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Caterina Forti Grazzini

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Dekan:	Prof. Dr. Dieter Nautz
Erstgutachter:	Prof. Dr. Dieter Nautz
	Professur für Ökonometrie
	Freie Universität Berlin
Zweitgutachter:	Prof. Dr. Alexander Kriwoluzky
	Professur für empirische Makroökonomie, Fiskal- & Geldpolitik
	Freie Universität Berlin und DIW Berlin
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To Hedi.

This thesis is for both of us

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Declaration of Co-Authorship and Publications

This dissertation consists of three research papers. Two papers were written in collaboration with one co-author. My contribution in conception, implementation and drafting can be summarized as follows:

- Caterina Forti Grazzini und Malte Rieth:

  "Interest Rates and Exchange Rates in Normal and Crisis Times"

  Contribution: 50 percent

- Caterina Forti Grazzini:

  "Unconventional Monetary Policy and Households' Financial Portfolio Choices"

  Contribution: 100 percent

- Caterina Forti Grazzini und Chi Hyun Kim:

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

  Contribution: 50 percent

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

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List of Abbreviations

ABS	Asset-Backed Securities
APP	Asset Purchase Programme
ARCH	Autoregressive Conditional Heteroscedasticity
BOT	Buoni Ordinari del Tesoro.
BTT	Buoni del Tesoro Poliennali
CPI	Consumer Price Index
CCT	Certificati di Credito del Tesoro
CTZ	Certificati del Tesoro Zero Coupon
EA	Euro Area
EC	European Crisis
ECB	European Central Bank
e.g.	exempli gratia
EUR	Euro
FE	Fixed Effects
FED	Federal Reserve
FEVD	Forecast Error Variance Decomposition
FOMC	Federal Open Market Commitee
FX	Foreign Exchange Rate
GARCH	Generalized Autoregressive Conditional Heteroscedasticity
GIPS	Greece, Ireland, Portugal, and Spain
HH	Household
i.e.	id est
LSAP	Large Scale Asset Purchases
MHH	Male-Headed Household
MOM	Month on Month
MP	Monetary Policy
IRA	Individual Retirement Account
LR	Likelihood Ratio
LTRO	Long-Term Refinancing Operation

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	Securities Market Program
SVAR	Structural Vector Autoregression
TLTRO	Targeted Longer-Term Refinancing Operation
UIP	Uncovered Interest Rate Parity
UK	United Kingdom
UMP	Unconventional Monetary Policy
US	United States of America
USD	US Dollar
VAR	Vector Autoregression
VIX	CBOE Volatility Index
YoY	Year on Year
ZLB	Zero Lower Bound

Summary

This thesis provides empirical evidence on the impact of the global financial crisis (2007-2009), the European sovereign debt crisis (2009-2012) and central banks's actions on different economic agents in the euro area and in the United States. It is composed by three independent chapters. The first chapter studies the impact of the above-mentioned two periods of crisis on the relation between euro area and US interest rates and the US dollar/euro exchange rate. The last two chapters focus on monetary policy and its impact on households' financial investment choices.

The first chapter is joint work with Malte Rieth. We use daily data from 2000 to 2016 to evaluate whether the structural relation between euro area and US interest rates and the US dollar/euro exchange rate changes across normal and crisis times. Using a structural VAR and exploiting the heteroskedasticity in the data for identification, we find that, on the one hand, positive US interest rate shocks appreciate the dollar in normal times, have no effect during the global financial crisis, and lead to a depreciation of the dollar during the European sovereign debt crisis; on the other hand, the impact of positive euro area interest rate shocks on the exchange rate surges during the European crisis, but we find no evidence of time-varying effect (in terms of sign) across crisis and non-crisis times. These patterns are consistent with the presence of a flight to safety channel towards the US during the two crisis samples that alters the traditional interest rate channel effect on exchange rates. Specifically, the results suggest that in times of crisis the US dollar is considered a safe haven currency.

In the second chapter I use 2006-2016 survey data on Italian households' financial portfolios to examine how the unconventional monetary policies implemented by the European Central Bank after 2007 affect households' asset allocation choices. I focus on two asset categories, Italian government bonds and Italian risky assets (equity, corporate bonds and mutual funds). First, I disentangle any change in households' investment in both categories into its active saving component (rebalancing) and its passive saving component (capital gains) using financial indexes. Then, I estimate the impact of unconventional measures on the active saving component. The empirical analysis finds that accomodative unconventional monetary policy exerts a substantial effect on the size and composition of households' financial investment. In particular, it induces a positive rebalancing into both Italian government bonds and risky assets, although this is true only for households at the top of the income distribution. Thus, the results show that, consistent with the confidence channel of unconventional monetary policy, European Central Bank's unconventional tools have contributed to reviving Italian households' appetite for some of the financial segments mostly hit by the crisis.

The third chapter is joint work with Chi Hyun Kim. We use US household survey data from 2001-2017 to investigate whether monetary policy has heterogeneous effects on the financial portfolio decisions of different household groups identified with a combination or their head's gender and marital status. On the one hand, we show that monetary policy affects stock market entry decisions of single female-headed households, while we find no impact for single and married male-headed households. On the other hand, we do not find any monetary policy's heterogeneous effect on household groups' exit decision nor in their stock market investment rebalancing choices. These results suggest that 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.

Keywords: International Transmission, Safe Haven, Crisis, United States, Euro Area, Monetary Policy, Unconventional Monetary Policy, Federal Reserve, European Central Bank, Households' Portfolio Choices, Gender.

JEL Classification: D14, E58, F31; G11, G51,J16.

Zusammenfassung

Im Rahmen der vorliegenden Dissertation habe ich empirisch untersucht, wie sich die globale Finanzkrise (2007-2009), die Europäische Schuldenkrise (2009-2012) und die von den Zentralbanken ergriffenen Maßnahmen auf verschiedene Wirtschaftsteilnehmer im Euroraum und in den Vereinigten Staaten ausgewirkt haben. Die Arbeit gliedert sich in drei voneinander unabhängige Kapitel. Im ersten Kapitel wird analysiert, welche Effekte die beiden aufeinanderfolgenden Krisen auf die Beziehung zwischen den Zinssätzen (in den USA und im Euroraum) und dem US-Dollar/Euro-Wechselkurs gehabt haben. In den beiden letzten Kapiteln liegt der Fokus auf der und Geldpolitik und darauf, wie sich diese auf die Investitionsentscheidungen privater Haushalte ausgewirkt hat.

Der Schwerpunkt des ersten Kapitels, welches in Zusammenarbeit mit Malte Rieth entstanden ist, liegt auf dem strukturellen Zusammenhang zwischen den Zinssätzen in den USA und im Euroraum sowie dem US-Dollar/Euro-Wechselkurs in “normalen” Zeiten und Krisenzeiten. Wir verwenden tägliche Daten von 2000 bis 2016. Durch Anwendung eines strukturellen vektorautoregressiven Modells (VAR-Modell) unter Ausnutzung der vorhandenen Heteroskedastizität in den Daten stellen wir fest, dass positive Zinsschocks in den USA in “normalen” Zeiten zu einer Aufwertung des Dollars führen, ein derartiger Effekt während der globalen Finanzkrise jedoch nicht zu beobachten war und es während der Europäischen Schuldenkrise zu einer Abwertung des Dollars kam. Zinsschocks im Euroraum haben während der Schuldenkrise einen vergleichsweise größeren Einfluss auf den Wechselkurs gehabt. Das Verhalten der Wechselkurse in Krisenzeiten steht im Einklang damit, dass während der Krisen ein “flight to safety channel” für Investoren Richtung USA vorhanden war, wodurch die standardmäßige Effekte des Zinskanals als Transmissionsmechanismus auf die Wechselkurse modifiziert wurde.

Im zweiten Kapitel wird basierend auf Umfragedaten über die Finanzportfolios italienischer Haushalte, die sich auf den Zeitraum von 2006 bis 2016 beziehen, untersucht, inwiefern sich die unkonventionellen geldpolitischen Maßnahmen der Europäischen Zentralbank (EZB) auf die Entscheidungen der Haushalte hinsichtlich der Aufteilung ihres

Vermögens auf verschiedene Anlagen ausgewirkt haben. Aus dabei auftretenden Vermögensänderungen wurden mithilfe von Finanzindizes die Effekte isoliert, die auf aktives Sparen (Portfolioumschichtungen) zurückgeführt werden können.

Anschließend wurden anhand von zwei ausgewählten Anlageoptionen - italienische Staatsanleihen und riskante Aktiva (Aktien, Unternehmensanleihen und offene Investmentfonds) - ermittelt, welchen Einfluss die unkonventionellen Maßnahmen auf Portfolioumschichtungen hatten. Im Rahmen der empirischen Untersuchung hat sich herausgestellt, dass die Maßnahmen der EZB einen erheblichen Einfluss auf den Umfang und die Zusammensetzung der Finanzportfolios von privaten Haushalten hatten, einschließlich eines prozyklischen Anstiegs der Investitionen in sowohl Staatsanleihen als auch in risikobehaftete Aktiva - jedoch nur bei den Haushalten am oberen Ende der Einkommensverteilung. Die Ergebnisse veranschaulichen die Wirkung der unkonventionellen geldpolitischen Maßnahmen der EZB. Die Maßnahmen haben seit 2007 zur Wiederbelebung Nachfrage in einigen Segmenten des Finanzmarktes, die am stärksten von der Krise betroffen waren, beigetragen.

Das dritte Kapitel ist in Zusammenarbeit mit Chi Hyun Kim entstanden. Wir haben Umfragedaten aus den USA für die Jahre 2001-2017 ausgewertet, um zu untersuchen, ob die Geldpolitik eine unterschiedliche Wirkung auf die Investitionsentscheidungen der verschiedenen Haushaltsgruppen gehabt hat. Zu diesem Zweck wurden die einzelnen Haushalte anhand des Geschlechts sowie des Familienstands des Haushaltsvorstands zu Gruppen zusammengefasst. Auf der einen Seite konnten wir so zeigen, dass die Geldpolitik die Entscheidung für den Einstieg in den Aktienmarkt im Falle von Single-Haushalten mit einem weiblichen Haushaltsvorstand beeinflusst, während geldpolitische Schocks die Entscheidungen von Single-Haushalten mit einem männlichen Haushaltsvorstand und die von Haushalten mit einem verheirateten männlichen Haushaltsvorstand diesbezüglich nicht beeinflussen. Auf der anderen Seite konnten wir für die geldpolitischen Schocks keine heterogenen Effekte in Bezug auf die Entscheidung der Haushaltsgruppen für den Ausstieg aus dem Aktienmarkt oder deren Entscheidungen zur Umschichtung ihrer Portfolios nachweisen. Diese Ergebnisse deuten darauf hin, dass Single-Haushalte mit einem weiblichen Haushaltsvorstand sensibler auf geldpolitische Zyklen reagieren als Haushalte mit einem männlichen Haushaltsvorstand, aber nur, wenn sie nicht schon an der Börse aktiv sind.

Schlagworte: Internationale Transmision, Safe Haven, Krisen, USA, Eurozone, Geldpolitik, Unkonventionelle Geldpolitik, Federal Reserve, Europäische Zentralbank, Portfolioentscheidungen von Haushalte, Geschlecht.

JEL Classification: D14, E58, F31; G11, G51,J16.

Introduction and Overview

In the first two decades of the 21st Century, the euro area and the United States have witnessed periods of strong financial and economic instability that forced policy makers and researchers to rethink macroeconomic policies, financial supervision and as well as central banks' role and objectives. First, the global financial crisis (2007-2009) showed how little was known about the dangers posed by the financial system and about the consequences of periods of crisis on financial markets. Then, the European sovereign debt crisis (2009-2012) revealed the weaknesses of the Euro area currency union and of its fiscal policy. To some extent, the real and financial consequences of these events are still unfolding as I write.

Exceptional times called for exceptional measures. The European Central Bank (ECB) and the Federal Reserve (as well as many other central banks around the world) have reacted to these challenges aggressively, cutting monetary policy rates and implementing the so called "unconventional monetary policies" (UMPs). In particular, the main objectives of these unconventional tools were to address problems in the monetary policy transmission mechanism and to provide additional monetary stimulus once standard policy rates could not be lowered further (Potter and Smets, 2019).

Motivated by these events, the three independent chapters of this thesis provide empirical evidence on the impact of periods of crisis and central banks' actions on different economic agents, with a focus on the euro area and the United States. The first chapter evaluates the impact of periods of crisis on financial markets and studies how the structural relation between interest rates and the US dollar/euro exchange rate changes across tranquil times, the global financial crisis and the European sovereign debt crisis. The last two chapters take a different perspective and focus on monetary policy and its impact on households' financial investment choices. Chapter 2 evaluates how the unconventional monetary policy implemented by the European Central Bank affects the financial portfolio choices of Italian households. Chapter 3 analyzes whether gender and marital status affect households' stock market investment decisions in response to monetary policy shocks.

From the methodological point of view, this thesis applies different econometric techniques, depending on the research focus and the type of data used for the analysis. Chapter 1 employs a structural VAR identified through the heteroskedasticity present in the data to reach identification (Rigobon, 2003; Lanne and Lütkepohl, 2008). Chapters 2 and 3 use cross-sectional data and panel data econometric techniques applied to microdata obtained from household-level surveys.

The **first chapter**, titled *Interest Rates and Exchange Rates in Normal and Crisis Time*, is a joint work with Malte Rieth. The global financial crisis and the subsequent European sovereign debt crisis (or European crisis) had substantial effects on global exchange rate configurations (Fratzscher, 2009; Ehrmann et al., 2014). In particular, the US-dollar/euro exchange rate decreased sharply after the fall of Lehman Brothers, with hitherto unseen increases in volatility that persisted also through the European crisis. Nonetheless, the literature that explains the determinants of the USD/EUR exchange rate behavior during these two crisis episodes is limited. Fratzscher (2009) and Kohler (2010) explain the 2008-2009 dollar appreciation with switch in the relationship between the currency and US macroeconomic announcements and with the safe haven status of the dollar, respectively. The studies that focus on the European crisis find contrasting results (Ehrmann et al., 2014; Stracca, 2015). Moreover, no study considers nor compares the behavior of the USD/EUR exchange rate in both crisis episodes.

Chapter 1 attemps to shed new light on exchange rate behaviour in times of crisis by focussing on how the structural relationship between the USD/EUR exchange rate, the euro area and the US interest rates changes across tranquil times, the global financial crisis and the European crisis. We employ a structural VAR model and we include the US dollar/euro exchange rate, the 2-year German Bund rate (as a proxy for the euro area interest rate), and the 2-year US Treasury rate. We collect data from January 2000 through October 2016 and we split the sample into four sub-periods: pre-crisis, global financial crisis, European crisis, and post-crisis. A main difficulty in identifying the causal effects between interest rates and exchange rates is the endogeneity of these asset prices even at the daily frequency. We circumvent this issue by using the heteroskedasticity in the data to identify the contemporaneous impacts of asset price changes (Rigobon, 2003; Lanne and Lütkepohl, 2008).

Our findings contribute to the literature by showing that the relation between interest rates and exchange rates changes during crisis episodes. In the two non-crisis samples we find that the US dollar/euro exchange rate responds in a manner consistent with the traditional interest rate channel, which is decreasing (increasing) in response to a positive shock to the US (euro area) interest rate. On the contrary, in the two crisis samples we find a striking asymmetry between the effects of the US and euro area interest rates. While during the global financial crisis both interest rate shocks cease to have an impact on the USD/EUR FX, during the European crisis both areas' shocks become significant, with the impact of the US interest rate shock having opposite sign with respect to the non-crisis samples. The documented patterns are consistent with a flight to safety channel towards the dollar in time of distress that alters the traditional interest rate channel of exchange rate. Specifically, the results suggest that during crises the US dollar is considered a safe haven currency.

The **second chapter**, titled *Unconventional Monetary Policy and Households' Financial Portfolio Choices* deals with the impact of monetary policy and households' financial choices. Unconventional monetary policy is expected to affect investors' portfolio choices, for example through the *portfolio rebalancing channel* or the *confidence channel* (Krishnamurthy and Vissing-Jorgensen, 2011; Krishnamurthy et al., 2017; Fratzscher et al., 2018). Yet, in the current academic debate on the impact of unconventional tools on households' wealth redistribution it is always assumed that households do not adjust their portfolios in response to monetary policy (Adam and Tzamourani, 2016; Casiraghi et al., 2018; Ampudia et al., 2018; Lenza and Slacalek, 2018). This chapter analyzes the impact of unconventional monetary policy on how much households decide to rebalance their portfolio in the context of a specific case: the impact of the European Central Bank's unconventional tools on the portfolio rebalancing of Italian households.

I combine the 2006-2016 waves of the Bank of Italy Survey on Household income and Wealth with financial indexes. This allows me to isolate the active saving component of households' investment in Italian government bonds and Italian risky assets (equity, corporate bonds and mutual funds). Then, I estimate the impact of unconventional monetary policy on these two active saving components. Analysing households' portfolio choices is important not only because it gives new insights into the effectiveness and the transmission channels of unconventional measures, but especially because it sheds new light on an under-explored channel through which unconventional tools can affect wealth redistribution: the one that stems from households' heterogeneous response to monetary easing.

Results show that in a period of financial turbulence in which households have drastically reduced their financial investment, accommodative unconventional tools boost the investment in both Italian government bonds and Italian risky assets. These findings are consistent with the so-called confidence channel of monetary policy and point toward the effectiveness of central banks in restoring Italian households' trust in the financial system. The study also finds that these results only hold for households in the top quartile of the income distribution. Moreover, I conduct a simulation exercise that evaluates the impact of unconventional tools on households' financial wealth considering both UMP-induced valuation and rebalancing effects. The simulation shows that the differences in household portfolio performance across the income distribution has to be attributed almost solely to the rebalancing component. This suggests that rebalancing decisions play a prominent role in the effects of monetary policy on households' wealth redistribution.

The **third chapter**, titled *The Effect of Monetary Policy on Stock Market Investment Decisions: The Role of Gender and Marital Status*, is a joint work with Chi Huyn Kim. One key aspect of the current debate on possible redistributional effects of central banks' actions (Yellen, 2016; Draghi, 2016) is to understand what are households' financial and demographic characteristics that interact with monetary policy. Several papers document monetary policy's heterogeneous effects along income, wealth, house ownership, employment status and age (see, among the others, Bivens, 2015; Ampudia et al., 2018; Wong, 2019). In this chapter we take a different perspective and we evaluate the impact of gender and marital status on monetary policy-driven financial portfolio decisions, focusing in particular on single female-headed households.

In the financial literature the link among gender, marital status, risk preferences and investment decisions is well established. The empirical findings show that men invest significantly more in financial assets than women. Moreover, women are more risk averse (Jianakoplos and Bernasek, 1998; Sunden and Surette, 1998; Fisher and Yao, 2017) and less confident in their investment decisions(Barber and Odean, 2001; Croson and Gneezy, 2009). In a prolonged period of low interest rates - and, thus, high asset prices - the single female-headed households' low propensity to bear financial risks and invest in financial assets may lead to a distributional divergence between different household groups.

Thus, we investigate whether monetary policy has a different impact on the stock market investment decisions of single-female headed households and both single and married male-headed households. We use survey data from the Panel Study of Income Dynamics (PSID) between 2001 and 2017. Employing different econometric techniques, we show that on the one hand, contractionary monetary policy negatively affects single female-headed households' stock market status, decreasing their probability of stock market entry (while this is not the case for male headed-households). On the other hand, we do not find any monetary policy's heterogeneous effect on household groups' exit decision or in their stock market investment rebalancing choices. 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 the single female-headed households are more sensitive to monetary policy cycles than male headed-households in their investment decisions, but only if they are not already participating in the stock market.

This thesis is an attempt to shed light over some recent economic events, which left behind many open questions in the fields of both macroeconomics and finance. I see these chapters as the beginning of a personal research agenda, as well as a modest attempt to better understand the consequences of the most prominent events unfolding in the past twenty years.

CHAPTER 1

Interest Rates and Exchange Rates in Normal and Crisis Times[1]

1.1 Introduction

A long-standing puzzle in international macroeconomics and finance is the difficulty to explain and predict exchange rate fluctuations (Rossi, 2006). In their seminal work, Meese and Rogoff (1983) show that a simple random walk beats various exchange rate models in terms of forecasting performance. A large number of subsequent papers introduced nonlinearities to improve the explanatory power of these models and to understand the determination of exchange rates. One strand focuses on reduced form forecasting models (Rossi, 2006, 2013). Several other studies center on changes in the structural relations between macroeconomic fundamentals and exchange rates (Fratzscher, 2009; Bacchetta et al., 2009; Mumtaz and Sunder-Plassmann, 2013). Finally, a different strand focuses on the relationship between interest rates (and especially their differential) and exchange rate (for example, Flood and Rose, 2002; Cheung et al., 2005; Engel, 2016), as the former are considered foundamental predictors of the latter (Rossi, 2013; Engel, 2014). In this Chapter, we empirically revisit the relationship between interest rates and exchange rate, focusing our attention on whether it changes across normal and crisis times.

