust.

[23] As data is scarce for children in general, there is not enough data points to analyze their entry- and exit behavior.

Table 4.5: Summary statistics of children, survey year 2017

	Experienced investor children		Telekom investor children	
	Mean	SD	Mean	SD
Demographics				
Sex	0.51	0.50	0.50	0.50
Age	33.14	6.71	30.52	4.33
Married	0.28	0.45	0.23	0.43
Number of children	0.36	0.75	0.29	0.65
Household size	1.96	1.04	1.88	0.95
High school degree (Abitur)	0.56	0.50	0.32	0.47
University degree	0.63	0.48	0.33	0.47
Degree in Economics	0.13	0.34	0.03	0.17
Employment	0.78	0.42	0.81	0.40
Former FRG/GDR	0.79	0.41	0.61	0.49
Total family income	39250.64	23907.93	33110.62	18839.40
Financial decisions				
Riskl. asset ownership	0.93	0.25	0.91	0.29
Stock ownership	0.57	0.50	0.18	0.39
Home ownership	0.28	0.45	0.26	0.44
Fin. risk taking	3.36	1.77	2.95	1.83
Wealth				
Financial wealth	39299.97	64719.38	15988.60	30009.77
Net wealth	138853-21	203192.91	78415.76	124049.70
Home equity	51113.51	117010.20	42505.03	93879.09
Real estate	6918.30	60340.55	6841.12	30149.25
Asset income	1162.49	2500.34	497.70	1550.65
Observations	230		69	

Notes: Telekom investors are defined as household heads who first owned security papers during the three IPOs of the Deutsche Telekom shares. Experienced investors are household heads with over 50% security papers ownership during their participation in the survey. For 2002 and 2017, the SOEP wealth module is available. *Source*: Own calculations based on SOEP data.

Table 4.6: Determinants of stock market participation of children, marginal effects at the mean

	SMP
Direct experience (β_1)	-0.53^{***}
Stock returns (β_2)	152.93
Controls	yes
Constant	yes
Region FE	yes
Observations	167
Pseudo R^2	0.38

Notes: This table presents the marginal effects of being a child of a Telekom investor (at the mean) on stock market participation. SMP stands for stock market participation. *, **, and *** represent statistical significance at the 10%, 5%, and 1% levels, respectively.

4.5 Conclusion

I utilize insights of a unique stock market event in Germany to examine whether emotionally-tagged events can have long-term effects on the investment behavior of households. My empirical analyses confirm its effect by showing that it is not just the experience itself, but also the degree of emotional attachment that matters for determining the long run investment patterns of affected households. Further, I also provide first-order evidence on the far-reaching consequences of such events by showing how the children of parents emotionally affected by the Telekom stock crash are still less likely to invest in the stock market than those children whose parents had less emotional attachment to the Telekom event.

Emotionally driven traders are not scarce in reality and stock market crashes are somewhat inevitable. Therefore, it is important to understand the underlying psychological channels that can provoke the emotional tagging of households and permanently close their doors to the stock market. Persistent non-participation can have severe negative consequences for households: according to a monthly report of the German Bundesbank, most German households have barely changed their savings and investment behavior since the emergence of the low interest rate environment, even though they are fully aware of the consequences of negative interest rates (Deutsche Bundesbank, 2015). This implies that German households - who mostly save their money in form of deposits - are missing out on equity premia that is especially important for long-term wealth accumulation in times of long-enduring low interest rates. Although I do not argue that the Telekom event is the main driver underlying such behavior, my empirical analysis introduces an important channel demonstrating that emotional-tagged events can persistently affect investment decisions.

For instance, insights of the Telekom event suggests *betrayal of trust* as an important determinant for the outbreak of long-lasting negative emotions after the crash. Indeed, many studies confirm the importance of trust on stock market participation (Guiso et al., 2008; Georgarakos and Pasini, 2011; Pevzner et al., 2015). Therefore, policies that help maintain a healthy trust relationship between investors and companies can help avoid such emotionally attached events .

4.A Sentiment analysis

In this appendix, I present how I conduct a sentiment analysis using newspaper articles from the *Handelsblatt*. The data is provided by the Forecasting and Economic Policy Department of DIW Berlin. All articles are converted in XML-format. For my purpose, I utilize articles from 1998 through 2002.