The choice to take a closer look to the possible changes in the interest rates/exchange rate relationship across different periods is motivated by the crisis episodes originating in the US (the 2007-2009 global financial crisis) and the in euro area (the 2009-2012 European sovereign debt crisis, or European crisis), as well as by the integration of global

[1]This chapter is joint work with Malte Rieth. We thank Menzie Chinn, Michael Ehrmann, Marcel Fratzscher, Peter Karadi, Angela Maddaloni, Lukas Menkhoff, Andreas Schrimpf. We are particularly grateful to Kenza Benhima and all participants to the WinE retreat, Lisbon, to the EEA annual conference, Lisbon, to the Jahreestagung of the Verein für Sozialpolitik, Vienna, as well as to participants of the DIW PhD internal seminar for helpful comments and suggestions.

financial markets. The latter has increased substantially since the turn of the century, with US markets, investors, and monetary policy playing a main role in the determination of international asset prices (Rey, 2015), including the exchange rates. During the global financial crisis, investors repatriated funds to the US (Fratzscher, 2009; McCauley and McGuire, 2009) and they fled the euro during the European debt crisis (Ehrmann et al., 2014). Reflecting these developments, the US-dollar/euro exchange rate decreased sharply after the fall of Lehman Brothers, with hitherto unseen increases in volatility in the following two periods of crisis (see Figure 1.1). Nonetheless, the literature explaining the determinants of the USD/EUR exchange rate behavior during these two crisis episodes is limited. Few works explain the 2008-2009 dollar appreciation with switch in the relationship between the currency and US macroeconomic announcements (Fratzscher, 2009), or with the safe haven status of the dollar (Kohler, 2010). Few other papers stress the role of US dollar shortages generated by the funding of net long US dollar exposure by European banks (McCauley and McGuire, 2009; Hui et al., 2011). The few studies that focus on the European crisis find contrasting results. On the one hand, Ehrmann et al. (2014) show that neither macroeconomic fundamentals, nor policy actions or the public debate by policy makers are able to explain euro exchange rate volatility between 2009 and 2012. On the other hand, Stracca (2015) finds that European crisis did have a consistent and depreciationary effect on the euro exchange rate. Moreover, no study encompasses nor compares these crisis episodes when analysing exchange rate developments.

Against this background, this Chapter attemps to shed new light on exchange rate behaviour in times of crisis by focussing on how the structural relationship between the USD/EUR exchange rate, the euro area and the US interest rates changes across tranquil times, the global financial crisis and the European crisis. In fact, understanding the determinants of the nominal exchange rate, especially in time of economic distress, is crucial from both the financial and the real point of view. On the one hand, the US dollar/euro exchange rate is the biggest and most important of all foreign exchange (FX) markets, with daily turnover of almost \$1.3 trillion and with a market share of 24% of all FX markets (2013). On the other hand, it is important for policymakers and practitioners alike since this price determines foreign demand for home goods if prices are sticky.

The empirical model includes three asset prices at the two-day frequency - the 2-year interest rate on the German Bunds (as a proxy for the euro area interest rate), the 2-year interest rate on the US Treasury and the US dollar/euro exchange rate - over the years 2000 to 2016.[2] We split the sample into four subperiods, the pre-crisis, global financial crisis, European crisis and post-Outright Monetary Transaction (OMT). Within each subsample, we model each asset with a multifactor model and then we use the heteroscedasticity in the data to identify the contemporaneous impacts of asset price changes (Sentana and Fiorentini, 2001; Rigobon, 2003; Lanne and Lütkepohl, 2008). In a nutshell, this identification strategy exploits the fact that periods of higher interest rate volatility contain additional information on the response of the exchange rate to interest rates changes. These shifts in the volatility of interest rates can then be used as a “probabilistic instrument" (Rigobon, 2003) to trace out the impact of interest rate shocks on the exchange rate. Similarly, the framework allows quantifying the strength of contemporaneous interest rate spillovers across the Atlantic.

The empirical analysis finds significant changes across subsamples in the structural relations among the endogenous variables. In the pre-crisis sample exchange rate movements are dominated by the US interest rate shocks, while in the post-crisis sample both euro area and US interest rates are of similar importance for the determination of the exchange rate. Moreover, in both non-crisis periods (pre-crisis and post-OMT) we find that the US dollar/euro exchange rate responds in a manner consistent with the traditional interest rate channel, which is decreasing (increasing) in response to a positive shock to the US (euro area) interest rate. On the contrary, in the two crisis samples we find a striking asymmetry between the effects of US and euro area interest rate shocks on the exchange rate. While during the global financial crisis both US and euro area shocks cease to have an impact on the USD/EUR FX, during the European crisis both areas' shocks become significant. What is interesting, though, is that in the second crisis subsample, the impact of the US interest rate flips sign with respect to the non-crisis samples, and a positive US interest rate shock has a *negative* impact the exchange rate. On the contrary, the euro area shock effects are in line with the non-crisis periods (although significantly bigger in magnitude). For the bidirectional spillovers between interest rates, we find that they are stronger from the US to the euro area than vice versa across all subsamples but that they do not change in magnitude or sign over time. These results are qualitatively and quantitatively robust to various alterations of the model and the data. Specifically, we assess the sensitivity of the estimates to changing the definition of the volatility regimes used for the identification, including a market sentiment variable, considering additional interest rate maturities and merging the two crisis periods in one.

[2]This Chapter was produced at an earlier stage of the dissertation. For this reason the analysed data sample ends at the end of 2016.

The findings are supported by the forecast error variance decomposition. US interest rate shocks explain a larger fraction of exchange rate and interest rates variability in the pre-crisis, lose their role during the global financial crisis and they gain it back during the European crisis. Euro area interest rate shock contribution to explaining exchange rate variation peaks during the European crisis. Finally, the regression of the US/EUR exchange rate on the euro area and US macroeconomic news confirms this pattern. In particular, in the post-OMT subsample, a positive news shock - i.e. a better than expected performance of the US economy - leads to an appreciation of the dollar and, in turn, to a decrease of the exchange rate. On the contrary, during the two crises there is evidence of a switch in sign, implying that some positive news about the US economy is perceived even better for the euro area, leading to an appreciation of the euro and pushing up the exchange rate.

Our findings show that the relation between interest rates and exchange rates changes during crisis episodes and provide a potential explanation for why uncovered interest parity may not hold on average (Engel, 2014). The documented patterns are consistent with a flight to safety channel towards the dollar in time of distress that alters the traditional interest rate channel of exchange rate. Specifically, the results suggest that during the crises the US dollar is considered a safe haven currency. While exogenous decrease in US interest rates typically lead to a depreciation of the dollar, during crisis times, i.e. when financial distress is high, investors might perceive them as a positive signal about the state of the US economy and increase their demand for what they perceive being a safer currencies. According to our results, the safe haven effect seems to offset the traditional interest rate channel for US interest rate shocks in the global financial crisis and dominate (and revert) it during the European debt crisis. The Chapter also shows that the ECB's announcement of the OMT program seems to have re-established normality in the foreign exchange market. After Summer 2012 the impact of both US and euro area interest rates shocks are consistent with the traditional interest rate channel and of similar importantance.

Our paper contributes to several strands of literature. Generally, it connects to empirical studies of the relationship between exchange rates and interest rates. Ehrmann et al. (2011), using a vector autoregressive model estimated on pre-crisis data, find that both US and euro area interest rate shocks are important for USD/EUR exchange rate developments. A long-standing literature analyses this relationship within the framework of the uncovered interest rate parity (UIP), which states that interest rates are fundamental in determining the exchange rate. UIP is a cornerstone of international finance, constituting an important building block of many exchange rate determination theories. However, the empirical evidence supporting UIP is mixed.

While several papers find evidence favoring this relation (Baillie and Bollerslev, 2000; Chinn and Meredith, 2004), others document its failure (Bekaert et al., 1997; Engel, 1996; Bekaert et al., 2007).

Another, more closely, related literature analyses whether the behavior of exchange rates changes during particular time periods. Focusing on the global financial crisis, Fratzscher (2009) shows that this episode triggered sharp and unexpected currency movements, with domestic macroeconomic fundamentals and the financial exposure of individual countries playing a key role in the transmission of US shocks; Kohler (2010) finds that safe haven flows and the role played by interest rate differentials contribute to explain why a large number of countries that were not at the centre of the crisis depreciated against the US dollar, the Japanese yen and the Swiss franc. Concentrating on the European crisis, Ehrmann et al. (2014) present evidence that the euro mainly danced to its own tune, with fundamentals and policy decisions possessing only a little explanatory power with respect to exchange rate fluctuations. Swanson and Williams (2014) investigate the effect of the zero lower bound on asset price formation using an event study design. They show that interest rates are partially constrained during this period, while exchange rates are not. Similarly, Stavrakeva and Tang (2016) show that the contemporaneous relation between yields at different maturities and exchange rates changed after the zero lower bound was hit in the US, based on a model with VAR-based expectations of short-run yields and inflation. Several other papers analyze whether there are differences between the effects of conventional and unconventional monetary policy shocks on exchange rates (Glick and Leduc, 2013; Kiley, 2013; Glick and Leduc, 2015; Ferrari et al., 2016). The main difference between these papers and our work is that we do not analyze the relationship between exchange rate and interest rates conditional on a particular shock, but more generally we study the reaction of the exchange rate to interest rate shocks.

The paper is organized as follows. In Section 1.2, we derive the main hypothesis for the analysis. In Section 1.3, we outline the empirical model and describe the data. Section 1.4 outlines the empirical findings. Section 1.5 contains the robustness checks, before the last section concludes.

1.2 Conceptual framework

This Chapter analyses the unconditional relationship between exchange rate and interest rates. In this section, we derive our main hypothesis for the empirical analysis. We first discuss the contemporaneous links between interest rates and exchange rates in normal times. Then, we outline potential effects of periods of crisis on this relationship.

Among financial market operators, folk wisdom establishes that a country's currency appreciates as its interest rates increase relative to those of another country (Stavrakeva and Tang, 2016). This basic and contemporaneous relationship derives from the uncovered interest rate parity (UIP) condition. The UIP provides a simple relationship between the interest rate on an asset denominated in a domestic country's currency unit, the interest rate on similar asset denominated in a foreign country's currency, and the expected rate of change in the spot exchange rate between the two currencies, conditional to the risk-neutrality and the market efficiency hypotheses. The UIP predicts that an increase in the domestic to foreign interest rate differential is, *ceteris paribus*, associated with an appreciation of the domestic currency. In the context of this Chapter, i.e., considering the US and the euro area, this can be written as

$$E_t(e_{t+h}^{US,EA}) - e_t^{US,EA} = \alpha + \beta(r_t^{US} - r_t^{EA})$$

where $e_t^{US,EA}$ is the natural logarithm of the spot exchange rate of US-dollar per unit of euro, r_t^{US} is the US nominal interest rate, r_t^{EA} is the euro area nominal interest rate proxied by the German interest rate, and h is the horizon. The UIP predicts that high interest rate currencies will depreciate relative to low interest rate currencies.

If the UIP condition holds, then $\alpha = 0$ and $\beta = 1$ and the UIP condition can be rewritten as

$$E_t(e_{t+h}^{US,EA}) - e_t^{US,EA} = r_t^{US} - r_t^{EA}. \tag{1.1}$$

After taking first differences, Equation (1.1) becomes

$$E_t(\Delta e_{t+h}^{US,EA}) - \Delta e_t^{US,EA} = \Delta r_t^{US} - \Delta r_t^{EA}. \tag{1.2}$$

For sufficiently large horizons, $E_t(\Delta e_{t+h}^{US,EA})$ should be very small because, conditional on t, when h increases the difference between $E_t(e_{t+h})$ and $E_t(e_{t+h-1})$ decreases (the value of the two forecasts should be very close); hence, we assume it is close to zero (Kiley, 2013).

Equation (1.2) can be written as

$$\Delta e_t^{US,EA} = -\Delta r_t^{US} + \Delta r_t^{EA}. \tag{1.3}$$

Equation (1.3) gives a (consistent with theory) contemporaneous relationship between changes in interest rates and exchange rate. It implies that an increase in Δr_t^{US} (Δr_t^{EA}) yields to an appreciation of the dollar (euro), that in turn leads to a decrease (increase) in the US dollar/euro exchange rate.

The following equation is the relationship we build our empirical framework on

$$\Delta e_t^{US,EA} = \alpha - \beta_1 \Delta r_t^{US} + \beta_2 \Delta r_t^{EA} + \epsilon_t. \tag{1.4}$$

It is important to note that, although derived from the UIP condition, our model relaxes the assumption of common coefficients between the two interest rates. In fact, separating the coefficients allows us to understand how the relative contribution to the exchange rate formation of the two interest rates evolves over time. Thus, from Equation (1.4), in normal times we expect a positive shock to the US (euro area) interest rate to appreciate the dollar (the euro) and, in turn, to decrease (increase) the exchange rate.

Results may change in times of financial distress. A relatively well-established literature finds that returns on low-interest rate currencies tend to be negatively correlated with global risk aversion, while high- yield currencies often crash exactly when global risk aversion is high (Brunnermeier and Nagel, 2008). This implies that, in times of crisis, large deviations from the UIP condition may occur, with low yield currencies appreciating during times of global financial stress (see, among others, Flood and Rose, 2002; Habib and Stracca, 2012). This may be due to an increase in market participants' risk aversion or to the failure of the rational expectation hypothesis. In particular, if market operators become risk averse, the UIP may not hold because agents require rates of return that are higher than the interest differential for holding a country asset (Fama, 1984; Taylor, 1995).

The failure of the UIP can lead to the safe haven phenomenon: a currency is considered a safe haven if, in periods of high uncertainty and risk aversion, it appears to be more attractive than others. In particular, the definition of safe haven implies that the relative price of such an asset should increase during crises. Moreover, in contrast to a hedge currency, which is expected to appreciate on average, a safe haven currency is expected to appreciate against others only in times of uncertainty (Beckmann and Czudaj, 2017). This definition of safe haven currency is in line with the dollar's behavior between at least 2008 and 2009. In fact, as noted by several observers (Fratzscher 2009; McCauley and McGuire 2009; Beckmann and Czudaj 2017; Reinhart and Reinhart 2008; Kohler 2010), one paradoxical aspect of the global financial crisis was the appreciation of the US currency and its behavior as safe haven, despite the fact that the financial crisis had been generated by, and was propagating from, the US to the rest of the world. What is less clear is if the dollar's role as safe haven can be extended also vis-à-vis the euro and if it holds not only during the global financial crisis but also throughout the subsequent European crisis.

The safe haven phenomenon operates in the direction opposite to the UIP one. If we consider the US dollar as a safe haven, this would imply that a positive shock to the US interest rate would lead to an *depreciation* of the dollar, leading to an increase in the exchange rate. Considering again Equation (1.4), this would imply a sign reversion of the β_1 coefficient, leading to the following equation

$$\Delta e_t^{US,EA} = \alpha + \beta_1 \Delta r_t^{US} + \beta_2 \Delta r_t^{EA} + \epsilon_t. \tag{1.5}$$

From Equation (1.4) and (1.5), it follows that in turbulent times both the standard (UIP-consistent) effect and the safe haven effect can be in place and that it is *ex-ante* unclear which one will prevail, as they could offset each other. Thus, in times of crisis, we expect an increase of the euro area interest rate to lead to an appreciation of the euro (and, thus, to an increase of exchange rate), while it is unclear what is the effect of a positive shock to the US interest rate on the dollar and, in turn, its final effect on the exchange rate.

1.3 Empirical analysis

This section presents the empirical model, the data and the estimation methodology.

1.3.1 Empirical model and data

The structural VAR model is

$$Ay_t = \tilde{c} + \tilde{A}_1 y_{t-1} + ... + \tilde{A}_p y_{t-p} + \tilde{\Gamma} x_t + \epsilon_t, \tag{1.6}$$

where y_t is the vector of endogenous variables, which includes: $e_t^{US,EA}$, the natural logarithm of the nominal exchange rate measured as the amount of US dollars per one euro (such that an increase in the variable reflects a depreciation of the US dollar vis-à-vis the euro); r_t^{EA}, the risk-free rate in the euro area, approximated through the 2-year rate on German government bond; r_t^{US}, the 2-year rate on the US Treasury bonds. All three variables are in first differences to account for the non-stationarity of data,

$$y_t = \begin{pmatrix} \Delta e_t^{US,EA} \\ \Delta r_t^{EA} \\ \Delta r_t^{US} \end{pmatrix};$$

x_t is a vector of exogenous variables; $\tilde{A}_t$ and $\tilde{\Gamma}$ with $i = 0, ..., p$ are coefficient matrices that capture the lagged and contemporaneous effects of the endogenous and exogenous variables, respectively; $\tilde{c}$ is a vector of constants; finally, ϵ_t is a vector of structural shocks with diagonal variance matrix $\Sigma_\epsilon = E(\epsilon_t \epsilon_t')$. Table 1.7 in the Appendix contains the summary statistics of the endogenous variables.

The vector of exogenous variables x_t includes the news or surprise component of macroeconomic announcements for both the US and the euro area about real, financial, and confidence variables. These surprises are computed as the actual realization of the economic indicators minus the financial market's expectations from few days before. The data are obtained from Bloomberg. Several papers show that macroeconomic surprises are important for the development of exchange rates and interest rates (among the others, Andersen et al. 2003 and Swanson and Williams 2014). For this reason, it is crucial to take them into account to rule out the possibility that asset price shocks are driven by common shocks, as this would imply the non-orthogonality of the structural shocks. Table 1.8 in the Appendix provides the complete list of included macroeconomic news, together with their summary statistics. We also include dummies for the day of the week.

Data is collected at daily frequency from Datastream and Bloomberg for the January 1, 2000, to October 31, 2016, period. We split the data into four subsamples: a pre-crisis period, the global financial crisis, the European sovereign debt crisis, and the period post-OMT (see Figure 1.1). The first subsample runs from the beginning of the sample until August 9, 2007, when the largest French bank BNP Paribas temporarily halted redemptions from three of its funds that were holding assets backed by US subprime mortgage debt. This event is seen by many commentators as the trigger of the global financial crisis (Trichet, 2010; Cecchetti, 2008). The second subsample follows the dating of Ehrmann and Fratzscher (2017) and it covers the period from August 10, 2007, until September 1, 2009, when the newly elected Greek government announced for the first time that there could be issues with the Greek government debt and deficit data. Then the European sovereign crisis subsample starts, continuing until September 30, 2012, which is after ECB President Draghi's "Whatever it takes" speech on July 26, 2012, and after the ECB's official announcement and implementation of the OMT program (August and September 2012). Finally, the post-OMT period runs from October 1, 2012, until the end of the sample. As we formally show in the main analysis, the data support this sample split, as the structural relations among the endogenous variables change significantly across subsamples. Throughout the Chapter, the analysis and the results are conducted and reported separately for the four subperiods.

We construct 2-day windows following Ehrmann et al. (2011). In fact, US shocks may occur after the closing of European markets, thus affecting the latter only on the next business day. To avoid losing information on macroeconomic surprises, all news that would happen on an excluded day are moved forward and reported as if they would occur on the next business day. Both endogenous and exogenous variables are standardized prior to estimation, as this enable us to compare the relative effects of variables having different units of measurement. The standardization is subsample specific.

1.3.2 Identification through heteroscedasticity

For estimation we pre-multiply the structural VAR model (1.6) by A^{-1}

$$y_t = A^{-1}\tilde{c} + A^{-1}\tilde{A}_1 y_{t-1} + ... + A^{-1}\tilde{A}_p y_{t-p} + A^{-1}\tilde{\Gamma} x_t + A^{-1}\epsilon_t$$

and re-write it as

$$y_t = c + A_1 y_{t-1} + ... + A_p y_{t-1} + \Gamma x_t + u_t, \tag{1.7}$$

where the vector of reduced form shocks $Au_t = \epsilon_t$ is related to the structural shocks through matrix A. The matrix A is the focus of the Chapter because it contains the contemporaneous effects of structural shocks on the endogenous variables (i.e., without taking into account all instantaneous feedbacks among variables). In particular, the off-diagonal elements of matrix A are of main interest for us, as they indicate the effect on impact of the interest rates on the exchange rate. After estimation, we are able to assess the effect of a positive shock to the euro area and US interest rates on the USD/EUR exchange rate in crisis and non-crisis times by testing whether, within each subsample, $\alpha_{1,2}$ and $\alpha_{1,3} \lesseqgtr 0$, where $\alpha_{j,k}$ is the respective element in the estimated A. Moreover, we are also able to test and compare estimates across subsamples.

The parameters of equation (1.7) and the covariance matrix of the reduced form, Σ_u, can be consistently estimated by ordinary least squares. In particular, based on the Bayesian information criterion, in the analysis of each subsample we include one lag of the endogenous variables, but in the post-OMT one, where we include two, so to obtain residuals free from autocorrelation. However, in order to recover the structural parameters, we need to identify the impact matrix A. From Equation (1.6) and (1.7), it follows that $\Sigma_u = A^{-1}\Sigma_\epsilon(A^{-1})'$. In this system, the number of unknowns is larger than the number of independent equations and it is necessary to add additional information to achieve the identification of the impact matrix A. A common practice in the structural VAR literature is to impose zero restrictions. However, as exchange rates and government interest rates react simultaneously to each other, it is difficult to defend the imposition of zero restrictions. One alternative option would be employing sign restrictions, which allows for contemporaneous effects among the variables. The side effect of this approach is that it constrains the response of the endogenous variables to be of an *ex-ante* determined sign. However, as there is no consensus in the literature on the bi-directional causality between asset price relationships, it is difficult to develop a robust sign scheme that would allow the identification of shocks. For example, Fratzscher (2009) finds that, in period of crisis, negative news about the US economy would make the dollar appreciate while Rogers et al. (2018) document that during the European crisis, the ECB's monetary policy easing would lead to a euro appreciation. Both findings are in contrast to what standard theory would predict. Moreover, the main hypothesis of this Chapter is that during crisis episodes, the relationship between exchange rates and interest rates may flip sign due to, as explained in Section 1.2, the safe haven phenomena. Thus, it is precisely the aim of this analysis to determine the signs of the relationships between endogenous variables and we thus do not, of course, impose them *a priori*.

To identify the shocks, we use the identification through heteroscedasticity (IH), as developed by Sentana and Fiorentini (2001) and Rigobon (2003). This methodology exploits the fact that financial variables are generally found to be heteroscedastic (Ehrmann et al., 2011) to obtain information on the response of variables to each other.[3] The idea behind it is that changes in the relative variance of the structural shock over time allows for determining the entries of A. A short illustration explains how different volatility regimes contain additional information that can be exploited. Defining k as the number of volatility regimes and considering our 3-variable system, when $k = 1$ we have only six moments on the LHS that can be estimated but nine parameters on the RHS that need to be determined (three structural shock variances and six off-diagonal elements in A, with the main diagonal normalized to unity). When $k \geq 2$, the system has as many moments to be estimated as unknowns and it can be solved. Thus, we assume that, within each subsample, the structural shocks have the following regime-dependent diagonal variance matrix in regime k:

$$\Sigma_{\epsilon,k} = E(\epsilon_t \epsilon_t') = \begin{pmatrix} \sigma_k^1 & 0 & 0 \\ 0 & \sigma_k^2 & 0 \\ 0 & 0 & \sigma_k^3 \end{pmatrix},$$

and that $\Sigma_{u,k} = A^{-1}\Sigma_{\epsilon,k}(A^{-1})'$.

The identification strategy relies on three assumptions: first, the structural shocks have to be uncorrelated; second, the ratio of the shock variances changes significantly across regimes; and, third, the contemporaneous impact matrix A must be stable across regimes, in order to ensure that each additional regime adds more equations than un-

[3]The specific form of heteroscedasticity is not of interest, as Rigobon (2003) shows that the estimates of the contemporaneous relationships are consistent regardless of the form of heteroscedasticity.

knowns. The first assumption is standard in the structural VAR literature. Moreover, to increase the likelihood that this assumption holds, we control for common effects through macroeconomic news and other exogenous variable (see Section 2.4). The second assumption can be formally tested after estimation. The third assumption is standard in both classical linear applications (no volatility regimes) and (G)ARCH models. Moreover, the choice of conducting the analysis separately on the four subsamples substantially increases the likelihood of this assumption to hold. Finally, the system is identified only up to a rotation of the matrix A and, therefore, we need to impose some additional restrictions to ensure that we pick the rotation that represents the true underlying economic relationships among variables.[4] Following Ehrmann et al. (2011), we impose a sign restriction on the structural coefficients and we assume that an increase in the 2-year US Treasury interest rate leads to an increase in the 2-year euro interest rate. This restriction is fairly uncontroversial and it helps us to correctly identify the structural parameters without restricting the coefficients of main interest. Moreover, the constraint is never binding in the estimation.

1.3.3 Identification of volatility regimes and estimation

Before the estimation, we need to define the volatility regimes. We apply a statistical approach, where the determination of the regimes is data-driven. The following procedure is applied separately to each of the four subsamples. We start by estimating the reduced form model (Equation 1.7). Then, we compute the rolling standard deviation for each reduced form residual u_t. We proceed by calibrating the threshold for the rolling standard deviation that defines whether the residual are classified into a high or low volatility. The threshold, as well as the window used to compute the rolling standard deviation, are subsample-specific and they are reported in the Appendix, Table 1.9. In particular, for each subsample we identify five volatility regimes and we define them as follows: 1) All residuals are classified in low volatility; 2) Only one residual is in high volatility, the others are in low; 3) All residuals display high volatility. Finally, observations that cannot be classified into any of the five regimes are excluded from the analysis (for an overview of the number of observations included in all volatility regimes and subsamples, see Table 1.10 in the Appendix). Windows and thresholds are chosen so to minimize the number of observations not included in the analysis while

[4] A short example using demand and supply problem can easily explain this issue. Consider the following two equations: $D = aP + e_1$ (supply) and $P = bD + e_2$ (demand). In this context there is another system (symmetric to the truthful one) that produce the same variance-covariance matrix. However, this implies that, empirically, it is not possible to disentangle the two systems from the variance-covariance matrix unless we impose a sign restriction that allows for identifying the correct rotation. Thus, this sign restriction is not used to identify the matrix, but only to make sure that a solution that is economically meaningful and consistent with theory is picked.

ensuring the minimum number of observations per regime (which, following Ehrmann et al. (2011), is set to 16 observations).

To see whether our regimes are supported by the data, we formally test for the constancy of the reduced form covariance matrix within each subsample. Following Lanne and Lütkepohl (2008), we perform a test on the joint null hypothesis that all five regimes have the same covariance matrix (Table 1.1). Moreover, we conduct pairwise likelihood ratio tests on the null hypothesis so that any two regimes have the same variance-covariance matrix (Table 1.11 in the Appendix). All null hypotheses are strongly rejected, implying that we can reject the equality of the variance-covariance matrices. In all subsamples, the data prefer a model with changes in volatility over the assumption of homoscedasticity.

Table 1.1: Likelihood ratio tests for constancy of the reduced form covariance matrix

	Pre-crisis	Global financial crisis	European crisis	Post-OMT
LR statistic	909.62	375.20	357.72	739.66
p-value	(0.00)	(0.00)	(0.00)	(0.00)

Note: The table shows results for all subsamples of a likelihood ratio test on the null hypothesis that, within a subsample, all regimes have the same reduced form covariance matrix. Bootstrap p-values are in parentheses.

After having determined the volatility regimes, within each subsample we estimate the model as in Ehrmann et al. (2011) by minimizing the following matrix norm:

$$\begin{aligned} &\min \quad g'g \quad \text{with} \quad g = \sum_{k=1}^{5} (A\Sigma_{u,k}A' - \Sigma_{\epsilon,k}), \\ &\text{s.t.} \quad \Sigma_{u,k} \quad \text{diagonal and } A \quad \text{sign restriction} \end{aligned} \tag{1.8}$$

where $\Sigma_{u,k}$ is the variance of the reduced form shocks in regime k, $\Sigma_{\epsilon,k}$ is the variance of the structural shocks in regime k, and A is the matrix of contemporaneous impact subject to the sign restriction defined in Subsection 1.3.2. Statistical inference is based on 200 bootstrapping replications. In each replication we use the regime-specific covariance matrix to generate new data from which we obtain estimates using the minimization procedure. We compute the p-values as the share of estimates beyond zero.

As mentioned above, one identification assumption is that all possible estimated variance-ratios of the uncorrelated structural shocks are sufficiently distinct across regimes (Lütkepohl and Netšunajev, 2014). To check if this condition is met, after the estimation, we consider all variance ratios $\phi_k^{S,S'} = \sigma_k^S/\sigma_k^{S'}$, for each pair of shock (S, S'), each regime k and each subsample. Then, for each shock pair, we use a Wald test on the joint null hypothesis that the variance ratio is the same across regimes, i.e., that $\phi_1^{S,S'} = \phi_2^{S,S'} = \phi_3^{S,S'} = \phi_4^{S,S'} = \phi_5^{S,S'}$, which would invalidate the identification of A. Inference on these tests is based on 200 bootstrap replications. Table 1.2 shows that, for each pair of shocks and each subsample, the null hypothesis of no changes in volatility is rejected. The model is statistically identified.

Table 1.2: Identification test

H_0	$\phi_1^{e,EA} = \phi_2^{e,EA} = \phi_3^{e,EA}$ $= \phi_4^{e,EA} = \phi_5^{e,EA}$	$\phi_1^{e,US} = \phi_2^{e,US} = \phi_3^{e,US}$ $= \phi_4^{e,US} = \phi_5^{e,US}$	$\phi_1^{EA,US} = \phi_2^{EA,US} = \phi_3^{EA,US}$ $= \phi_4^{EA,US} = \phi_5^{EA,US}$
		Pre-crisis	
Wald stat	27.13	21.12	16.10
p-value	(0.00)	(0.00)	(0.00)
		Global financial crisis	
Wald stat	22.01	17.75	18.67
p-value	(0.00)	(0.00)	(0.00)
		European crisis	
Wald stat	7.85	9.36	24.42
p-value	(0.10)	(0.05)	(0.00)
		Post-OMT	
Wald stat	31.87	21.38	42.42
p-value	(0.00)	(0.00)	(0.00)

Note: The table shows the Wald statistics and the associated p-values of linear Wald tests on the joint null hypothesis that, within subsample, the estimated variance ratios of two structural shocks, $\phi_k^{s,s'} = \sigma_k^s/\sigma_k^{s'}$, are the same across volatility regimes, for each pair of structural shocks. σ_k^s is the estimated variance of shocks s = USD/EUR exchange rate, 2-year German interest rate, 2-year US Treasury rate in regime $k = 1, ..., 5$. The tests are based on 200 boostrap replications. P-values are in parentheses.

1.4 Results

We now present the empirical results. We first discuss the estimates of the direct effects, focusing specifically on the impact of the interest rates on the exchange rate;

then, we move to the variance decomposition. Finally, we analyze the sensitivity of the exchange rate to macroeconomic news.

1.4.1 Direct effects

Table 1.3 presents the estimated A-matrix for all four subsamples. These coefficients are the direct effects of a one standard deviation shock (in columns) on the endogenous variables (in rows). All other variables are kept constant within period and there is no contemporaneous feedback among endogenous variables. For ease of interpretation, we reverse the signs of the off-diagonal elements. Statistical significance at the 1%, 5%, and 10% level is denoted by *a*, *b*, and *c*, respectively.

The following set of equations presents the estimates of contemporaneous response of the exchange rate to interest rate shocks for each subsample. These correspond to the first row of each of the four panels in Table 1.3. We highlight parameters that are at least significant at the 10% level using bold font:

$$\textit{Pre-crisis}: e_t^{US,EA} = -0.02 r_t^{EA} - \mathbf{0.22} r_t^{US} + ... \tag{1.9}$$

$$\textit{Glob fin crisis}: e_t^{US,EA} = +0.21 r_t^{EA} - 0.01 r_t^{US} + ... \tag{1.10}$$

$$\textit{Euro crisis}: e_t^{US,EA} = +\mathbf{0.29} r_t^{EA} + \mathbf{0.32} r_t^{US} + ... \tag{1.11}$$

$$\textit{Post-OMT}: e_t^{US,EA} = +\mathbf{0.21} r_t^{EA} - \mathbf{0.20} r_t^{US} + ... \tag{1.12}$$

Equation (1.9) shows that, in the pre-crisis period, a one standard deviation increase of the 2-year Treasury directly decreases the exchange rate by 0.22 standard deviation on impact. In contrast, the effect for the euro area is not significant. Thus, the evidence for the pre-crisis and the US supports the traditional interest rate channel on exchange rate, with a positive shock to the 2-year Treasury rate leading to a decrease of the USD/EUR exchange rate. With the unfolding of the global financial crisis the situation changes (Equation 1.10). Not only does the euro area interest rate remains insignificant, but now also the US interest rate ceases to affect the exchange rate. One possible explanation is that the two channels described in Section 1.2, the traditional interest rate channel and the safe haven channel offset each other so that the final outcome is insignificant. During the European crisis (Equation 1.11), the situation is reversed. The euro area gains importance and the exchange rate now responds to developments of the German rates, probably as a consequence of the fear of a possible euro break-up. Regarding the US interest rate, the results suggest that the safe haven channel prevails over the traditional one. In fact, the US coefficient turns positive, implying that a positive one standard deviation shock to the Treasury leads to a *depreciation* of the dollar and, in turn, to an increase of 0.32 standard deviation of the exchange rate. Finally, after the announcement of the ECB's OMT program (Equation 1.12) the situation reverts back to normality in terms of signs, with the euro area estimate being positive and significant and the US coefficient being negative. In terms of magnitude, the two areas are now of similarly importance for the exchange rate determination.

To formally compare results across subsamples, we conduct a pairwise t-test on the null hypothesis that two estimates are equal. The tests for the euro area and US coefficients are run separately. From Table 1.4, it emerges that all tested differences are significantly different from zero and that they support our hypothesis of a change in the relationship between interest rates and exchange rate change during crisis episodes. Moreover, the table also shows that the US interest rate impact on the exchange rate in the pre-crisis and post-OMT subsamples is similar both for sign and magnitude, but that their difference is statistically different from zero.

Summing up, the results show that the relationship between the USD/EUR exchange rates and the US/euro area interest rates changes over time and that, in tranquil times, it is consistent with the folk wisdom that a country's currency appreciates when its interest rates increase relative to that of another country. In particular, the analysis finds that while US economic conditions dominate exchange rate movements before the global financial crisis, the euro area variable gains importance in the post-crisis sample. The two crisis samples behave very differently and we find a striking asymmetry between the effects of the two interest rate shocks on the exchange rate. While, during the global financial crisis, the US shocks cease to have an impact on the currency, interest rate shocks from the euro area increase their impact during the European crisis. The results of the two crisis subsamples are in line with Stavrakeva and Tang (2016)'s finding that after the first quarter of 2009 (i.e., during the zero lower bound period, which, for us, starts at the end of the global financial crisis period, continues throughout the European crisis, and partially overlaps the post-OMT subsample), the increase in longer bonds (one year or above) leads to a depreciation of the dollar. On the contrary, our crisis-period findings are in contrast with Glick and Leduc (2015), which find that both accommodative conventional and unconventional monetary policy shocks (measured as the change in long term Treasury rate futures in a tight window around a monetary policy announcement) in the ZLB period lead to a depreciation of the dollar. Nonetheless, two points are worth noting. First, this Chapter considers the USD/EUR exchange rate (in contrast to only the dollar). Second, we study the *unconditional* relationship between exchange rate and interest rates, which we do not condition with respect to any particular shock (as the monetary policy one). Finally,

Table 1.3: Contemporaneous effects among endogenous variables - main results

Response	Impulse: USD/EUR exchange rate	Impulse: 2-year euro area rate	Impulse: 2-year US interest rate
		Pre-crisis	
USD/EUR ex rate	1.00	-0.02	-0.22
p-value	.	.	.b
2-year euro rate	-0.13	1.00	0.37
p-value	.	.	.a
2-year US rate	-0.09	0.27	1.00
p-value	.	.a	.
		Global financial crisis	
USD/EUR ex rate	1.00	0.21	- 0.01
p-value	.	.	.
2-year euro rate	0.16	1.00	0.45
p-value	.	.	.a
2-year US rate	-0.15	0.33	1.00
p-value	.	.b	.
		European crisis	
USD/EUR ex rate	1.00	0.29	0.32
p-value	.	.b	.b
2-year euro rate	0.22	1.00	0.34
p-value	.b	.	.a
2-year US rate	-0.41	0.24	1.00
p-value	.a	.b	.
		Post-OMT	
USD/EUR ex rate	1.00	0.21	-0.20
p-value	.	.a	.b
2-year euro rate	0.19	1.00	0.33
p-value	.	.	.a
2-year US rate	0.08	0.23	1.00
p-value	.	.a	.

Note: The table shows, for each of the four subsamples, the estimated direct effects of a one standard deviation structural shocks on the endogenous variables, based on a three-variable SVAR identified through heteroscedasticity. The impulse variables are in columns, response variables are in rows. For ease of interpretation, the signs of the off-diagonal elements are reversed. The subsample periods are as follows. Pre-crisis: 3 Jan 2000 - 8 Aug 2007; global financial crisis: 9 Aug 2007 - 30 Sep 2009; European crisis: 1 Oct 2009 - 30 Sep 2012; post-OMT: 1 Oct 2012 - 31 Oct 2016. .a, .b, .c below point estimates denote significance at the 1%, 5%, 10% level, respectively.

Table 1.4: Test for the equality of A-matrix coefficients

H_a	$\beta^{GFC} - \beta^{pre} > 0$	$\beta^{EC} - \beta^{GFC} > 0$	$\beta^{post} - \beta^{EC} > 0$	$\beta^{post} - \beta^{pre} \neq 0$
Euro area				
Difference between βs	0.24	0.10	-0.09	0.25
T stat	12.67	5.68	- 8.20	17.06
p-value	(0.00)	(0.00)	(0.00)	(0.00)
United States				
Difference between βs	0.22	0.29	-0.54	-0.02
T stat	12.86	14.20	-34.67	-2.21
p-value	(0.00)	(0.00)	(0.00)	(0.03)

Note: The table shows results for the t tests on the null hypothesis that two coefficients are equal. The test statistic is:

$$t = \frac{\beta_1 - \beta_2}{\sqrt{s_1^2/n_1 + s_2^2/n_2}}$$

where β_1 and β_2 are the two means, s_1 and s_2 are unbiased estimators of the variances of the two samples and n_1 and n_2 are the number of observations. In our case, $n_1 = n_2$. In the first row we report the alternative hypothesis we test for. Then, we report the actual difference between the two coefficients. Moreover, we report the value of the t statistic. For ease of interpretation, the signs of the differences and of the t statistics are reversed. P-values are in parentheses.

the announcement of the OMT program seems to have re-established normality in the foreign exchange market, because, in the last subsample, signs revert back to what the traditional UIP-consistent interest rate channel would predict. This supports once more the safe haven behavior of the dollar vis-à-vis the euro (in contrast, for example, to its possible use as a hedge currency), because it implies that the dollar appreciates with respect to the euro only in times of crisis and not also in tranquil periods (Beckmann and Czudaj, 2017).

The following sets of equations present the results of contemporaneous responses of the 2-year euro area rates and of the 2-year US Treasury rates. They correspond to the second rows (Equation 1.13 to 1.16) and the third row (Equation 1.17 to 1.20) of the four panels in Table 1.3. The euro area rate responds to the exchange rate and to its US counterpart in the following way:

$$\textit{Pre-crisis}: r_t^{EA} = -0.13 e_t^{US,EA} + \mathbf{0.39} r_t^{US} + ... \tag{1.13}$$

$$\textit{Glob fin crisis}: r_t^{EA} = -0.16 e_t^{US,EA} + \mathbf{0.45} r_t^{US} + ... \tag{1.14}$$

$$\textit{Euro crisis}: r_t^{EA} = \mathbf{0.22} e_t^{US,EA} + \mathbf{0.34} r_t^{US} + .. \tag{1.15}$$

$$\textit{Post-OMT}: r_t^{EA} = +0.19 e_t^{US,EA} + \mathbf{0.33} r_t^{US} + ..., \tag{1.16}$$

while the US Treasury rate responds to euro rate and exchange rate as follows:

$$\textit{Pre-crisis}: r_t^{US} = -0.09 e_t^{US,EA} + \mathbf{0.27} r_t^{EA} + ... \tag{1.17}$$

$$\textit{Glob fin crisis}: r_t^{usUS} = 0.15 e_t^{US,EA} + \mathbf{0.33} r_t^{EA} + ... \tag{1.18}$$

$$\textit{Euro crisis}: r_t^{US} = \mathbf{-0.41} e_t^{US,EA} + \mathbf{0.24} r_t^{EA} + ... \tag{1.19}$$

$$\textit{Post-OMT}: r_t^{US} = 0.08 e_t^{US,EA} + \mathbf{0.23} r_t^{EA} + ... \tag{1.20}$$

The spillovers from the United States to the euro area are generally larger than in the other direction. This result is in line with the US as the leading market and consistent with , among others, findings in Chinn and Frankel (2003), Ehrmann and Fratzscher (2005), and Ehrmann et al. (2011), while holding across both tranquil and turbulent times. Moreover, although we constrain the parameter of the effect of US interest rate shocks on the euro interest rate to be positive, the restriction is not binding in any of the subperiods. Moreover, the relationship between the two interest rates does not change sign during period of turmoil. The reason is most likely found in the peculiar role of safe haven held by the German interest rate within the eurozone during crisis periods (see, among others, Von Hagen et al. 2011 and De Santis 2012). In fact, as both countries are perceived safe havens for the euro area, this implies that the relationship between them is invariant. Furthermore, the magnitude of the coefficient is fairly constant, slightly peaking only during the global financial crisis period. Finally, as to the effect of a shock to the exchange rate, the estimates are only significant during the European crisis subsample, with the effects on the euro area and the US interest rates of opposite sign.

1.4.2 Variance decomposition

To quantify the average economic significance of the different types of structural shocks and to understand which shock contributes the most in explaining the volatility of the endogenous variables, we compute the 1-step ahead forecast error variance decomposition (FEVD). The FEVD shows how much of the forecast error variance of each variable can be explained by exogenous shocks to the other variables. As we have five different regimes, we obtain five forecast decompositions for each subsample. Thus, we compute a weighted average of the regime specific decomposition using the number of observations per regime as weights. Only this last set of results are reported in Table 1.5.

Table 1.5: Weighted forecast error variance decomposition

	USD/EUR exchange rate	2-year euro area rate	2-year US interest rate
		Pre-crisis	
USD/EUR ex rate	0.95	0.00	0.05
2-year euro rate	0.04	0.81	0.15
2-year US rate	0.02	0.06	0.92
		Global financial crisis	
USD/EUR ex rate	0.94	0.05	0.01
2-year euro rate	0.01	0.82	0.16
2-year US rate	0.02	0.12	0.87
		European crisis	
USD/EUR ex rate	0.77	0.09	0.14
2-year euro rate	0.01	0.84	0.15
2-year US rate	0.16	0.01	0.82
		Post-OMT	
USD/EUR ex rate	0.93	0.05	0.02
2-year euro rate	0.04	0.87	0.09
2-year US rate	0.02	0.11	0.88

Note: The table shows the weighted forecast error variance decompositions over an horizon of 2 days for each subsample. For each subsample, the weighted FVED is computed averaging over the forecast error variance decompositions calculated for each regime, using the number of observations per regime as weights, based on a structural VAR identified through heteroscedasticity. For each subsample, we identify five volatility regimes and we define them as follows: 1) all residuals are classified in low volatility; 2) only one residual is in high volatility, the others are in low; 3) all residuals display high volatility. Observations that cannot be classified into any of the five regimes are excluded from the estimation. The windows and thresholds used to define the volatility regimes are subsample-specific (see Table 1.9 in the Appendix).

The weighted FEVD results are in line with previous findings, corroborating and mirroring the evidence for a time varying impact of interest rates on exchange rate across normal and crisis times. Table 1.5 shows that, as it is usually found, each variable variance is mainly explained by its own idiosyncratic shock (see, among the others, Ehrmann et al., 2011). The importance of the two interest rates in explaining the variation of the exchange rate evolves over time. The US accounts for a larger fraction before the crisis but it loses its dominance with the unfolding of the global financial crisis. During the European crisis, the importance of the US shocks spikes again, returning to very low levels in the aftermath of the OMT announcement. Further, the contribution of the euro area shocks follows closely the findings discussed in previous section: it is equal to zero in the pre-crisis subsample, it increases until its maximum (9%) in the European crisis, and then it reverts to a lower level in the last subsample. In general, the overall contributions of the two areas to the exchange variation is increasing over time, moving from 5% in the pre-crisis sample to 23% during the European crisis. In the post-OMT period, the explanatory power of both interest rates goes back to pre-crisis levels.