1. Import all articles (XML-Document) using Python. The XML format includes tags that indicate the topic of a certain article. Use category "Firms" and extract all articles that report news about Deutsche Telekom within this category. This includes all sort of information of the company, starting from sport sponsorship to IPOs. Afterwards, manually filter out articles that exclusively include information on Telekom shares. See Table 4.7.

2. Tokenize and lemmatize all texts using Natural Language Toolkits in Python (SpaCY).

3. Apply lexicon-based sentiment analysis using *SentiWS* (Sentiment WortSchatz), which is a publicly available German-language resource provided by Remus et al. (2010). This provides a list of positive and negative sentiment-bearing words that are weighted within the interval of $[-1, 1]$, which is used to evaluate the average tone of an article.

4. Consistency check: Manual screening of all relevant articles and categorize between (i) negative, (ii) neutral, and (iii) positive. Compare my evaluation with others and with the results produced with *SentiWS*.[24]

5. Calculate the yearly average of all Telekom shares articles that are available within a year. Afterwards, compute the growth rate of the yearly sentiment measures.

Table 4.7: Newspaper articles on Deutsche Telekom

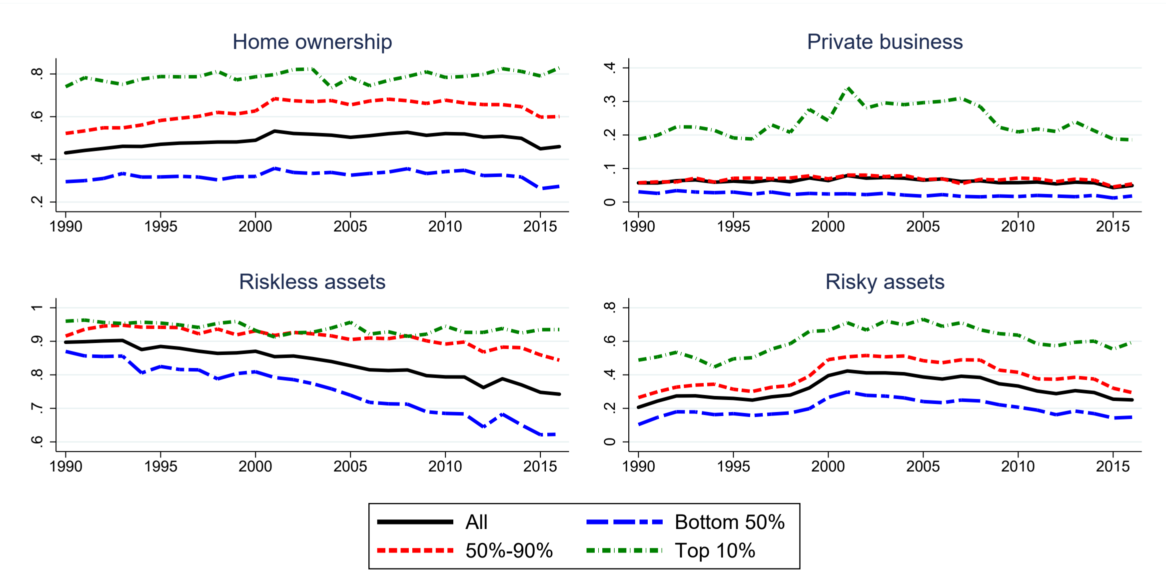
Year	# Deutsche Telekom	# Telekom shares
1998	242	44
1999	286	77
2000	322	120
2001	299	115
2002	269	124

[24]I thank Till Baldenius for his excellent support.

4.B Asset ownership of German households: 1990-2016

How do German households invest and which asset type is particularly popular? Did investment decisions change over the last 26 years? I provide answers to these questions using by utilizing information from the SOEP. To be specific, I provide descriptive statistics of German households' ownership of four different asset categories: (i) riskless assets, which include banking saving deposits, mortgage saving plans ("Bausparvertrag"), life insurance policies, and saving bonds, (ii) risky assets (security papers of listed companies including stocks, bonds, and equity warrants held directly or through mutual funds), (iii) home ownership, and (iv) private businesses. In addition, I divide the sample into three income groups (bottom 50%, 50%-90%, and top 10%). Figure 4.8 presents the time series from 1990 to 2016.[25] The black solid line represents the German average. On average, 49% of German households have reported to own a main dwelling, 6% have a business, 84% have invested in riskless financial assets, and 32% own risky assets.