Turning to the interest rates, on average the share of their variance not explained by idiosyncratic shocks is larger than in the exchange rate case. The spillovers from the US to the euro interest rate are the strongest. Results also show that, over time, the two areas show a diverging trend: while a larger share of the behavior of the US interest rate is explained by the other two shocks, the opposite is true for the euro area rate. Moreover, while the contribution of the US interest rate is fairly constant over time (with the exception of the last subsample), the spillovers from the euro area increase up to 12% during the global financial crisis, then drop to almost zero during the European crisis, implying that the US rates are almost only determined domestically.

1.4.3 The sensitivity of exchange rate to macroeconomic news

So far, the analysis shows that the impact of interest rate shocks on the exchange rate is time varying, pointing toward the role of the US-dollar as a safe haven vis-à-vis the euro in times of crisis. In this section, we check whether the same result emerges when considering the impact of the macroeconomic surprises. As explained in Section 2.4, the SVAR includes US and euro area macroeconomic news shocks as exogenous variables. Here we present the effects of these macro surprises on the USD/EUR exchange rate. The estimates are obtained estimating the first row of the reduced form model (Equation 1.7) with robust standard errors, in order to account for the heteroscedasticity present in the data,

$$e_t = c + \alpha e_{t-1} + \beta r_{t-1}^{EA} + \gamma r_{t-1}^{US} + \delta X_t + u_{e,t}. \tag{1.21}$$

The vector X_t contains all included macroeconomic news (see Table 1.8 in the Appendix for the complete list).[5] The results are reported in Table 1.6. Each of the four columns reports the results of a separate regression of the USD/EUR exchange rate on the macroeconomic data surprises (listed in the first column) in the subsample specified in the first row of the table. All variables are standardized prior to estimation and the standardization is subsample-specific. We do not interpret all single coefficients, as we are mostly interested in the general trend. Moreover, we only show the estimates of the surprises being significant at least in one subsample. The significant coefficients are reported in bold font to facilitate reading. In line with the interpretation of the results in Section 1.4.1, a positive (negative) coefficient, i.e. a stronger than expected performance of European (American) economy, implies an appreciation of the euro (US dollar) and, in turn, an increase (decrease) of the exchange rate.[6] We also experiment with excluding the lag values of the exchange rate, the German bund and the Treasury, and excluding those days on which no major macroeconomic data was released (i.e., days on which X_t is identically zero). Results are very similar to those presented in Table 1.6.

The results show that, over time, the exchange rate sensitivity to macroeconomic surprises coming from the euro area is U-shaped and it reaches its minimum during the European crisis, while the responsiveness to US surprises increases over time. This implies that, on the on hand, our results are in line with Swanson and Williams (2014)'s findings and that the exchange rate response to the macroeconomic surprises was constrained neither by the US zero lower bound nor by the global financial crisis; on the other hand, we also find that, consistent with Ehrmann et al. (2014), the USD/EUR is not responsive to the news coming from the euro area during the European crisis. In general, the effects of the news coming to the euro area is mixed, as there are several cases where positive news tend to depreciate the euro. Regarding the effect of the surprises coming from the other side of the Atlantic, while positive news about the US economy tend to appreciate the dollar in the post-OMT, during the two crisis subsamples, there is evidence of a partial switch in signs. This implies that a better than expected performance of several US variables (though not all) leads to a depreciation of the US dollar. This suggests that, in line with previous results, positive news for the US economy may be perceived as even better for the euro area, inducing a euro strengthening. These results are consistent with the dollar being a safe haven currency vis-à-vis the euro.

[5] Contrary to the main analysis, here we use daily changes instead of 2-day changes data, in order to capture the effect of the news on the day of their release.

[6] This reasoning does not hold for the Eurostat Unemployment Rate news, as a positive surprise is generally perceived as negative for the euro area, depreciating the euro.

Table 1.6: Effects of selected surprises on the USD/EUR exchange rate - estimation of the reduced form model

	Pre-crisis	Global financial crisis	European crisis	Post-OMT
Eura area surprises				
EC Business Climate Ind	0.0026	0.0611	0.0204	**0.0602***
	(0.0199)	(0.0611)	(0.0300)	(0.0330)
Euro Consumer Price Index YoY	-0.0039	0.0026	**0.0704***	**-0.0639****
	(0.0246)	(0.0463)	(0.0402)	(0.0274)
Gross Fixed Capital Formation QoQ	**-0.0142***	-0.0503	0.0201	0.0061
	(0.0073)	(0.0370)	(0.0259)	(0.0277)
Retail Sales Volume MoM SA	-0.0008	**-0.119*****	-0.0479	0.0107
	(0.0390)	(0.0442)	(0.0523)	(0.0415)
Trade Balance with non Eurozone	**0.0315****	**-0.0636***	-0.0174	-0.0221
	(0.0132)	(0.0360)	(0.0311)	(0.0475)
PPI Eurozone Industry Ex Construction MoM	0.0384	**-0.214***	-0.0020	-0.0456
	(0.0252)	(0.126)	(0.0714)	(0.0450)
Eurostat Unemployment Rate	-0.0086	-0.0195	-0.0278	**-0.0893****
	(0.0284)	(0.0408)	(0.0318)	(0.0369)
US surprises				
CPI Urban Consumers YoY	-0.0201	0.0348	**-0.141*****	-0.0089
	(0.0329)	(0.1480)	(0.0489)	(0.0536)
CPI Urban Consumers no Food&Energy YoY	0.0240	0.0195	0.0494	**-0.0634***
	(0.0201)	(0.0464)	(0.0428)	(0.0366)
Personal Consumption Expenditure CPI YoY	0.0070	-0.0036	-0.0440	**-0.0764*****
	(0.0201)	(0.0336)	(0.0366)	(0.0280)
CPI Urban Consumers MoM SA	0.0153	-0.104	**0.0941****	0.0286
	(0.0322)	(0.1740)	(0.0440)	(0.0673)
S Government Budget Balance	-0.0051	0.0510	-0.0321	**-0.0441***
	(0.0116)	(0.0347)	(0.0284)	(0.0249)
US Consumer Spending Growth Rates MoM	**-0.0293***	0.0131	0.0243	**0.0536****
	(0.0165)	(0.0288)	(0.0357)	(0.0247)
Industrial Production MoM	-0.0133	0.0433	-0.0136	-0.0010
	(0.0288)	(0.0780)	(0.0463)	(0.0340)
Core Producer Price Index	0.0020	0.0010	**0.0546***	-0.0253
	(0.0208)	(0.0465)	(0.0320)	(0.0177)
Producer Price Index - Finished Goods	-0.0138	**0.2050***	0.0472	-0.0065
	(0.0288)	(0.1240)	(0.0664)	(0.0965)
GDP Chained 2009 Dollars QoQ	**0.0426*****	0.0097	-0.0366	**-0.0446***
	(0.0149)	(0.0305)	(0.0249)	(0.0256)
Capacity Util % of Total Capacity	-0.0208	-0.0285	**-0.0817****	-0.0166
	(0.0281)	(0.0643)	(0.0347)	(0.0301)
Business Inventories MoM	0.0098	**0.0728****	-0.0303	-0.0307
	(0.0192)	(0.0306)	(0.0230)	(0.0194)
Prod Output Per Hour Nonfarm Bus QoQ	**0.0366****	-0.0095	**0.0529***	-0.0233
	(0.0182)	(0.0360)	(0.0291)	(0.0472)
Nonfarm Payrolls Total MoM	-0.0447	0.0054	0.0495	**-0.144*****
	(0.0294)	(0.0473)	(0.0338)	(0.0253)
Observations	1979	560	782	1065
R^2	0.026	0.084	0.037	0.064
F	1.758	2.026	1.922	2.707
p-value	0.06	0.00	0.00	0.00

Note: The table shows the effects of statistically significant variables in each subsample on the USD/EUR exchange rate from the baseline VAR, obtained from estimating the first row of the reduced form model (Equation 1.7) with robust standard error, in order to account for heteroscedasticity: Only estimates of the exogenous regressors have been reported here. All surprises have been standardized. $.^{***}$, $.^{**}$, and $.^{*}$ denote significance at the 1%, 5%, and 10% levels, respectively.

The R^2 of the regressions is low, ranging from the 0.026 of the pre-crisis subsample to the 0.084 of the global financial crisis subsample, consistent with the standard finding that it is difficult to explain exchange rate movements with fundamentals (Andersen et al., 2003; Swanson and Williams, 2014). Nonetheless, the regressions have a very high degree of statistical significance overall, which peaks in the two crisis subsamples and in the post-OMT. This indicates that the USD/EUR does responds systematically to many macroeconomic announcements. The results support the findings described in Subsection 1.4.1.

1.5 Robustness checks

In what follows we perform a number of robustness analyses, which can be classified into four different groups. First, we introduce in our analysis the Vix, in order to control for investor sentiment and market volatility. Second, we test whether our findings are robust to using different maturities of US and German interest rates. Third, we analyze the robustness of the main results to a different calibration and definition of the volatility regimes. Finally, we specify a different crisis sample. Given the large number of estimates, in the discussion we mainly focus only on the contemporaneous impact of the two interest rates on the exchange rate.[7]

1.5.1 Including market sentiment variable and considering additional interest rate maturities

One possible concern about our main model specification is the lack of inclusion of any market sentiment variable. For this reason, we include the Vix as an endogenous variable. The results show that estimates are comparable to the baseline model, although the point estimates are slightly different and the effect of both US and euro area rates are amplified (Table 1.13 in the Appendix).

We repeat the analysis using different US and German interest rate maturities. In particular, instead of the 2-year, we use the 5- and the 10-year German and US Treasury rates (the estimates are reported in Table 1.14 in the Appendix). Once again, the results found in the baseline analysis are here confirmed, implying that changes in the structural relations between exchange rate and interest rate was not peculiar to the 2-year maturity, but that they can be extended to longer maturities. The euro area shocks are not significant in the pre-crisis and in the global financial crisis, but they gain importance during the European crisis. The pattern for the 5- and 10-year US shocks

[7] All the calibrated thresholds and window lengths needed for these new sets of estimations are reported in Table 1.12 in the Appendix.

are in line with the 2-year interest rates, the only difference being that the coefficients during the European crisis are insignificant. Finally, the estimates in Table 1.14 are in magnitude bigger than in the baseline analysis, pointing to a positive correlation between the rates' maturity and the absolute value of the estimated coefficients.

1.5.2 Alternative definition of volatility regimes

As a second robustness check, we test whether our choice of windows used to calculate the rolling standard deviations and the thresholds used to classify the residuals in high or low volatility affects the findings. Thus, we investigate how the results change when we modify the two parameters, one at the time. The main results are robust to the alternative specifications of the windows. In fact, the pattern that emerges from the table is very similar to that described in Section 1.4: the US economic conditions dominate exchange rate movements in the pre-crisis sample, while the euro area fails to be significant in the first two periods. Moreover, there is again evidence of safe haven phenomena in crisis times: during the global financial crisis the estimate of the impact of a Treasury shock on the exchange increases, getting close to zero. In the post-OMT period estimates revert back to pre-crisis behavior, with the euro area being now significant. The same pattern is confirmed when different thresholds, the only difference being that now the US impact on the exchange rate fails to be significant in the post-OMT subsample. Both sets of results are in the Appendix, Table 1.15.

1.5.3 Alternative definition of the crisis period

As a final sensitivity check, we repeat the analysis merging the two crisis period into one. Table 1.17 in the Appendix presents only the results for the new merged crisis subsample, as the pre-crisis and post-OMT results are equal to that which is reported in Table 1.3. Moreover, the table also reports how this last set of estimates varies when changing the threshold or window (one at the time) used for the estimation. From the results, it emerges that in this new crisis sample the behavior of the coefficient attached to the US interest rates is dominated by the global financial crisis sample, meaning negative, close to zero, and insignificant, while the response of the exchange rate to a euro interest rate shock is clearly mainly influenced by the European crisis period. The table also shows that these results are robust to different windows and thresholds. To conclude, none of our key results, regarding both magnitude and the direction of the coefficients, are affected in any of the robustness checks, thus implying that the main findings are robust.

1.6 Conclusion

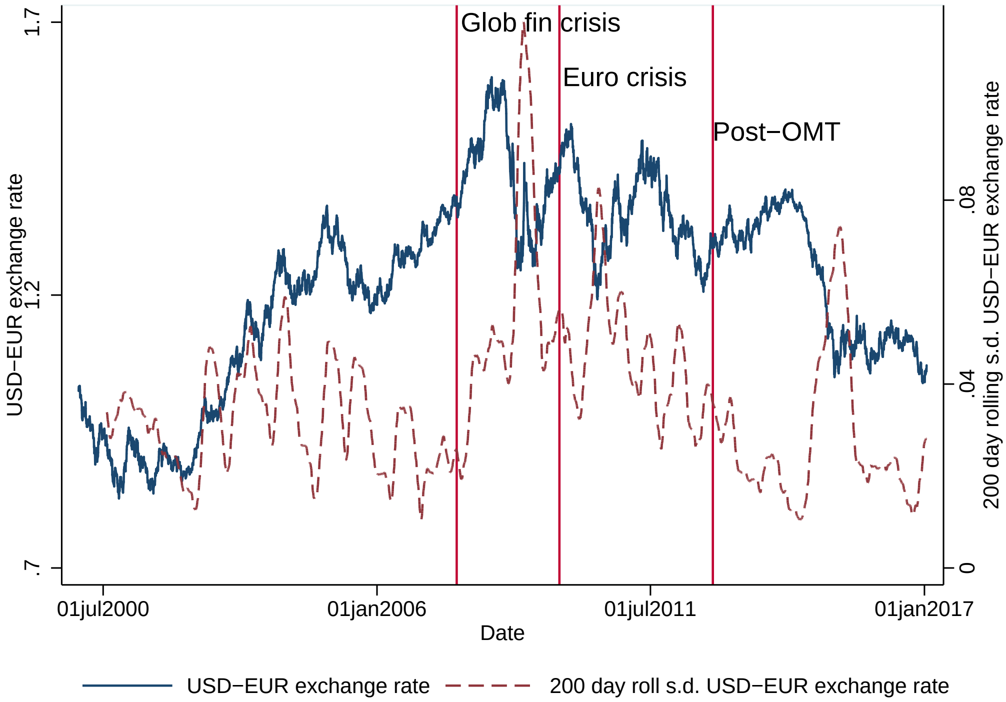
The paper studies the structural relationship between the US-dollar/euro exchange rate and the US and euro area interest rates over time. We collected daily data from January 1, 2000, to October 31, 2016, splitting it into four subsamples: pre-crisis, global financial crisis, European debt crisis, and post-crisis. For each subsample, we specify each asset price within a multifactor model and estimate the causal contemporaneous coefficients by exploiting the heteroscedasticity that is present in the data.

The key finding of the Chapter is that the relationship between the exchange rates and the interest rates changes significantly across crisis and non-crisis times. In both non-crisis periods the US dollar/euro exchange rate responds to interest rate shocks in a manner consistent with the traditional interest rate channel, that is decreasing (increasing) in response to a positive shock to the US (euro area) interest rate. In particular, US economic conditions dominate exchange rate movements before the global financial crisis (and the euro are interest rate shocks have no significant effect), while in the post-crisis sample both areas are similarly important. In the two crisis periods, results look very different. On the one hand, during the global financial crisis both US and euro area shocks have no impact; on the other, during the European crisis not only both areas have a significant effect on the exchange rate, but the impact of the US interest rate flips sign with respect to the non-crisis samples. This implies that a positive US interest rate shock has now a *negative* impact the exchange rate. These results are confirmed by the variance decomposition and by the analysis of the impact of macroeconomic news on the exchange rate.

The results are consistent with a flight to safety channel towards the dollar that alters the traditional interest rate channel of exchange rate, suggesting that in times of crisis the US dollar is considered a safe haven currency. These findings contribute to a better understanding of exchange rate behavior. In fact, they indicate the existence of time-variation in the international transmission of unconditional interest rates shocks which can play a prominent role in explaining why the UIP condition is generally found not to hold on average.

1.A Figures

Figure 1.1: Evolution of the USD/EUR exchange rate over time



Note: The figure shows the behavior of the daily USD/EUR exchange rate (blue line, left axis) together with its 200 days rolling standard deviations (red dashed line, right axis). The subsample periods are as follows. Pre-crisis: 3 Jan 2000 - 8 Aug 2007; global financial crisis: 9 Aug 2007 - 30 Sep 2009; European crisis: 1 Oct 2009 - 30 Sep 2012; post-OMT: 1 Oct 2012 - 31 Oct 2016.

1.B Tables

Table 1.7: Summary statistics of the endogenous variables used in the analysis

	Pre-crisis		Global financial crisis		European crisis		Post-OMT	
Variable	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev.
Baseline analysis								
Δ lnUSD/EUR ex rate	0.002	0.893	0.021	1.109	-0.033	0.973	-0.030	0.780
Δ2-year euro rate	-0.002	0.062	-0.011	0.102	-0.003	0.064	-0.001	0.023
Δ2-year US rate	-0.003	0.074	-0.013	0.121	-0.002	0.042	0.001	0.035
Robustness checks								
Δvix	-0.001	2.282	0.015	3.875	-0.025	2.531	0.011	1.976
Δ5-year euro rate	-0.002	0.066	-0.007	0.095	-0.005	0.082	-0.002	0.042
Δ5-year US rate	-0.002	0.086	-0.008	0.132	-0.004	0.076	0.001	0.058
Δ10-year euro rate	-0.002	0.061	-0.004	0.075	-0.005	0.077	-0.002	0.053
Δ10-year US rate	-0.002	0.083	-0.005	0.121	-0.004	0.085	0.001	0.060

Note: The table provides summary statistics for the endogenous variables used in the analysis. The reported statistics are subsample specific. The subsample periods are as follows. Pre-crisis: 3 Jan 2000 - 8 Aug 2007; global financial crisis: 9 Aug 2007 - 30 Sep 2009; European crisis: 1 Oct 2009 - 30 Sep 2012; post-OMT: 1 Oct 2012 - 31 Oct 2016.

Table 1.8: Summary statistics of euro area and US news included in the analysis

	Pre-crisis		Global financial crisis		European crisis		Post-OMT	
Macroeconomic news	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev.
Euro area								
EC Business Climate Ind	-0.002	0.199	-0.033	0.296	0.010	0.301	0.011	0.292
EC Consumer Confidence Ind	-0.003	0.139	-0.034	0.304	0.006	0.296	0.034	0.294
Euro Consumer Price Index YoY	-0.004	0.157	0.007	0.109	-0.008	0.137	0.002	0.124
Euro Consumer Price Index MoM	-0.001	0.166	0.007	0.137	0.000	0.168	0.000	0.115
Gross Fixed Capital Formation QoQ	0.001	0.094	0.006	0.116	-0.040	0.286	0.007	0.157
EC Services Confidence Ind	0.000	0.124	-0.032	0.259	-0.002	0.291	0.021	0.296
Retail Sales Volume YoY WDA	-0.011	0.247	-0.048	0.332	-0.024	0.296	0.001	0.293
Retail Sales Volume MoM SA	-0.009	0.256	-0.046	0.323	-0.039	0.305	0.009	0.269
Trade Balance with non Eurozone	-0.008	0.207	0.023	0.246	0.006	0.226	0.006	0.194
ECB M3 Money Supply 3 Month Moving Avg	0.015	0.235	-0.020	0.280	0.001	0.238	-0.004	0.210
PPI Eurozone Industry Ex Construction YoY	0.005	0.192	-0.023	0.280	0.027	0.271	-0.016	0.249
PPI Eurozone Industry Ex Construction MoM	-0.001	0.176	-0.022	0.248	0.010	0.266	-0.010	0.246
Eurostat Unemployment Rate	-0.017	0.204	0.034	0.222	0.011	0.166	-0.022	0.239
United States								
GDP SA QoQ	0.003	0.108	0.007	0.141	0.011	0.159	0.011	0.160
CPI Urban Consumers YoY	0.001	0.168	0.003	0.248	-0.002	0.217	-0.002	0.229
CPI Urban Consumers no Food&Energy YoY	-0.003	0.155	0.006	0.205	0.000	0.235	-0.007	0.197
Personal Consumption Expenditure CPI YoY	0.000	0.125	0.000	0.153	-0.005	0.198	-0.005	0.194
CPI Urban Consumers MoM SA	-0.006	0.232	0.007	0.235	-0.008	0.220	-0.012	0.221
Uni of Michigan Consumer Conf Ind	0.006	0.213	0.029	0.275	0.040	0.287	0.012	0.272
S Government Budget Balance	0.014	0.298	-0.009	0.253	-0.006	0.272	0.012	0.292
US Consumer Spending Growth Rates MoM	0.000	0.228	0.009	0.195	-0.015	0.273	0.006	0.214
Industrial Production MoM	-0.003	0.243	-0.016	0.277	-0.014	0.287	0.001	0.237
Trade Balance of Goods& Services	-0.008	0.275	0.016	0.301	0.001	0.291	-0.003	0.267
Core Producer Price Index	-0.004	0.150	0.027	0.272	0.008	0.245	-0.007	0.135
Producer Price Index - Finished Goods	0.004	0.185	0.005	0.284	0.003	0.261	-0.012	0.140
Initial Jobless Claims SA	0.009	0.622	0.043	0.619	0.046	0.631	-0.024	0.617
Housing Starts/Permits	0.017	0.292	-0.013	0.291	-0.005	0.296	-0.011	0.283
GDP Chained 2009 Dollars QoQ	0.002	0.133	0.008	0.084	-0.016	0.172	0.001	0.156
PPI Finished Goods MoM %	0.003	0.268	0.008	0.288	0.009	0.282	-0.011	0.134
Capacity Util % of Total Capacity	-0.003	0.235	-0.023	0.276	-0.008	0.266	-0.005	0.263
Business Inventories MoM	0.011	0.231	-0.028	0.231	0.007	0.241	0.010	0.246
Construction Spending Tot MoM	-0.004	0.201	0.021	0.295	0.009	0.290	-0.038	0.320
Durab Goods New Ord Indust MoM	-0.006	0.288	0.009	0.286	-0.029	0.301	0.012	0.274
Conf Board Leading Ind MoM	-0.002	0.238	0.006	0.253	0.014	0.246	0.020	0.262
Prod Output Per Hour Nonfarm Bus QoQ	0.003	0.157	0.007	0.174	0.001	0.160	0.005	0.154
Unit Labor Costs Nonfarm BusiQoQ %	0.007	0.170	0.002	0.170	0.004	0.170	-0.004	0.164
Personal Income MoM	0.009	0.250	0.035	0.293	-0.010	0.269	0.002	0.247
Nonfarm Payrolls Total MoM	-0.011	0.211	-0.010	0.181	-0.011	0.198	0.008	0.173

Note: The table provides summary statistics for the euro area and the US exogenous variables used in the analysis. The reported statistics are subsample specific. The subsample periods are as follows. Pre-crisis: 3 Jan 2000 - 8 Aug 2007; global financial crisis: 9 Aug 2007 - 30 Sep 2009; European crisis: 1 Oct 2009 - 30 Sep 2012; post-OMT: 1 Oct 2012 - 31 Oct 2016.

Table 1.9: Windows and thresholds used for the computation and the classification of reduced form residuals

	Pre-crisis	Global financial crisis	European crisis	Post-OMT
Window (days)	20	8	24	16
Threshold (st. dev.)	0.8	0.9	1.1	0.9

Note: The table reports the subsample-specific windows and thresholds used to determine the volatility regimes. First, we estimate the reduced form model. Then, we compute the rolling standard deviation for each reduced form residual u_t, using the window reported above. Finally, we calibrate the threshold for the rolling standard deviation that defines whether the residual should be classified into a high or low regime. For each subsample we identify five volatility regimes and we define them as follows: 1) all residuals are classified in low volatility; 2) only one residual is in high volatility, the others are in low; 3) all residuals display high volatility. Observations that cannot be classified into any of the five regimes are dropped from the estimation.

Table 1.10: Summary statistics of statistically defined regimes

	Pre-crisis		*Global financial crisis*	
Volatility Regime	Observation	%	Observation	%
1	131	13.21	77	27.50
2	123	12.40	37	13.21
3	16	1.61	37	13.21
4	51	5.14	31	11.07
5	430	43.35	30	10.71
Dropped	241	24.29	68	24.29
Total	992	100	280	100
	European crisis		*Post-OMT*	
Volatility Regime	Observation	%	Observation	%
1	163	41.69	124	23.26
2	32	8.18	33	6.19
3	45	11.51	74	13.88
4	52	13.30	77	14.45
5	37	9.46	43	8.07
Dropped	62	15.86	182	34.15
Total	391	100	533	100

Note: The table shows the number and the share of observations per volatility regime and subsample. The definition of the volatility regimes is based on a statistical procedure that uses the rolling standard deviations (in days) and a thresholds defined in Table 1.9 to classify the reduced-form residuals. For each subsample we identify five volatility regimes and we define them as follows: 1) all residuals are classified in low volatility; 2) only one residual is in high volatility, the others are in low; 3) all residuals display high volatility. Observations that cannot be classified into any of the five regimes are dropped from the estimation.

Table 1.11: Pairwise likelihood ratio tests for constancy of the reduced form covariance matrix

Volatility regime		*Volatility regime* 1	2	3	4
		Pre-crisis			
2	LR statistic	29.69			
	p-value	(0.00)			
3	LR statistic	27.03	25.43		
	p-value	(0.04)	(0.00)		
4	LR statistic	25.44	32.74	31.24	
	p-value	(0.00)	(0.00)	(0.00)	
5	LR statistic	171.12	114.49	35.91	52.45
	p-value	(0.00)	(0.00)	(0.00)	(0.00)
		Global financial crisis			
2	LR statistic	36.91			
	p-value	(0.00)			
3	LR statistic	26.90	35.14		
	p-value	(0.00)	(0.00)		
4	LR statistic	13.53	44.52	33.40	
	p-value	(0.00)	(0.00)	(0.00)	
5	LR statistic	78.75	67.70	46.95	36.17
	p-value	(0.00)	(0.00)	(0.00)	(0.00)
		European crisis			
2	LR statistic	20.03			
	p-value	(0.02)			
3	LR statistic	50.32	16.72		
	p-value	(0.00)	(0.01)		
4	LR statistic	23.21	13.98	24.36	
	p-value	(0.00)	(0.03)	(0.00)	
5	LR statistic	106.80	27.32	39.35	29.97
	p-value	(0.00)	(0.00)	(0.00)	(0.00)
		Post-OMT			
2	LR statistic	42.46			
	p-value	(0.00)			
3	LR statistic	40.60	44.61		
	p-value	(0.00)	(0.00)		
4	LR statistic	57.49	48.03	106.02	
	p-value	(0.00)	(0.00)	(0.00)	
5	LR statistic	123.71	30.50	76.25	78.92
	p-value	(0.00)	(0.00)	(0.00)	(0.00)

Note: The table shows results of likelihood ratio tests on the null hypothesis that, within a subsample, pairwise regimes have the same reduced form covariance matrix. The first row from the top and the fist column from the left indicate the volatility regime. For instance, the cell at the intersection between regime 1 and regime 2 in the pre-crisis subsample (the value shown in that cell is 29.69) contains the likelihood ratio test statistic of the test on the null hypothesis that regime 1 and regime 2 have the same variance-covariance matrix. Bootstrap p-values are in parentheses.

Table 1.12: Windows and thresholds used for the computation and the classification of reduced form residuals - sensitivity analysis

	Pre-crisis	Global financial crisis	European crisis	Post-OMT
			Including VIX	
Window (days)	14	6	12	4
Threshold (st. dev.)	0.7	0.9	1.1	0.7
			Different windows	
Window (days)	16	12	26	8
Threshold (st. dev.)	0.8	0.9	1.1	0.9
			Different threshold	
Window (days)	20	8	24	16
Threshold (st. dev.)	0.7	0.9	1	1.2
			Merged crisis	
Window (days)	.	.	14	.
Threshold (st. dev.)	.	.	1	.
			Merged crisis - different window	
Window (days)	.	.	16	.
Threshold (st. dev.)	.	.	1	.
			Merged crisis - different threshold	
Window (days)	.	.	14	.
Threshold (st. dev.)	.	.	1.1	.
			5-year interest rates	
Window (days)	8	14	10	4
Threshold (st. dev.)	0.9	0.9	1	0.9
			10-year interest rates	
Window (days)	18	4	10	16
Threshold (st. dev.)	0.8	0.9	0.9	1

Note: The table reports the windows and the thresholds used to determine the volatility regimes in the sensitivity analysis, Section 1.5. First, we estimate the reduced form model. Then we compute the rolling standard deviation for each reduced form residual u_t, using the window reported above. Then we calibrate the threshold for the rolling standard deviation that defines whether the residual should be classified into a high or low volatility regime. For each subsample, we identify five volatility regimes and we define them as follows: 1) all residuals are classified in low volatility; 2) only one residual is in high volatility, the others are in low; 3) all residuals display high volatility. Observations that cannot be classified into any of the five regimes are excluded from the estimation.

Table 1.13: Contemporaneous effects among endogenous variables - including the VIX

	Impulse			
Response	USD/EUR exchange rate	2-year euro area rate	2-year US interest rate	Vix
		Pre-crisis		
USD/EUR ex rate	1.00	0.06	-0.32	-0.13
p-value	.	.	.a	.c
2-year euro rate	-0.11	1.00	0.31	-0.7
p-value	.	.	.a	.
2-year US rate	-0.04	0.31	1.00	-0.28
p-value	.	.a	.	.
Vix	0.11	-0.13	0.13	1.00
p-value	.	.	.	.
		Global financial crisis		
USD/EUR ex rate	1.00	0.12	-0.06	-0.24
p-value	.	.	.	.
2-year euro rate	0.05	1.00	0.22	-0.16
p-value	.	.	.b	.
2-year US rate	-0.12	0.33	1.00	-0.9
p-value	.	.a	.	.
Vix	-0.07	-0.05	-0.12	1.00
p-value	.	.	.	.
		European crisis		
USD/EUR ex rate	1.00	0.45	0.34	0.12
p-value	.	.b	.c	.
2-year euro rate	-0.01	1.00	0.39	-0.53
p-value	.	.	.b	.
2-year US rate	-0.58	0.08	1.00	-0.19
p-value	.b	.	.	.
Vix	-0.27	0.40	-0.35	1.00
p-value	.	.	.	.
		Post-OMT		
USD/EUR ex rate	1.00	0.17	-0.24	-0.04
p-value	.	.b	.a	.
2-year euro rate	0.19	1.00	0.22	0.17
p-value	.c	.	.b	.c
2-year US rate	-0.15	0.22	1.00	-0.11
p-value	.c	.a	.	.
Vix	0.03	-0.19	0.02	1.00
p-value	.	.b	.	.

Note: The table shows, for each of the four subsamples, the estimated direct effects of a one standard deviation structural shocks on the endogenous variables, based on a four-variable SVAR identified through heteroscedasticity. The impulse variables are in columns, response variables are in rows. For ease of interpretation, the signs of the off-diagonal elements are reversed. The subsample periods are as follows. Pre-crisis: 3 Jan 2000 - 8 Aug 2007; global financial crisis: 9 Aug 2007 - 30 Sep 2009; European crisis: 1 Oct 2009 - 30 Sep 2012; post-OMT: 1 Oct 2012 - 31 Oct 2016. .a, .b, .c below point estimates denote significance at the 1%, 5%, 10% level, respectively.

Table 1.14: Contemporaneous effects among endogenous variables - different interest rate maturities

	Impulse					
	5 year rate			10 year rate		
Response	USD/EUR ex rate	5y euro area rate	5y US int rate	USD/EUR ex rate	10y euro area rate	10y US int rate
	Pre-crisis					
USD/EUR ex rate	1.00	0.03	-0.30	1.00	-0.02	-0.38
p-value	.	.	.b	.	.	.a
Euro area int rate	-0.1	1.00	0.43	-0.09	1.00	0.45
p-value	.	.	.b	.	.	.a
US int rate	0.04	0.23	1.00	0.07	0.39	1.00
p-value	.	.	.b	.	.a	.
	Global financial crisis					
USD/EUR ex rate	1.00	0.24	-0.02	1.00	0.12	-0.08
p-value	.	.	.	.	.	.
Euro area int rate	-0.01	1.00	0.23	0.03	1.00	0.30
p-value	.	.	.c	.	.	b
US int rate	-0.08	0.35	1.00	-0.10	0.40	1.00
p-value	.	.b	.	.	.a	.
	European crisis					
USD/EUR ex rate	1.00	0.47	0.06	1.00	0.51	-0.42
p-value	.	.a	.	.	.a	.
Euro area int rate	0.22	1.00	0.42	0.13	1.00	0.41
p-value	.	.	.a	.	.	.a
US int rate	-0.11	0.35	1.00	0.27	0.35	1.00
p-value	.	.a	.	b	.a	.
	Post-OMT					
USD/EUR ex rate	1.00	0.19	-0.36	1.00	0.35	-0.36
p-value	.	.a	.a	.	.b	.b
Euro area int rate	0.24	1.00	0.47	0.05	1.00	0.34
p-value	.a	.	.a	.	.	.a
US int rate	0.09	0.27	1.00	-0.08	0.61	-1.00
p-value	.	.b	.	.	.a	.

Note: The table shows, for each of the four subsamples, the estimated direct effects of a one standard deviation structural shocks on the endogenous variables, based on two distinct structural VARs, both identified through heteroscedasticity. The interest rates in the response column are of the same maturity of the shocked interest rates. For instance, the value 0.23 at the intersection between the third row and the second column corresponds to the response of the 5-year US interest rate to a shock to the 5-year euro area interest rate. For ease of interpretation, the signs of the off-diagonal elements are reversed. For each subsample, the threshold and the windows used to classify residual in low or high volatility regime are reported in Table 1.12. The subsample periods are as follows. Pre-crisis: 3 Jan 2000 - 8 Aug 2007; global financial crisis: 9 Aug 2007 - 30 Sep 2009; European crisis: 1 Oct 2009 - 30 Sep 2012; post-OMT: 1 Oct 2012 - 31 Oct 2016. .a, .b, .c below point estimates denote significance at the 1%, 5%, 10% level, respectively.

Table 1.15: Contemporaneous effects among endogenous variables - different window

Response	Impulse: USD/EUR exchange rate	Impulse: 2-year euro area rate	Impulse: 2-year US interest rate
Pre-crisis			
USD/EUR ex rate	1.00	-0.12	-0.24
p-value	.	.	.b
2-year euro rate	-0.09	1.00	0.42
p-value	.	.	.a
2-year US rate	-0.05	0.25	1.00
p-value	.	.a	.
Global financial crisis			
USD/EUR ex rate	1.00	0.14	-0.07
p-value	.	.	.
2-year euro rate	0.07	1.00	0.42
p-value	0.39	.	.a
2-year US rate	-0.04	0.32	1.00
p-value	0.52	.a	.
European crisis			
USD/EUR ex rate	1.00	0.30	0.31
p-value	.	.a	.c
2-year euro rate	0.21	1.00	0.29
p-value	.c	.	.b
2-year US rate	-0.43	0.32	1.00
p-value	.b	.a	.
Post-OMT			
USD/EUR ex rate	1.00	0.17	-0.29
p-value	.	.a	.b
2-year euro rate	0.29	1.00	0.39
p-value	.b	.	.a
2-year US rate	0.16	0.21	1.00
p-value	.	.b	.

Note: The table shows, for each of the four subsamples, the estimated direct effects of a one standard deviation structural shocks on the endogenous variables, based on a three-variable SVAR identified through heteroscedasticity. For each subsample, the thresholds and the windows used to classify residual in low or high volatility regime are reported in Table 1.12. The impulse variables are in columns, response variables are in rows. For ease of interpretation, the signs of the off-diagonal elements are reversed. The subsample periods are as follows. Pre-crisis: 3 Jan 2000 - 8 Aug 2007; global financial crisis: 9 Aug 2007 - 30 Sep 2009; European crisis: 1 Oct 2009 - 30 Sep 2012; post-OMT: 1 Oct 2012 - 31 Oct 2016. .a, .b, .c below point estimates denote significance at the 1%, 5%, 10% level, respectively.

Table 1.16: Contemporaneous effects among endogenous variables - different threshold

Response	Impulse: USD/EUR exchange rate	Impulse: 2-year euro area rate	Impulse: 2-year US interest rate
Pre-crisis			
USD/EUR ex rate	1.00	0.10	-0.23
p-value	.	.	.c
2-year euro rate	-0.19	1.00	0.37
p-value	.	.	.a
2-year US rate	-0.14	0.23	1.00
p-value	.b	.b	.
Global financial crisis			
USD/EUR ex rate	1.00	0.21	-0.01
p-value	.	.	.
2-year euro rate	0.16	1.00	0.45
p-value	.	.	.a
2-year US rate	-0.15	0.33	1.00
p-value	.	.b	.
European crisis			
USD/EUR ex rate	1.00	0.29	0.23
p-value	.	.b	.c
2-year euro rate	0.15	1.00	0.37
p-value	.	.	.a
2-year US rate	-0.34	0.24	1.00
p-value	.a	.b	.
Post-OMT			
USD/EUR ex rate	1.00	0.32	-0.22
p-value	.	.a	.
2-year euro rate	0.29	1.00	0.38
p-value	.a	.	.a
2-year US rate	-0.22	0.29	1.00
p-value	.	.a	.

Note: The table shows, for each of the four subsamples, the estimated direct effects of a one standard deviation structural shocks on the endogenous variables, based on a three-variable SVAR identified through heteroscedasticity. For each subsample, the thresholds and the windows used to classify residual in low or high volatility regime are reported in Table 1.12. The impulse variables are in columns, response variables are in rows. For ease of interpretation, the signs of the off-diagonal elements are reversed. The subsample periods are as follows. Pre-crisis: 3 Jan 2000 - 8 Aug 2007; global financial crisis: 9 Aug 2007 - 30 Sep 2009; European crisis: 1 Oct 2009 - 30 Sep 2012; post-OMT: 1 Oct 2012 - 31 Oct 2016. .a, .b, .c below point estimates denote significance at the 1%, 5%, 10% level, respectively.

Table 1.17: Contemporaneous effects among endogenous variables - merged crisis subsample

Response	USD/EUR exchange rate	2-year euro area rate	2-year US interest rate
	Impulse		
	Crisis subsample, main result		
USD/EUR ex rate	1.00	0.35	-0.06
p-value	.	.a	.
2-year euro rate	-.01	1.00	0.29
p-value	.	.	.a
2-year US rate	-0.11	0.24	1.00
p-value	.	.	.
	Crisis subsample, different window		
USD/EUR ex rate	1.00	0.30	-0.07
p-value	.	.b	.
2-year euro rate	0.08	1.00	0.37
p-value	.	.	.a
2-year US rate	-0.04	0.19	1.00
p-value	.	.b	.
	Crisis subsample, different threshold		
USD/EUR ex rate	1.00	0.30	-0.05
p-value	.	.b	.
2-year euro rate	0.02	1.00	0.21
p-value	.	.	.b
2-year US rate	-0.07	0.23	1.00
p-value	.	.a	.

Note: The table shows, for each of the four subsamples, the estimated direct effects of a one standard deviation structural shocks on the endogenous variables, based on a three-variable SVAR identified through heteroscedasticity. In this case, the two crisis subsamples, the GFC subsample and the EC subsample have been merged into a more general crisis subsample. Only results for the crisis subsample are shown, as the results for the pre-crisis subsample and the post-OMT are identical to results shown in Table 1.3. For each subsample, the thresholds and the windows used to classify residual in low or high volatility regime are reported in Table 1.12. The impulse variables are in columns, response variables are in rows. For ease of interpretation, the signs of the off-diagonal elements are reversed. The subsample periods are as follows. Pre-crisis: 3 Jan 2000 - 8 Aug 2007; global financial crisis: 9 Aug 2007 - 30 Sep 2009; European crisis: 1 Oct 2009 - 30 Sep 2012; post-OMT: 1 Oct 2012 - 31 Oct 2016. .a, .b, .c below point estimates denote significance at the 1%, 5%, 10% level, respectively.

CHAPTER 2

Unconventional monetary policy and households' financial portfolio choices[1]

2.1 Introduction

The unconventional monetary policies (UMPs) introduced by the European Central Bank (ECB) with the aim of restoring confidence in the European financial system have spurred an intense public and academic debate about their financial and real implications on different economic agents. In particular, the discussion on whether ultra-loose monetary policy impacts income and wealth inequality has caught the attention of the public and policy-makers in the Eurozone (Draghi, 2015; Panetta, 2015; Draghi, 2016; Constâncio, 2017). Among other concerns, commentators point out that expansionary measures that positively affect the price of financial assets would only benefit the owners of financial wealth, thus increasing the gap between households that invest in financial markets and those who do not. In order to analyze the issue in depth, Adam and Tzamourani (2016), Casiraghi et al. (2018), Ampudia et al. (2018), and Lenza and Slacalek (2018) perform simulation exercises where they focus on the UMP's effects on financial wealth through capital gains. These papers, though, assume that households are only passively affected by monetary policy through its impact on asset prices but that this effect does not translate into an active choice of portfolio rebalancing. However, this view is hard to reconcile with the fact that unconventional tools are also expected to work by affecting investors' portfolio allocation choices (e.g.,

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through the portfolio rebalancing channel or the signalling channel of monetary policy) and investors' risk appetite (e.g., through the confidence channel), both domestically (Krishnamurthy and Vissing-Jorgensen 2011; Krishnamurthy et al. 2017) and internationally (Fratzscher et al., 2018).

This Chapter analyzes how unconventional tools affect households' portfolio choices, addressing a specific case: the impact of ECB's unconventional policies on Italian households' financial choices between 2006 and 2016. The contribution of this work is threefold. First, this is the first paper to focus on, and emphasize, the role that unconventional monetary policy plays in households' portfolio rebalancing decisions. Second, I use survey data on Italian households' portfolio composition and augment it with extra financial data to construct a novel dataset that, for each household and financial asset class, disentangles any change in financial wealth into its capital gain and active rebalancing components. Third, this work brings new evidence to the limited literature that evaluates UMPs effect on Italy, one of the so-called peripheral countries heavily affected by the crisis and, thus, most suitable for evaluating the effectiveness of ECB's actions.

Understanding how unconventional tools shape households' financial choices is important from several points of view. First, households hold an important stock of financial wealth and their investment decisions may have a significant impact on asset prices (Kogan et al., 2006; Kumar and Lee, 2006) and even on the macroeconomy (Korniotis and Kumar, 2010). To give an example, between 2007 and 2014, Italian households invested on average EUR 3,7 trillion in financial assets, of which EUR 205 billion in Italian government securities, financing roughly 10% of the government debt.[2] Thus, their rebalancing decisions can have important consequences at both the financial and real economy levels. Second, it provides new insights into the effectiveness and the transmission channels of unconventional measures in a country that was at the core of the sovereign crisis. Third, this Chapter contributes to the debate on inequality by pointing toward an under-explored additional channel through which UMPs might affect wealth inequality: the one stemming from investors' heterogeneous reaction to monetary stimulus. An extended literature documents that richer and better educated households (the so-called sophisticated investors) rebalance their portfolio more frequently (Vissing-Jorgensen, 2002; Campbell, 2006; Calvet et al., 2009b; Bilias et al., 2010) and make smaller investment mistakes (Calvet et al., 2007, 2009b). Thus, only a small group of households might decide to shape their investment decisions so to

[2] Bank of Italy, Supplement to the Statistical Bulletin, Household wealth in Italy in 2014. `https://www.bancaditalia.it/pubblicazioni/ricchezza-famiglie-italiane/2015-ricchezza-famiglie/en_suppl_69_15.pdf?language_id=1`.

take better advantage of the new unconventional environment. This could have serious consequences on the wealth distribution, as it might exacerbate inequality.