Figure 4.8: Ownership of assets across different income groups



Source: Own calculations using SOEP data

When it comes to non-financial assets (main dwelling and private businesses), ownership rates seem to have barely changed between 1990 and 2016. The only noticeable change observed is in business ownership of households in the top 10%; however, it is more of a temporary increase in business holdings during 1996-2007 that reverts after the Global Financial Crisis period, instead of a persistent change.

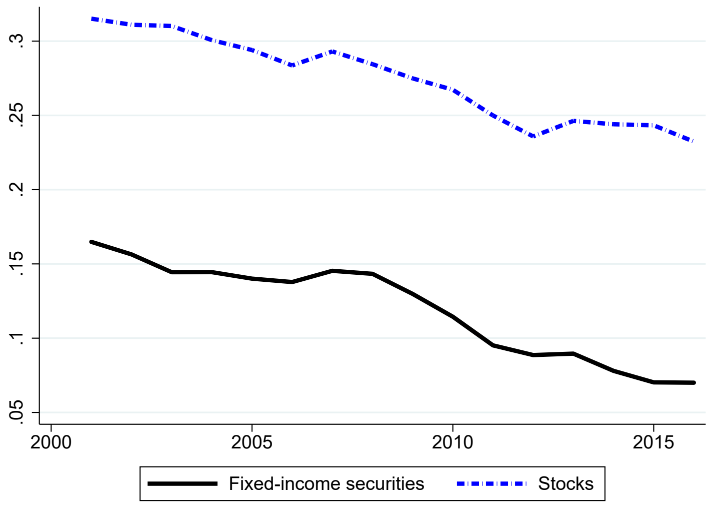
A large fraction of German households own riskless financial assets, with more than 90% of the top 50% of the income distribution claiming ownership of these assets. While these numbers also remain quite stable over time, I observe a slightly declining

[25]Strictly speaking, the survey waves are 1991 to 2017. Households report their asset ownership of the *previous* year.

trend in riskless asset ownership among the bottom 50% income distribution. While 87% of them owned riskless assets in 1990, only 63% did so in 2016. In case of risky financial asset holdings, the share increases between 1995 and 2000 for all income groups. Afterwards, ownership declines - again for all income groups - in a gradual manner.

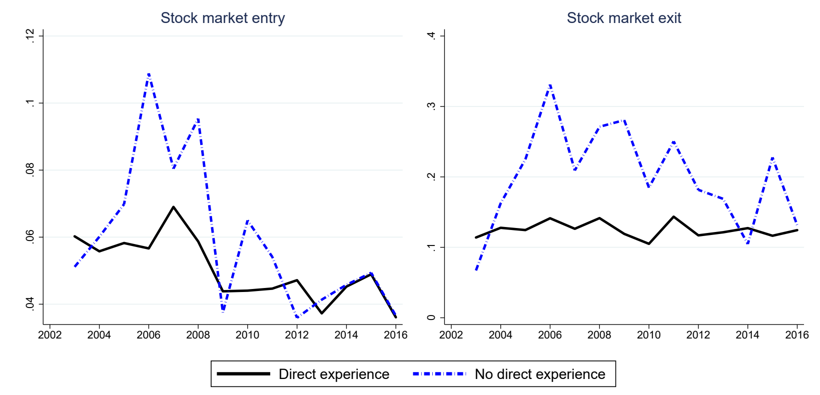
4.C Figures

Figure 4.9: Fixed securities ownership vs. stock ownership



Source: Own calculations using SOEP data

Figure 4.10: Stock market entry and exit



Source: Own calculations using SOEP data

Figure 4.11: DAX performance index

Source: Datastream

4.D Tables

Table 4.8: Identification of TE

	Hypothesis	Identification	Control group
Direct experience	The Telekom event led to less stock market participation among German households	Household heads who were younger than 20 in 2000 were old enough to directly experience the Telekom event	Household heads who were younger than 20 in 2000
Emotional experience	Households that have higher emotional attachment to the Telekom event invest less in stocks	Households that have invested in Telekom shares are likely to have high emotional attachment to the event, as they experienced real losses to their wealth (*Telekom investors*)	Households that experienced the Telekom event, but did not invest in T-shares
Indirect experience	Traumatic stock market experiences can even have an impact on investment decision of households that did not directly experience it.	Children of Telekom investors may have noticed the emotional turbulence of their parents and anticipate this experience when they make their own investment decisions	Children of experienced investors, i.e. parents who have high participation rate in the stock market throughout the survey waves