I use the micro data contained in the Bank of Italy Survey on Household Income and Wealth (SHIW) for the 2006-2016 period. This dataset gathers data every other year on wealth and other aspects of households' economic and financial behavior. Each wave includes roughly 8,000 households. This data allows constructing financial portfolios, but it does not include any additional information that can be used to distinguish whether a change in financial wealth is due to valuation effects (passive saving component/capital gains) or to an active reallocation (active saving component/active rebalancing). This distinction is crucial for this Chapter, as the aim of this study is to capture how UMPs shape households' financial decisions, i.e. how much households actively and voluntarily decide to rebalance their portfolio. Following Guiso et al. (2002b), Berben et al. (2006), and Juster et al. (2006), I approximate the return on several asset classes included in the survey with financial indexes. This allows me to isolate the active saving component of the ten most important Italian asset classes (four types of government bonds, equity, corporate bonds, liquidity funds, flexible funds, bond funds, and equity funds). Furthermore, the richness of the financial wealth information included in the SHIW and the possibility to isolate the pure rebalancing component supports the choice of focusing on the Italian case.

One of the challenges of this Chapter is the identification of the causal relationship between UMPs and households' portfolio rebalancing due to the low frequency of the household data and concerns over simultaneity and omitted variable bias. Therefore, I employ a two step approach. First, I follow Altavilla et al. (2014), and Fratzscher et al. (2018) by using an event-study approach to estimate the impact of monetary policy shocks. UMP shocks are proxied by announcements and measured through their valuation impact on the return of the financial indexes used to construct the dataset. This gives a clear picture of the effects that UMPs have on the asset classes I include in the analysis. The choice of a high-frequency, financial market-driven procedure ensures that any change in the indexes around the time of any unconventional monetary policy announcement can be only attributed to the unexpected component of monetary policy revealed by the announcement itself, while the magnitude of the shock is extracted directly from the financial indexes' response. Then, I exploit households' cross sectional heterogeneity in their exposure to unconventional monetary policy due to their heterogeneous stock of financial investment to sharpen the identification. The intuition is that the more financial wealth households hold, the more they are affected by monetary policy changes due to the valuation effect. Comparing the behavior of households expected to be more affected with those less affected should shed light on the effects of UMPs. This two step procedure allows for a straightforward interpretation of the identification: it isolates the portion of capital gains due to exogenous UMP-valuation effects. Thus, the empirical analysis estimates how much of this exogenous change in financial wealth is passed through to portfolio rebalancing.

My main result is that unconventional tools directly affect households' portfolio composition. In line with the literature finding that portfolio rebalancing is positively correlated with income and wealth, this is true only for households in the top 25% of the income distribution. An accommodative ECB announcement that increases return of financial indexes by 1% induces, on average, a positive investment of EUR 5000, of which one third in government bonds and two thirds in risky assets. These findings speak in favor of ECB's ability to restore household confidence in the financial system and are consistent with the confidence channel of unconventional monetary policy. Further evidence supporting this channel comes from limiting the estimation sample to the end of 2014. Excluding the last rounds of unconventional measures - initiated to sustain economic growth and inflation rather than to relieve sovereign stress and fight the redenomination risk - quantitatively increases households' response to unconventional shocks. Finally, I use a qualitative measure of the households' change in risk aversion to gain more insights on the relationship between household risk taking and unconventional measures. The analysis finds that UMP has a significant and negative effect on households' risk aversion.

The empirical finding that only top income earners rebalance after a UMP shock can be interpreted as a first evidence that unconventional monetary policy could amplify the (financial) wealth inequality between sophisticated and non-sophisticated investors. This idea is corroborated by the results of a simulation exercise (Section 2.6) finding that, when considering both UMP-induced valuation and rebalancing effects, rich households' portfolio return cumulated over the 2006-2016 period is ten times higher than that of low/middle income investors. The simulation also shows that this difference is only attributable to portfolio rebalancing. This Chapter does not consider the total financial portfolio but only some of the asset classes households invest in. Thus, it is not possible to draw direct conclusions about financial wealth inequality. Nevertheless, my findings point toward the heterogeneity of portfolio rebalancing as important channel through which unconventional monetary policy can affect wealth inequality.

This Chapter relates to different strands of literature. Few works evaluate the impact and transmission channels of ECB programs on the Italian financial markets and macroeconomic variables (see, among others, Altavilla et al., 2014; Fratzscher et al., 2016; Falagiarda and Reitz, 2015; Casiraghi et al., 2016; Krishnamurthy et al., 2017).

They all find that ECB policies are positively associated with an increase in financial prices, real activity, and credit. Most closely related to this Chapter is the study by Casiraghi et al. (2016), who assess the impact of the main unconventional measures on the Italian economy. In line with this Chapter, they first use a high-frequency event study approach to estimate the UMP's direct effects on financial markets and then estimate the impact of UMPs on the main Italian macroeconomic variables using a much lower frequency model. They find that UMPs have, to varying degrees, served to counteract the increase in government bond yields and had a large positive effect at the macro level. My findings are in line with theirs and, in addition, I show that the ECB was not only effective in sustaining Italian financial segments, but also in restoring confidence to private investors.

There is a flourishing literature showing that households' risk aversion is time-varying and that macro/financial events affect portfolio decisions through their impact on households' wealth. This literature is nicely connected to this work not only because, consistent with several papers (Fratzscher et al., 2016; Falagiarda and Reitz, 2015), I find that unconventional tools negatively affect agents' risk aversion, but also because I identify UMP shocks as an exogenous change in wealth induced by monetary policy on portfolio rebalancing. There are no studies linking unconventional (or conventional) monetary policy shocks to households' portfolio rebalancing, but a few papers consider the relationship between financial fluctuations and portfolio rebalancing in the context of the 2007-2009 financial crisis. Bucher-Koenen and Ziegelmeyer (2013) show that crisis episodes might have a negative impact on households' participation in financial markets both in the short- and long-run. This is especially true for households with lower levels of financial literacy. Interestingly, the authors point out that for financially illiterate households, this behavior might have serious consequences regarding wealth distribution, as they would fail to benefit from market resurgence in the short-run and from equity premium in the long-run. Guiso et al. (2018) also find that, after the 2008 drop in stock prices, Italian investors rebalanced their portfolio in a way consistent to a fear model (i.e. selling stocks). On the contrary, using Dutch data, Hoffmann et al. (2013) find that individual investors continue to trade actively and do not de-risk their investment portfolios. Consistent with the first three papers, I show that households respond to an exogenous increase in financial wealth (thus, to a positive shock) driven by unconventional measures by buying more assets. However, this is only true for richer households, as for the bottom 75% of income distribution "*inertia seems to be the main driver of portfolio allocation*" (Brunnermeier and Nagel, 2008).

Finally, this Chapter contributes to the growing literature that uses detailed balance sheet data to estimate the distributional effects of unconventional monetary policy on wealth inequality. Bivens (2015) for the US, Domanski et al. (2016) for a set of advanced countries, and Ampudia et al. (2018) for the Euro Area, provide data-driven simulation exercises. On the one hand, they find that, given the initial distribution of financial wealth, unconventional tools mainly accrue wealthiest investors. On the other, this negative effect is outweighed by UMP ability to sustain economic activity, employment, and house prices, which is especially beneficial for low-income households. They conclude that unconventional tools tend to reduce income and wealth inequality. Casiraghi et al. (2018) (on Italian data) and Bunn et al. (2018) (on UK data) use elasticities from a large-scale econometric model of the Italian and UK economy, respectively, and reach the same conclusions. My work is complementary to this literature and uncovers an additional channel through which UMP exacerbates wealth inequality.

The paper proceeds as follows. Section 2.2 reviews ECB's main policy actions during the crisis and UMPs main transmission channels. Section 2.3 describes the data and the construction of the dataset. Section 2.4 discusses unconventional monetary policy, the empirical framework, and the results. Section 2.5 presents robustness checks. Section 2.6 quantifies the heterogeneous UMPs impact on financial wealth along the income distribution. Section 2.7 concludes.

2.2 The crises, ECB's unconventional monetary policy, and its transmission channels

Between 2008 and 2016, the ECB faced three different crisis phases (Praet, 2018). The first phase coincides with the liquidity crisis triggered by the onset of the global financial cycle and the collapse of Lehman Brothers. The principle ECB response was to lower its main refinancing rate to 1% (May 2009), expand the number of eligible collateral for refinancing operations, and provide liquidity to the banking sector. The second phase is the sovereign debt crisis of 2011-12. Italy and the other GIPS countries (Greece, Ireland, Portugal, and Spain) saw large increases in their government bond yields. The ECB's main policy response included the direct purchases of government bonds through the Securities Market Program (SMP, May 2010), two three-year refinancing operations (LTROs, December 2011 and February 2012) and the announcement of the conditional Outright Monetary Transactions (OMT, August 2012). The explicit aim of these programs was to reduce the perception of redenomination risk and financial market anomalies, such as fragmentation and illiquidity. Although the extraordinary injection of liquidity proved useful, acting as powerful circuit breaker stopping the downward spiral, by mid-2014, a credit crunch was looming and the economic recovery was losing momentum. With the risk of low inflation and de-anchored inflation expectations on the rise, the ECB decided to ease its monetary stance further by directly intervening across the whole range of interest rates affecting the financing conditions of the economy. First, it implemented a negative interest rate policy, lowering the interest rate paid on the deposit facility to -0.1% (2014). Then it promoted two rounds of longer-term refinancing operations (TLTROs, September 2014 and March 2016), in order to support bank lending to the private sector. Finally, it introduced the asset purchase programme (APP, January 2015) for both private and public sector securities. These policies helped to compress further premia along the yield curve.

There is a growing literature analyzing possible channels through which UMPs might impact financial markets (for an extensive literature review, see Haldane et al. 2016), but less is known about the mechanisms through which they affect households' portfolio decisions. In this Chapter, I focus on two transmission channels that are the most likely to affect the households' portfolio rebalancing choices: the confidence channel (and the closely connected sovereign credit risk channel) and the portfolio rebalancing channel. These two channels have opposite implications for active rebalancing following an increase in asset prices driven by UMPs valuation effect. Although the empirical strategy does not allow me to directly test/compare them, in the discussion of the results (Section 3.4) I still argue whether the empirical findings are consistent with one channel or the other.

The first channel is the *confidence channel*. It affects and influences the perception of risk and uncertainty, restoring agents' confidence in the financial system. Consequently, risk premia decline and asset prices increase (Bluwstein et al. 2016; Fratzscher et al. 2018). It was heavily stressed, especially during the first and the second phases of the crisis. A very connected channel is the so-called *sovereign credit risk channel*. One of the key goals for UMPs in the second phase of the crisis was to reduce sovereign risk premia of peripheral counties, considered excessive and not in line with fundamental but more reflecting unfounded fears of Eurozone break-up. Given Italy's peculiar role during the period of crisis, these two channels predict that investors would respond in a pro-cyclical manner to price changes. Thus, accommodative unconventional monetary policy should increase households' investment toward all asset classes, including government bonds.

The second channel is the *portfolio rebalancing channel*. Many commentators indicate it is one of the main transmission mechanisms (see, among others, Bernanke et al. 2010; Bernanke 2012; Gagnon et al. 2010; Joyce et al. 2012; Draghi 2014, 2015). The main idea is that central bank purchases affect risk premia and yields of key financial segments, inducing investors to rebalance their portfolio away from assets not directly affected by unconventional stimulus and toward investments with higher return. Thus, this channel prescribes that investors would respond in a counter-cyclical manner to price changes, selling government bonds and buying riskier assets.

2.3 Data description and portfolio composition

The empirical analysis in Section 2.4 requires disaggregated and detailed data of a representative sample on households' portfolio composition, wealth, and demographics. For this reason, I make use of the Bank of Italy's Survey of Households Income and Wealth. In this chapter, I provide an overview of the dataset and of the methodology I apply to distinguish between active and passive saving components.

2.3.1 Data sources

The main dataset is obtained from the Survey of Households Income and Wealth (SHIW), a large-scale household survey run by Banca d'Italia every second year. It contains detailed information on demographic, consumption, labor supply, income, real wealth, and financial wealth characteristics of a stratified random sample of the Italian population. SHIW is conducted since 1960 and the sample used in most recent years comprises about 8,000 households; as I focus on the unconventional monetary policy period, I rely on the waves from 2006 through 2016 (the last available survey wave). Their contents, methodology, and variable definition are broadly homogeneous. The survey contains a rotating panel component and in each wave around half of the households are participating for at least their second time. This panel component proves very useful for the determination of the change in wealth active saving component needed for the analysis. Balance-sheet entries are reported as of the end of the previous year (for example, December 31st, 2006), while flows of income and consumption refer to the previous year. The unit of observation is the household.[3]

In the empirical analysis, I focus only on the rebalancing of the financial portfolio. In fact, although for the majority of Italians the residence is the only held asset, it is acquired mainly for living necessities rather than as an investment. For this reason, throughout the Chapter it is assumed that housing is a highly illiquid asset and that it is not rebalanced in a strategic manner. Nonetheless, it will be used as a control in the empirical analysis. The same applies to deposits: I assume that increases or decreases in saving and checking accounts are linked to liquidity reasons, unrelated to investment policies. For this reason, they are also excluded from the analysis. In the SHIW

[3]The household is defined to include all persons residing in the same dwelling who are related by blood, marriage, or adoption. Individuals selected as partners or other common-law relationships are also treated as households.

questionnaire, households are presented with a fixed list of forms of financial saving and investment and they are asked if the household held any of them on December 31st of the previous year. If the answer is positive, they are asked to provide the approximate value. Table 2.5 in the Appendix contains an overview of all asset classes among which the households can choose for the years 2006 to 2016.

The asset classes included in the analysis are: Italian government bonds (bot, btp, cct, ctz), Italian mutual funds (liquidity funds, mixed funds, bond funds, equity funds), Italian equity and Italian corporate bonds.[4] [5] The other asset classes are excluded for the following reasons:

1. It is not possible to calculate the active saving component due to lack of extra information. As described later in this paragraph, this calculation is done using data available on Bloomberg. If the data provider does not have any information about these asset classes they are dropped. This is the case of shares of unlisted and private companies, shares of partnerships, other government bonds, managed portfolios, and loans to cooperatives.

2. The percentage of households holding the asset class is very low. This is the case of bonds, government bonds, shares, other assets and investment funds issued by non-residents.

3. The asset class does not appear in all waves. This is the case of shares in privatized listed companies indexed funds and non-harmonized funds.

Table 2.7 (in the Appendix) shows the percentage of households holding each asset class included in the SHIW. The table is divided in two, the upper part reporting the holding of the asset classes included in the analysis while the bottom part showing the excluded ones. Most asset classes are held by only a small fraction of households. This can be partially due to a problem of underreporting, but it is mostly to be attributed to the fact that several Italian financial markets are thin. Moreover, the table shows that, on average, between 2006 and 2016 the share of households holding financial assets decreased, implying that during the years of crisis households have partially abandoned financial investments. This trend is visible not only in the asset classes included in the analysis (with the main exception of mutual funds), but also in the excluded markets. Table 2.7 (in the Appendix) also shows that the share of households directly investing in assets issued by non-residents has not increased over time (it has decreased, if possible),

[4] The asset class mixed funds is given by the sum of balanced funds, balanced bond funds, balanced equity funds and flexible funds.

[5] For a detailed description of the asset classes included in the analysis, as well as the final classification with respect to the analysis, see Table 2.6 in the Appendix.

suggesting that there has been no rebalancing between domestic and foreign markets, but, most likely, a general reduction of financial investment. Moreover, Figure 2.6 shows that, in times of crisis, Italians have not drastically changed the composition of their financial portfolio, and this pattern is consistent across different groups of households along the income distribution (Group 1, Group 2, and Group 3, i.e. households in the bottom two quartiles, third quartile and top quartile, respectively). All groups exhibit the strongest reduction in government bonds in favor of other assets between 2010 and 2012, the period of higher stress for Italian markets, but the trend starts reversing already in 2014. Moreover, the percentage invested in foreign assets and in other assets is constant over time. Interestingly, the figure also shows that, contrary to expectations, low income earners do not invest higher share of their portfolio in government bonds compared to other households.

2.3.2 Passive and active rebalancing of portfolio

To understand the impact of UMPs on the portfolio rebalancing of household it is necessary to isolate the change in financial wealth due to an active decision of rebalancing (either positive or negative). An asset can change in value for two reasons: either some of it is sold or purchased (active saving or rebalancing) or the price of the asset changes (passive saving or capital gain). Thus, by definition the change in the financial wealth composition between two points in time is always given by the sum of the active rebalancing and the passive saving

$$\underbrace{X^j_{i,t} - X^j_{i,t-1}}_{\text{change in wealth}} = \underbrace{x^j_{i,t}p^j_{i,t}}_{\text{active rebalancing}} + \underbrace{X^j_{i,t-1}r^j_{i,t}}_{\text{capital gains}}, \tag{2.1}$$

where $X^j_{i,t}$ is the stock of money held by household i in asset j, $x^j_{i,t}$ is the flow of asset j, $p^j_{i,t}$ denotes the price of asset j at time t paid by household i and $r^j_{i,t}$ is the annual return between $t-1$ and t of asset j. Unfortunately, the SHIW does not contain questions about purchases or sales of assets. Moreover, it does not include any information about asset prices and returns. This implies that Equation (2.1) cannot be used directly to compute the active saving component. To solve this problem I first rearrange Equation (2.1) to obtain the expression for active saving:

$$x^j_{i,t}p^j_{i,t} = X^j_{i,t} - X^j_{i,t-1} - X^j_{i,t-1}r^j_{i,t}; \tag{2.2}$$

then, I approximate the missing variables. In fact, for each wave the survey provides me with the money invested in several asset classes ($X^j_{i,t}$ and $X^j_{i,t-1}$). Thus, the only thing that is missing in order to calculate the active saving is the return on assets, $r^j_{i,t}$.

I replace the return on each of the ten asset classed included in the analysis j with the return of financial indexes that summarize the behavior of the asset segment.[6] Then, I can finally apply Equation (2.2) and obtain ten active rebalancing components. It is important to notice that to use Equation (2.2) it is necessary to follow the same household for at least two consecutive waves, and that is where I make use of the rotating panel component of the SHIW. This procedure is applied to the ten classes included in the analysis (bot, btp, cct, ctz, liquidity funds, mixed funds, bond funds, equity funds, equity and corporate bonds). Moreover, it is also important to stress that my final dataset does not have a panel structure, but that it is constituted by repeated cross sections. The indexes used to approximate the returns, together with a short description, is contained in Table 2.8 and they are plotted in Figure 2.7 (both in the Appendix).

Finally, the active saving components are aggregated further more into two final categories that are then used for the empirical analysis, the government bonds active saving component (including btp, bot, ctz, and the cct) and the risky asset active saving component (including corporate bonds, equities, and mutual funds). This procedure also allows obtaining the capital gain components of the two asset categories. This classification between government bonds and risky assets is maintained throughout the entire paper.

I impose some requirements for households to be considered in the analysis. First, only households followed for at least two consequent waves are included in the analysis for the aforementioned reason. Second, both net wealth and income must be positive. Third, all positive rebalancing of risky or government bonds assets smaller than 500 Euro is considered a mistake of the household reporting of its financial variables and it is replaced with a rebalancing equal to zero.[7] Pooling all waves together, I am left with a sample of around 20,000 households.

Figure 2.1 shows the cross-section distribution of the newly created stock of active rebalancing variables.[8] Table 2.1 reports the summary statistics when considering the full sample and different groups of households along the income distribution (Group 1, Group 2, and Group 3, including households in the bottom two quartiles, third quartile and top quartile, respectively) over the 2006-2016 period. Both mean and median of

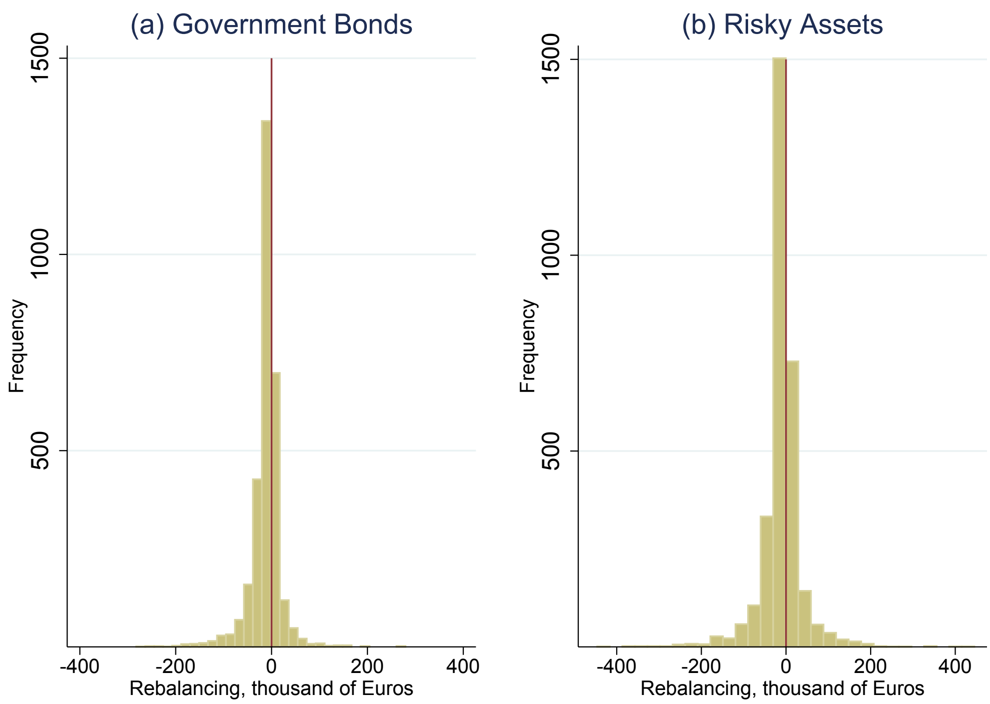
[6] The return on the asset class mixed funds is given by the average of the balance funds and flexible funds indexes returns, $r_t^{\text{mixed funds}} = \frac{1}{2}(r_t^{\text{balanced funds}} + r_t^{\text{flexible funds}})$.

[7] The idea is the following. Assume that household i has not rebalanced its portfolio between $t-1$ and t. If i does not take into account the valuation effect of its investment when reporting its investment value in t, the application of Equation (2.2) will result in an active saving different from zero. For this reason, rebalancing lower than 500 Euro are considered here as misreporting and not as an investment or disinvestment of small amounts.

[8] For a plot of the distribution of households's active rebalancing across different years and household groups, please refer to the Appendix 2.8 and 2.9, respectively.

the active rebalancing are negative for both asset categories and across all groups of households. The same results emerge when considering the evolution over time of the same two summary statistics (Figure 2.10, in the Appendix). To conclude, it appears that in the decade under analysis households have progressively reduced their financial investement in the two asset categories under analysis.

Figure 2.1: Italian households' active rebalancing, 2006-2016



Note: The figure shows the distribution of Italian households' active rebalancing for the two asset categories included in the empirical analysis, Italian government bonds (bot, ctz, btp,cct) and Italian risky assets (liquidity funds, mixed funds, bond funds, equity funds, equity, and corporate bonds) over the 2006-2016 period. The unit is thousands of euro. For a detailed explanation of the methodology used to construct the two active rebalancing categories, please refer to Section 2.3.2. Only data included in the estimation sample is reported here. Thus, rebalancing equal to zero as well as positive rebalancing in t conditioning on not being invested in the asset category in $t-1$ have been excluded from the picture.
Source: Bank of Italy's Survey on Household Income and Wealth, Bloomberg, Datastream and own calculation.

Table 2.1: Italian households' active rebalancing - summary statistics

Variable	Obs	Mean	Median	Std. Dev.	Min	Max
Government Bonds						
Group 1	966	-9207.9	-6061.9	24268.9	-193365.9	200150.6
Group 2	1163	-14567.4	-7551.9	33100.9	-282372.7	118234.9
Group 3	894	-12376.0	-8410.4	40724.9	-258490.20	279394.1
Full sample	3023	-12048.0	-7282.9	33835.8	-282372.7	279394.10
Risky Assets						
Group 1	538	-6326.0	-6765.1	28208.8	-110462.9	188918.4
Group 2	1184	-8580.4	-6078.6	47168.0	-269932.6	445047.4
Group 3	1371	-11630.8	-6356.7	57088.1	-446124.8	354559.9
Full sample	3093	-9821.5	-6259.4	50329.8	-446124.8	445047.4

Note: The table reports the summary statistics for the active rebalancing of the two categories included in the analysis, Italian government bonds (Bot, Ctz, Btp, Cct) and Italian risky assets (Liquidity funds, mixed funds, bond funds, equity funds, equity and corporate bonds) for the year 2006-2016. Summary statistics are reported for the full sample as well as for different groups of households along the income distribution: Group 1 (bottom two quartiles of the income distribution), Group 2 (third quartile) and Group 3 (top quartile). For a detailed explanation of the construction of the two active rebalancing categories, please refer to Section 2.3.2. Only data included in the estimation sample is used for the calculations. Thus, rebalancing equal to zero as well as positive rebalancing in t conditioning on not being invested in the asset category in $t-1$ have been excluded from the table. *Source*: Bank of Italy's Survey on Household Income and Wealth, Bloomberg, Datastream and own calculation.

2.4 Empirical analysis

In this section, I first present the identification strategy. Then I proceed describing the empirical framework. Finally, I discuss the baseline results.

2.4.1 The identification of unconventional monetary policy

The identification of unconventional monetary policy in this context poses several issues. First, unlike in conventional times, in the unconventional period there is not a clean single indicator of the overall stance of monetary policy. Moreover, using low frequency data could lead to endogeneity issues and omitted variable bias. Not only could the responses that follow the central bank intervention be attributed to other changes in the economy around the same time, but monetary policy could also be responding to important news affecting both monetary policy itself and the other variables under investigation (Gürkaynak et al., 2005).

In order to construct a UMPs measure and achieve identification, I build on the literature that estimates the impact of monetary policy changes using high frequency data. Here the effect of unconventional tools is extracted directly from the high frequency response of financial markets to unconventional announcements. Many authors rely on announcements as a source of UMPs identification, using an event study approach or structural VAR models (see, among others, Wright 2012; Gertler and Karadi 2015; Rogers et al. 2018; Hachula et al. 2020). The idea is that any high frequency financial indicator (indexes, prices, yields, etc...) close before a monetary announcement has already priced in the endogenous response of monetary policy to the state of the economy. Thus, any variation that occurs in a (small enough) window around a monetary policy announcement must reflect only the unexpected component of monetary policy revealed by the announcement itself and it is interpreted as exogenous with respect to the economy. Second, I borrow from the literature evaluating the impact of UMPs on banks. Here the heterogeneity across banks is exploited to assist with the identification (see, among others, Albertazzi et al. 2018; Chakraborty et al. 2020; Acharya et al. 2018). The idea is that some banks are expected to be more affected than others (due to a different composition of their balance sheet) and their different reaction to monetary shocks can, in turn, help identify the causal effect of monetary policy. The same reasoning can be applied to households: it is possible to exploit their financial portfolio composition to assess their exposure to unconventional monetary. Then, comparing the behavior of more and less affected HHs, allows shedding light on the effects of unconventional measures on portfolio rebalancing choices. In particular, I employ as exposure measure the stock of assets invested in $t-1$ in both the categories under analysis, risky assets and government bonds.

In order to have a clear understanding of the effects of UMPs on the asset classes included in the analysis, I estimate the effect of unconventional monetary policy announcements on the ten indexes used to approximate the returns (see Section 2.3.2). Following, among others, Fratzscher et al. (2018) and Altavilla et al. (2014), I employ the following univariate model

$$r_{j,t} = \alpha_j + \sum_{a=1}^{A} \beta_a D_{a,t} + \gamma \Delta \text{Eonia}_t + \sum_{n=1}^{N} \delta_n z_{n,t} + \eta_{j,t} \tag{2.3}$$

where $r_{j,t}$ is the daily return of financial index j at time t; $j=$ bot index, btp index, ctz index, cct index, equity index, corporate bond index, liquidity funds index, mixed funds index, bond funds index, and equity funds index; α_j is the index specific constant; $D_{a,t}$ is a dummy variable equal to 1 if the unconventional monetary policy announcement $a = 1, .., A$ takes place, zero otherwise. The event dummies reflect the major unconventional monetary policy announcement-related events that occurred between 2007 and 2016. The events are chosen following Hachula et al. (2020) and listed in Table 2.10; the Eonia in first difference, ΔEonia_t accounts for conventional monetary policy; $z_{n,t}$ is the release of macro news for Italy, Euro Area, and the US. See Table 2.9 (in the Appendix) for a list of all included macro surprises. The coefficients of interest are the $\beta_{a,j}$, as each of them captures the change in the return $r_{j,t}$ in response to the ECB announcements a.

After estimating Equation (2.3), the following procedure is applied:

1. The estimated vector $(\beta_{j,1}, .., \beta_{j,A})^T$ is transformed into a daily binary variable $m_{\beta,j,t}$ that takes value 0 on non-announcement days and value β_a on the day of announcement a.

2. The vector $m_{\beta,j,t}$ is then aggregated into a biennial series, $ump_{j,t}$, by summing within two years. Thus, each one of the five data points (2008, 2010, 2012, 2014, 2016) composing the vector $ump_{j,t}$ is the cumulated impact of UMPs announcements on asset j in the previous two year window.

3. Finally, two UMPs impact measures, one for risky assets and one for government bonds, are constructed by averaging out the single UMPs impact measures

$$\text{UMP impact measure}_t^{(f)} = UMP_{t,average}^{(f)} = \sum_{j=1}^{J} \frac{ump_{j,t}^{(f)}}{J}$$

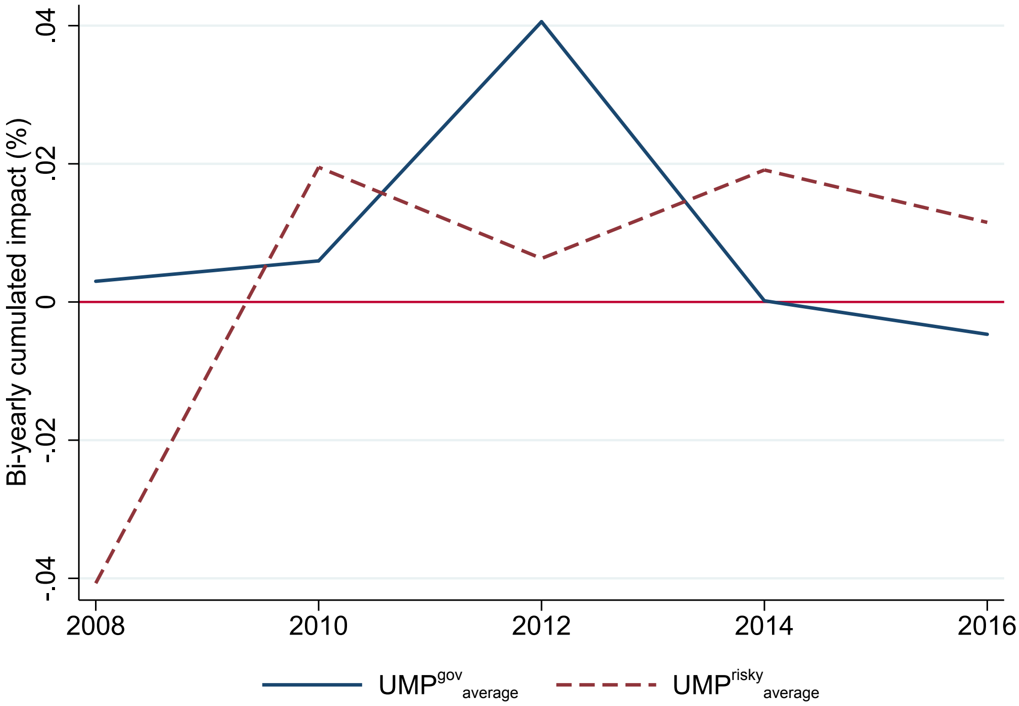
where f={gov, risky} and j= bot index, btp index, ctz index, cct index if f=gov bonds and j= equity index, corporate bond index, liquidity funds index, mixed funds index, bond funds index and equity funds index if f=risky. For a plot of $UMP_{t,average}^{gov}$ and $UMP_{t,average}^{risky}$, see Fig. 2.2.

After having constructed the (common) UMPs impact measures, I interact them with the household-specific financial investment in government bonds and risky assets at time $t-1$ and construct the UMP variable,

$$\text{UMP Var}_{i,t} = UMP_t^{gov} \times \text{gov bonds}_{i,t-1} + UMP_t^{risky} \times \text{risky}_{i,t-1}.$$

The way the UMP $\text{Var}_{i,t}$ variable is constructed allows for a straightforward interpretation. It captures the realized capital gains on the two asset categories in t due to unconventional monetary policy. Referring again to the decomposition in Eq. (2.1) and assuming $Y_{i,t}$ is household i total portfolio, y is i' total inflow/outflow of assets, p

Figure 2.2: The biennial unconventional monetary policy measures



Note: The graph depicts the biennial cumulated impact of UMP announcements' unexpected component on Italian government bonds and risky asset categories. The blue and red lines are obtained by estimating at the daily frequency the effect of ECB's UMP announcements on the returns of several financial indexes (bot, btp, ctz, cct in the case of $UMP^{gov}_{average}$ and equity, corporate bond, liquidity funds, mixed funds, bond funds and equity funds in the case of $UMP^{risky}_{average}$), estimating Equation (2.3). The daily impact is then aggregated into a biennial series by summing within two year window, as explained in Section 2.4.1. For a complete list of the ECB's announcements included in the analysis, see Table 2.10.

the price of assets and r the total portfolio return, then

$$\begin{aligned}
\underbrace{Y_{i,t} - Y_{i,t-1}}_{\text{change in wealth}} &= \underbrace{y_{i,t}p_{i,t}}_{\text{active rebalancing}} + \underbrace{Y_{i,t-1}r_{i,t}}_{\text{capital gains}} \\
&= \underbrace{y_{i,t}p_{i,t}}_{\text{active rebalancing}} + \underbrace{Y_{i,t-1}r_{i,t}^{non-UMP}}_{\text{capital gains}^{non-UMP}} + \underbrace{Y_{i,t-1}r_{i,t}^{UMP}}_{\text{capital gains}^{UMP}} \\
&= y_{i,t}p_{i,t} + Y_{i,t-1}r_{i,t}^{non-UMP} + \\
&+ \underbrace{UMP_{t,average}^{gov} \times \text{gov bonds}_{i,t-1} + UMP_{t,average}^{risky} \times \text{risky}_{i,t-1}}_{\text{capital gains}^{UMP}}.
\end{aligned}$$

Thus, this Chapter estimates the UMP-induced financial wealth effect on household portfolio choices, i.e. how much of the exogenous change in the value of assets (windfall gains) due to UMPs is passed through to financial inflows or outflows. The two component of the UMPs variable, UMPs-exposure through government bonds and risky assets are included separately in the empirical model (see Section 2.4.2). First, households might react differently to different realized UMPs-capital gains, implying a different wealth effect for different asset categories. Second, including only one UMPs- capital gains, for example, the UMPs-exposure through the government bonds, I would impose *a priori* that the UMP transmission happens only through government bond exposure, and, in turn, that households only invested in risky assets would not be affected by the ECB monetary action.

2.4.2 Econometric framework and results

2.4.2.1 Unconventional monetary policy and portfolio rebalancing

This section addresses the empirical question of whether UMPs affects household portfolio rebalancing. Following Juster et al. (2006) and Calvet et al. (2009a), I employ the following univariate model:

$$\begin{aligned}
as_{i,t}^{(f)} = c^{(f)} + \alpha \text{HHs Controls}_{i,t-1}^{(f)} + \beta_1 UMP_{\text{t,average}}^{\text{gov}} \times \text{gov bonds}_{i,t-1}^{(f)} \\
+ \beta_2 UMP_{\text{t,average}}^{\text{risky}} \times \text{risky}_{i,t-1}^{(f)} + \delta_t + \eta_{i,t}^{(f)}
\end{aligned} \tag{2.4}$$

where i denotes the household, $t = 2006 - 2016$, $f = \{\text{risky, gov}\}$ asset categories; $j=$ bot index, btp index, ctz index, cct index if f=gov and $j=$ equity index, corporate bond index, liquidity funds index, mixed funds index, bond funds index, and equity funds index if f=risky; $as_{i,t}^{j}$ indicates the (stock of) active saving of category f between $t-1$ and t; the interactions $\text{UMP}_{\text{t, average}}^{\text{gov}} \times \text{gov bonds}_{i,t-1}$ and $\text{UMP}_{\text{t, average}}^{\text{risky}} \times \text{risky}_{i,t-1}$ capture the effect of unconventional monetary policy on the dependent variable. The vector HHs $\text{Controls}_{i,t-1}^{(f)}$ contains one period lagged household financial and demographic characteristics (so to mitigate the impact of reverse causality) that reflect factors likely to shape rebalancing decisions. The first category includes disposable income, net wealth, and a dummy equal to 1 if the household has a mortgage. The second category includes a dummy for post high-school education, dummies for the sex and marital status, as well as the age of the household head and the family size. HHs $\text{Controls}_{i,t-1}^{(f)}$ also contains the lag investment in government bonds and risky assets, the two variables included in the interaction terms and the capital $\text{gains}^{non-UMP}$, i.e. the total portfolio capital gains minus the portion that is to be attributed to UMPs. Moreover, I include dummies capturing the household head's attitude toward risk.[9] The effect of $UMP_{t,average}^{\text{risky}}$ and $UMP_{t,average}^{\text{gov}}$ are absorbed by the time fixed effect δ_t. The coefficients of interest are β_1 and β_2, the interaction of the UMPs impact measures with the households (idiosyncratic) exposure. Finally, all financial variables are trimmed at the bottom and top 1%.

Table 2.2 includes the baseline results for risky assets (Column 1) and government bonds (Column 2). Only some coefficients are reported. In general, a positive sign indicates an increase in investment (purchases), while a negative sign indicates a disinvestment (sales). Moreover, based on the way the variables $UMP_{t,average}^{gov}$ and $UMP_{t,average}^{risky}$ are constructed, the coefficients attached to the interactions are the effect of an accommodative unconventional monetary policy announcement that increases the return of, respectively, risky assets and government bonds by 1%. Column (1) and (2) show that unconventional monetary policy has a positive impact on the rebalancing of both asset classes. In particular, a positive UMP shock induces a positive investment in risky assets (6.3%) and government bonds (4.6%), for an increase in the total portfolio of roughly 11%. Considering the average investment in risky assets (41000 EUR), it implies a positive investment of 2600 EUR in risky assets and 1900 EUR in government bonds. The results can also be interpreted through the lens of the confidence channel of monetary policy: by reducing the amount of volatility in markets or uncertainty about both financial and real variables outlook, UMPs are able to boost financial investment in those segments that were mostly affected by the crisis. Table 2.2 also shows that the UMPs wealth effect happens only through risky assets. One possible explanation is that, given that the average investment in government bonds amounts to roughly

[9]In SHIW, households are asked, “In managing your financial investments, would you say you have preferences for investments that offers: 1) A very high returns, but with a high risk of losing part of the capital; 2) A good return, but also a fair degree of protection for the invested capital; 3) a fair returns with a good degree of protection for the invested capital; or 4) low returns with no risk of losing the invested capital." The answers are used to construct dummies capturing the household head's attitude toward risk.

27000 EUR, the gains induced by UMPs are too little to trigger a wealth effect. This result ia in line with households' inertia in portfolio rebalancing.

Table 2.2: UMP and active rebalancing - baseline results

	Risky assets (1)	Government bonds (2)
Risky, lag × $\text{UMP}^{\text{risky}}_{\text{average}}$	6.283***	4.599***
	(2.088)	(1.712)
Gov, lag × $\text{UMP}^{\text{gov}}_{\text{average}}$	-2.443	-1.782
	(2.724)	(2.567)
Risky, lag	-0.504***	0.0713*
	(0.0415)	(0.040)
Gov, lag	0.223***	-0.693***
	(0.068)	(0.050)
Δ Income	0.303***	0.0870
	(0.077)	(0.060)
Δ Net wealth	0.0672***	0.0418***
	(0.007)	(0.007)
Net wealth, lag	0.0182***	0.0111**
	(0.004)	(0.004)
Income, lag	0.139**	0.145***
	(0.067)	(0.046)
Constant	yes	yes
Demographics, lag	yes	yes
Risk aversion, lag	yes	yes
Time FE	yes	yes
Observations	3,093	3,023
R^2	0.338	0.370

Note: The table reports the estimates from Equation (2.4) including data from 2008 to 2016. The dependent variables are the stock (in euro) of active rebalancing of risky assets (Column 1) and of government bonds (Column 2). The variables $\text{UMP}^{\text{risky}}_{\text{average}}$ and $\text{UMP}^{\text{gov}}_{\text{average}}$ are constructed estimating Equation (2.3) with daily data and then following the procedure explained in Section 2.4.1 to construct a biennial series. Only the coefficients of interest and the households' financial controls are reported here. Standard errors (in parenthesis) are robust to heteroskedasticity. Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 2.11 in the Appendix (that is the extended version of Table 2.2 where all estimates are reported) also shows that for both asset categories the rebalancing is negatively affected by the initial investment in the same category and positively by the other one, implying that, in line with Calvet et al. (2009b), the initial investment has a quantitative effect on the active change and that, after a shock, the investment might gradually revert to a long-term mean and that the ratio between risky assets and government bonds is (partially) re-adjusted. In line with previous findings (Juster et al., 2006; Calvet et al., 2009a,b; Bilias et al., 2010), household financial characteristics are positively correlated with the portfolio rebalancing. In general, all other controls have the expected sign, but the capital gain variables deserve some deeper considerations. Following Calvet et al. (2009b) and Juster et al. (2006), the relationship between active and passive saving components is expected to be negative, but in this context the magnitude of the coefficient is mechanically inflated by the relationship between the two. Households' financial wealth is measured with error, as are passive and active savings. Measurement error introduces bias due to the method used to construct capital gains and active rebalancing. In fact, for a given change in wealth, a positive error in the passive saving necessarily lowers active rebalancing by an equal amount, artificially inducing a negative correlation between the two (Juster et al., 2006). This explains the strong effect of capital gains on risky asset rebalancing. Finally, it is worth noting that financial and demographic characteristics, including risk attitude, play a more important role in risky asset rebalancing decisions compared to government bonds, implying that for the latter category other factors might be partially affecting the active saving choices, like the traditional precautionary saving motive.[10]

One possible concern is that the outcome might differ for different groups of households. Thus, the analysis splits the sample into three groups: "poor" households, defined as those in the bottom 50% of the income distribution, "middle class" households, defined as the third quartile, and "rich" households, defined as the 25% richest households according to the income distribution. Results are reported in Table 2.3 for risky assets and in Table 2.4 for government bonds. For ease of comparability, the tables also report the baseline results. For both asset categories, the full sample results is driven by the behavior of the rich households, with the UMPs being significant only for HHs in the third group. This finding is consistent with the well-established literature finding that richer, better educated households (the so-called sophisticated households) better diversify their financial investments (Vissing-Jørgensen and Attanasio, 2003; Calvet et al., 2007) and rebalance more frequently (Vissing-Jorgensen, 2002; Campbell, 2006; Calvet et al., 2009b; Bilias et al., 2010) than other households. Finally, more sophisticated agents tend to be more aware of financial products (Guiso

[10]Unfortunately it is not possible to capture the precautionary saving motive due to data limitation.

and Jappelli, 2005), invest more aggressively, and make smaller mistakes (Calvet et al., 2007, 2009b). According to Table 2.3 and 2.4, UMPs induce a rebalance of the total portfolio of roughly 10% (6.6% and 3.16% towards, respectively, risky assets and government bonds). Considering that, for this group, the investment in risky assets ranges from 1000 EUR to 418000 EUR (with an average investment of 50000 EUR), it implies a rebalancing of total portfolio that ranges between 100 and 41000 EUR. As in the baseline results, the impact of UMPs happens only through the exposure to risky assets. Moreover, it appears that the effect of financial covariates on the rebalancing is similar across household groups.