Table 4.9: The effect of direct experience: different λ

	SMP (1)	Entry (2)	Exit (3)
$\lambda = 0$			
Direct experience (β_1)	-0.52	-0.90^{***}	-0.86^{***}
Stock returns (β_2)	-1.94	-0.12	-6.36
Observations	4466	2577	1889
Pseudo R^2	0.19	0.12	0.16
$\lambda = -1$			
Direct experience (β_1)	-0.42^{***}	-0.92^{***}	-0.88^{***}
Stock returns (β_2)	-0.19	-0.00	-0.55
Observations	4466	2577	1889
Pseudo R^2	0.19	0.12	0.17
Controls	yes	yes	yes
Constant	yes	yes	yes
Region FE	yes	yes	yes

Notes: *, **, and *** represent statistical significance at the 10%, 5%, and 1% levels, respectively.

Table 4.10: The effect of direct experience: different age threshold equal 30

	SMP (1)	Entry (2)	Exit (3)
Direct experience (β_1)	-0.24	-0.98^{***}	-0.88^{***}
Stock returns (β_2)	-3.11	38.46	32.46
Controls	yes	yes	yes
Constant	yes	yes	yes
Region FE	yes	yes	yes
Observations	4466	2577	1889
Pseudo R^2	0.19	0.10	0.15

Notes: *, **, and *** represent statistical significance at the 10%, 5%, and 1% levels, respectively.

Table 4.11: Summary statistics

	German average		Experienced investors		Telekom investors	
	mean	sd	mean	sd	mean	sd
Demographics						
Sex	0.63	0.48	0.66	0.47	0.67	0.47
Age	46	13.33	47	13.02	45	12.39
Married	0.57	0.50	0.62	0.48	0.55	0.50
Number of children	0.54	0.91	0.46	0.84	0.47	0.83
Household size	2.41	1.27	2.39	1.19	2.34	1.19
High school degree (*Abitur*)	0.17	0.38	0.33	0.47	0.22	0.41
University degree	0.18	0.39	0.37	0.48	0.23	0.42
Employment	0.63	0.48	0.72	0.45	0.74	0.44
Former FRG/GDR	0.80	0.40	0.84	0.37	0.77	0.42
Total family income	34791.66	20697.43	47277.93	26572.32	37746.20	19961.38
Financial decisions						
Riskl. asset ownership	0.87	0.33	0.97	0.16	0.98	0.15
Stock ownership (proxy)	0.26	0.44	0.81	0.39	1.00	0.00
Home ownership	0.41	0.49	0.56	0.50	0.47	0.50
Asset income	1393.97	8107.94	4293.37	17811.90	1378.77	4015.32
Fin. risk taking	2.35	1.99	3.15	2.09	2.66	1.97
Observations	5730		882		1261	

Notes: Experienced investors are household heads with over 50% security papers ownership during their participation in the survey. Household heads who have reported to own security papers in 1996, 1999, and 2000 (but not before) are defined as Telekom investors. For the German average and experienced investors, I present the summary statistics of 1995, before the beginning of the Telekom event. Summary statistics of the Telekom investors refer to their characteristics in the year of their stock market entry. *Source*: Own calculations based on SOEP data.

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Eidesstattliche Erklärung

Hiermit erkläre ich, dass ich die vorgelegte Dissertation auf Grundlage der angegebenen Quellen und Hilfsmittel selbstständig verfasst habe. Alle Textstellen, die wörtlich oder sinngemäß aus veröffentlichten oder nicht veröffentlichten Schriften entnommen sind, sind als solche kenntlich gemacht. Die vorgelegte Dissertation hat weder in der gleichen noch einer anderen Fassung bzw. Überarbeitung einer anderen Fakultät, einem Prüfungsausschuss oder einem Fachvertreter an einer anderen Hochschule zum Promotionsverfahren vorgelegen.

Chi Hyun Kim
Berlin, March 2021

Liste verwendeter Hilfsmittel

- Statistische Programme

  Matlab, Python, R, Stata
- Datenbanken

  Datastream, Bloomberg
- Texteditoren

  LaTeX, Microsoft Excel, Microsoft Word
- Siehe auch Literatur- und Quellenangaben