Another possible concern is that the second and the third phase of unconventional tools might have different impacts on rebalancing. Unlike the previous phase, which sought to relieve financial and sovereign stress and fight redenomination risk and financial markets' geographical fragmentation triggered by crises, after Summer 2014 ECB action was designed to support economic growth and inflation. Specifically, in this third phase, the ECB introduced the Asset Purchase Programme (APP) for private and public sector securities with the aim of further depressing the term structure of interest rates, not only in those market segments where there is a direct intervention, namely covered bonds, asset-backed securities (ABS), sovereign and corporate bonds, but also into non-targeted markets through the portfolio rebalancing channel (Praet, 2017). Due to data limitations, it is not possible to directly compare the UMPs impact on rebalancing in the two different phases, but it is possible to re-run the analysis excluding data after 2014. Results for risky assets and government bond rebalancing are in Tables 2.12 and 2.13, respectively. The two tables show that excluding the last rounds of unconventional monetary policy increases rich households rebalancing of risky and government bonds by roughly 20% and 45% , respectively. This is further evidence on the effectiveness of ECB's policy in restoring markets' confidence and that, while the positive rebalancing in the period before the end of 2014 is driven by the confidence channel of monetary policy, after 2014 other channels, such as the portfolio rebalancing channel, might be attenuating (or even reversing) the sign of the rebalancing.

A final possible concern is that unconventional monetary policy might affect households' financial portfolio not only through the change in financial asset prices, but also through other direct and indirect effects. For example, UMP might impact households' saving incentives (the intertemporal substitution effect) by affecting interest rates on newly originated fixed rate mortgages or adjustable rate mortgages. In the former case, households might decide to invest these extra available resources in their financial investment, and this could have a positive impact on the portfolio adjustment of both the exposures to risky asset and goverment bonds. In the latter case, a newly originated

Table 2.3: UMP and risky assets active rebalancing - different household groups

	Full Sample	Group 1 1-50%	Group 2 51-75%	Group3 76-100%
	(1)	(2)	(3)	(4)
Risky, lag × $UMP^{risky}_{average}$	6.283***	1.739	5.929	6.619**
	(2.088)	(4.959)	(3.650)	(2.789)
Gov, lag × $UMP^{gov}_{average}$	-2.443	-0.962	-4.462	0.599
	(2.724)	(11.09)	(2.798)	(4.016)
Risky, lag	-0.504***	-0.312***	-0.490***	-0.516***
	(0.0415)	(0.103)	(0.0776)	(0.0512)
Gov, lag	0.223***	0.182*	0.182***	0.227**
	(0.0680)	(0.107)	(0.0659)	(0.103)
Δ Income	0.303***	0.348**	0.254*	0.324***
	(0.0775)	(0.171)	(0.146)	(0.0996)
Δ Net wealth	0.0672***	0.0679***	0.0763***	0.0601***
	(0.00700)	(0.0135)	(0.0116)	(0.00928)
Net wealth, lag	0.0182***	0.0180*	0.0131**	0.0180***
	(0.00435)	(0.00988)	(0.00540)	(0.00641)
Income , lag	0.139**	0.295	0.256	0.169
	(0.0686)	(0.246)	(0.226)	(0.130)
Constant	yes	yes	yes	yes
Demographics, lag	yes	yes	yes	yes
Risk aversion,lag	yes	yes	yes	yes
Time FE	yes	yes	yes	yes
Observations	3,093	538	1,184	1,371
R^2	0.338	0.211	0.367	0.355

Note: The table reports the estimates from Equation (2.4) including data from 2008 to 2016. In all four columns the dependent variable is the stock (in euro) of Italian risky assets active rebalancing. The first column shows the results when considering the full sample. In the following three columns households are split according to the value of their disposable income: Group 1 includes households in the bottom two quartiles of the income distribution, Group 2 contains households in the third quartile, and Group 3 comprises households in the top quartile. The variables $UMP^{risky}_{average}$ and $UMP^{gov}_{average}$ are constructed estimating Equation (2.3) with daily data and then following the procedure explained in Section 2.4.1 to construct a biennial series. Only the coefficients of interest and the households' financial controls are reported here. Standard errors (in parenthesis) are robust to heteroskedasticity. Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 2.4: UMP and government bonds active rebalancing - different household groups

	Full Sample	Group 1 1-50%	Group 2 51-75%	Group3 76-100%
	(1)	(2)	(3)	(4)
Risky, lag × $\text{UMP}^{\text{risky}}_{\text{average}}$	4.599***	11.36	6.371	3.164*
	(1.712)	(7.210)	(4.593)	(1.911)
Gov, lag × $\text{UMP}^{\text{gov}}_{\text{average}}$	-1.782	8.667	-3.431	-2.411
	(2.567)	(6.069)	(4.452)	(2.784)
Risky, lag	0.0713*	0.00677	0.120	0.0439
	(0.0398)	(0.0883)	(0.0836)	(0.0459)
Gov, lag	-0.693***	-0.662***	-0.674***	-0.719***
	(0.0502)	(0.0844)	(0.0923)	(0.0786)
Δ Income	0.0870	0.327***	0.134	-0.00699
	(0.0602)	(0.116)	(0.0875)	(0.0934)
Δ Net wealth	0.0418***	0.0309***	0.0385**	0.0441***
	(0.00719)	(0.00852)	(0.0158)	(0.00918)
Net wealth, lag	0.0111**	0.00599	0.0124	0.0101**
	(0.00434)	(0.00520)	(0.0102)	(0.00508)
Income, lag	0.145***	0.484***	0.202	0.128
	(0.0456)	(0.124)	(0.188)	(0.0971)
Constant	yes	yes	yes	yes
Demographics, lag	yes	yes	yes	yes
Risk aversion	yes	yes	yes	yes
Time FE	yes	yes	yes	yes
Observations	3,023	966	1,163	894
R^2	0.370	0.391	0.349	0.418

Note: The table reports the estimates from Equation (2.4) including data from 2008 to 2016. In all four columns the dependent variable is the stock (in euro) of Italian government bonds active rebalancing. The first column shows the results when considering the full sample. In the following three columns households are split according to the value of their disposable income: Group 1 includes households in the bottom two quartiles of the income distribution, Group 2 contains households in the third quartile, and Group 3 comprises households in the top quartile. The variables $\text{UMP}^{\text{risky}}_{\text{average}}$ and $\text{UMP}^{\text{gov}}_{\text{average}}$ are constructed estimating Equation (2.3) with daily data and then following the procedure explained in Section 2.4.1 to construct a biennial series.Only the coefficients of interest and the households' financial controls are reported here. Standard errors (in parenthesis) are robust to heteroskedasticity. Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

mortgage (for example for the purchases of a new home) might have a negative effect on portfolio adjustment, as households might decide to disinvest their financial wealth to finance the down payment. The indirect effects of UMP operate through the general equilibrium responses of prices and wages, hence of labour income and employment. With an accomodative monetary policy, the direct increase in households' expenditure and firms' investment will lead to an increase in output, sustainin employment and wages, which, in turn, might have a positive impact on financial portfolio adjustment. In order to control for these second-order effects of monetary policy I re-run the analysis for different subsamples. Results for risky assets and for government bonds can be found in Tables 2.14 and 2.15 (both in the Appendix), respectively. Both tables present three sets of results. In the first one (Column 1-4) I consider only households where the head is employed at time $t-1$ and t; in the second one (Column 5-8) I include households that have no mortgage at time $t-1$ and t; in the third one (Column 9-12) I take into account only household that are home owners both in $t-1$ and t. In each set of results I report results for all households grouped together (full sample) and for the “poor", “middle" and “rich" household groups. Comparing results in Tables 2.14 with the baseline outcome obtained when using all available observations (Table 2.3), it is possible to notice that results are very similar. Only the interaction term Risky, lag $\times$ $\text{UMP}^{\text{risky}}_{\text{t, average}}$ is significant, the coefficients are qualitatively and quantitatively very close and significant only for the “rich" households. This suggests that none of the possible second order effect of UMP taken into consideration play a role risky asset investment adjustment decisions. The same is true when comparing Tables 2.15 and 2.4. As in the baseline results, what drives governement bond investment decisiond for households in the three analysed subsamples is previous year investment in risky asset. Moreover, results are significant only for households at the top of the income distribution and the coefficients are qualitative similar to the baseline case, although quantitatively a bit higher. This suggests that for households where the head is always employed, that has no mortgage and that are home owners the positive effect of unconventional policies on goverment bonds invesment might be stronger. This last set of results confirm baseline findings.

2.4.2.2 Unconventional monetary policy and change in household's risk aversion

Results in previos section show that Italian households respond in a pro-cyclical way to UMP, increasing their invesment into both government bonds and risky assets. As discussed in Section 2.2, this behaviour is consistent with the *confidence channel* of monetary policy. To further explore the the relationship between UMP and households' risk aversion I make use of the measure of households' willingness to take financial risk reported in the SHIW. In each wave, respondents are asked, "In managing your financial investments, would you say you have preferences for investments that offers: 1) A very high returns, but with a high risk of losing part of the capital; 2) A good return, but also a fair degree of protection for the invested capital; 3) a fair returns with a good degree of protection for the invested capital; or 4) low returns with no risk of losing the invested capital". Responses are coded with integers from 1 to 4, with higher score indicating a higher aversion to risk. This qualitative measure of risk aversion is extensively used in the literature as a risk attitude measure (e.g., Malmendier and Nagel 2011 and Guiso et al. 2018). Then, I use this risk aversion measure to construct the variable "Change in risk aversion":

$$\text{Change in risk aversion}_{i,t} = \begin{cases} -1, & \text{if risk aversion decreases between } t-1 \text{ and } t \\ 0, & \text{if risk aversion does no change between } t-1 \text{ and } t \\ 1, & \text{if risk aversion increase between } t-1 \text{ and } t. \end{cases}$$

Figure 2.3 plots the distribution of the risk aversion and the change in risk aversion variables. Panel (a) shows a very skewed distribution of the risk aversion variable, with only 2% of the sample choosing "high return and high risk" and with 80% of the households choosing between "moderate return and moderate risk" and "low return and no risk". Panel (b) shows that between 2006 and 2016, almost 30% of the sample experiences an increase in risk aversion, in contrast to almost 20% that sees a decrease. The remaining 50% reports no change in its risk attitude.

I use the newly constructed variable to understand how UMP affects change in risk aversion. I model the comulative probability of these ordinal outcomes with an ordered probit model,

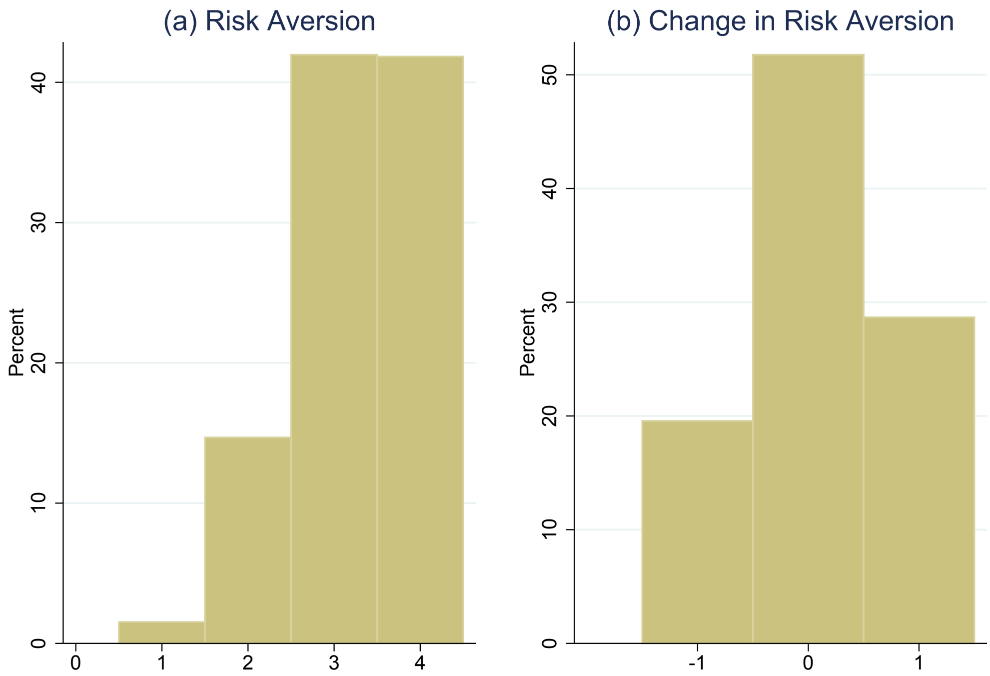
$$\begin{aligned} &\Pr(y_{i,t} \leqslant |x_{i,t-1}, UMP^{\text{risky}}_{\text{t,average}} \times \text{risky}_{i,t-1}, UMP^{\text{gov}}_{\text{t,average}} \times \text{gov bonds}_{i,t-1}) = \\ &\Phi(\alpha_j - \beta_1 UMP^{\text{risky}}_{\text{t,average}} \times \text{risky}_{i,t-1} - \beta_2 UMP^{\text{gov}}_{\text{t,average}} \times \text{gov bonds}_{i,t-1} - \gamma x_{i,t-1}), \quad (2.5) \end{aligned}$$

where $y_{i,t} \in \{-1, 0, 1\}$ denotes the change in risk aversion, $\Phi(.)$ is the cumulative standard distribution function, the α_j denote the cutoff points that must be estimated ($\alpha_1 < \alpha_2 < \alpha_3 < \infty$) and $x_{i,t-1}$ is a vector containing all control variables described in Section 3.4. I estimate the model with maximum likelihood to obtain estimates of β_1, β_2, γ and the cutoff points. The analysis is performed on the full sample and when eliminating households that reported inconsistent answers to the risk aversion measure/financial portfolio distribution.[11] Because coefficients do not have a direct

[11] I eliminate answers where households define themselves as risk lovers but only invest in goverment bonds and where households define themselves as highly risk averse but only invest in risky assets.

economic interpretation, I focus especially on the marginal effect of the increase of one unit in the UMP measures, leaving all other variables at their actual sample realization.

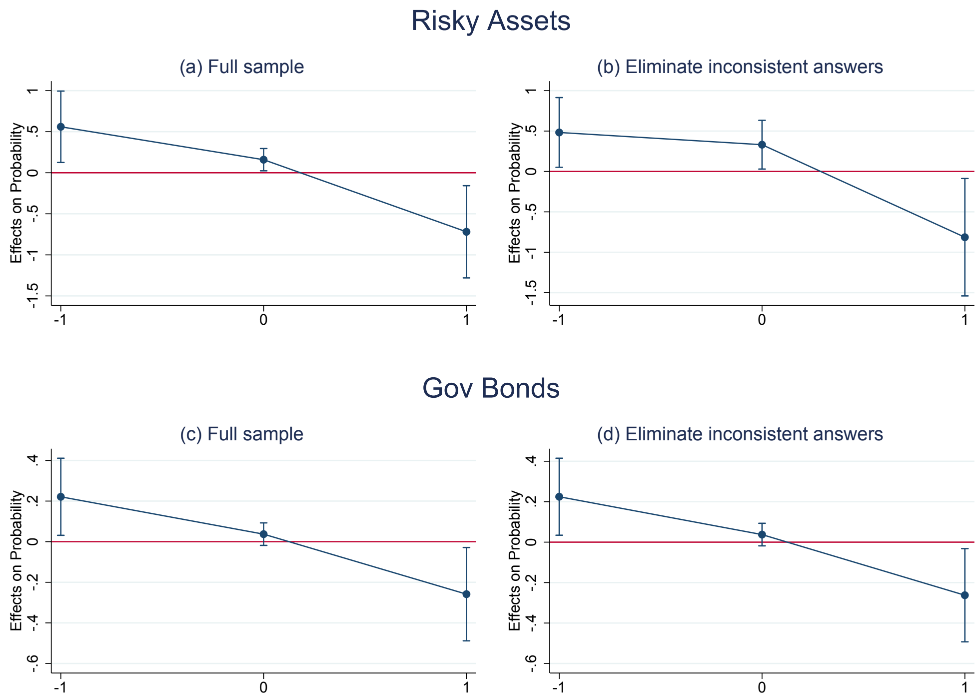
Figure 2.3: Italian households' risk aversion and change in risk aversion, 2006-2016



Note: The figure shows the distribution of of Italian households' level of risk aversion (left panel) and the change in risk aversion (right panel) between 2006 and 2016. In SHIW, households are asked, "In managing your financial investments, would you say you have preferences for investments that offers: 1) A very high returns, but with a high risk of losing part of the capital; 2) A good return, but also a fair degree of protection for the invested capital; 3) a fair returns with a good degree of protection for the invested capital; or 4) low returns with no risk of losing the invested capital". In order to construct the variable "risk aversion", responses are coded with integers from 1 to 4, with higher score indicating a higher aversion to risk. The variable "change in risk aversion" is equal to -1 if the household has experienced a decrease in risk aversion, is equal to 0 if household's risk aversion has remained unchanged and is equal to 1 if risk aversion has increased between $t-1$ and t.
Source: Bank of Italy's Survey on Household Income and Wealth, 2006-2016.

Table 2.16 (in the Appendix) presents the results of the ordered probit estimated on the 2006-2016 sample. The estimates of the parameters of interest are at the top of the table. Results are presented separately for two different household groups (investors in risky assets and investors in government bonds), when using the full sample (Column 1 and 3) and when dropping inconsistent answers (Column 2 and 4). The table shows that an accomodative UMP shock increases the probability of decreasing risk aversion. Moreover, consistent with all previous results, this effect happens only through risky assets. Figure 2.4 plots the marginal effect of a one unit increase of the variable the $UMP^{\text{risky}}_{\text{t,average}}$. The economic magnitudes are sizeble. For households investing in Italian risky assets (Panel a), an accomodative UMP shock, implies, on average a 72% lower probability of increasing their risk aversion, a 16% higher probability of not changing their risk aversion and a 56% increase in the probability in decreasing their risk aversion. When I drop inconsistent answers (Panel b), results are quantitatively and qualitatively similar. For households investing in Italian goverment bonds (Panel c), results are qualitatively similar although the point estimates appear to be smaller in magnitude. An accomodative UMP shock, implies, on average a 26% lower probability of increasing their risk aversion and a 22% increase in the probability of decreasing risk aversion. Thus, results show that households' risk aversion is negatively affected by accomodative monetary policy and that this result is stronger for investors in risky assets. These findings are not only consistent with the increase in financial investment discussed in the previous section, but are also in support of the confidence channel of monetary policy: by reducing risk perception and uncertainty, monetary policy lowers risk aversion, restores households' confidence and boosts financial investment in financial segments strongly hit by the crisis.

Figure 2.4: UMP marginal effects on Italian households' change in risk aversion, 2006-2016



Note: The figure plots the marginal effects of the variable $\text{UMP}^{\text{risky}}_{\text{average}}$ on the change in risk aversion. The variable "change in risk aversion" is equal to -1 if the household has experienced a decrease in risk aversion, is equal to 0 if household's risk aversion has remained unchanged and is equal to 1 if risk aversion has increased between $t-1$ and t. The marginal effects are computed from an ordered probit model estimated with maximum likelihood (Equation 3.4). The coefficients of interest are reported for the two different samples used for the baseline estimation (Risky assets and gov bonds, please refer to Table 2.2) and for the full sample (Panel a and c) and when eliminating inconsistent answers (Panel b and d). 90% confidence intervals.

2.5 Robustness checks

This section reports a series of robustness checks to the results. First, I employ a two-step model to correct for possible sample-selection bias. Second, I construct the UMPs impact measures using different aggregation and estimation techniques. Third, I include sampling weights to correct for survey data construction. Finally, I control for the quality of survey answers. All sensitivity tests report results quantitatively and qualitatively similar to the baseline findings.

2.5.1 Robustness checks to the model

In the portfolio allocation literature, it is well established that not all households invest in risky asset markets. Furthermore, this literature also finds that the decision whether to participate or not in risky markets and how much to invest are correlated. This, in turn, creates a problem of self-selection into investing that should be taken into account in the empirical analysis. The same issues may apply when considering the choice of rebalancing. In fact, the investment and the rebalancing decision problems are very similar; with the latter analyzing in terms of (financial investment) flows what the former analyzes in terms of (financial investment) levels. In presence of self-selection, the use of OLS would lead to inconsistent parameter estimates. For this reason, in this section I make use of a latent variable model with a probit selection equation (Guiso et al., 2000, 2002a,b; Guiso and Jappelli, 2005). I deal with the joint decision of whether to rebalance and how much to rebalance using a Heckman selection model (Heckman, 1979). I estimate a probit model for the binary choice of rebalancing conditioning on not being invested in the asset category in $t-1$ (extensive margin decision) and then a rebalancing equation for the participants accounting for selection between $t-1$ and t. Formally, the model is the following:

$$as_{it}^{(f)} = \beta_2 x_{2,i,t-1} + \rho_{12}\lambda(\hat{\beta}_1 x_{1,i,t-1}) + e_{2,i,t}^{(f)}$$

with $as_{it} \neq 0$ indicates the (stock of) active saving of the asset category $f = \{\text{risky}, \text{gov}\}$, $\lambda = \phi(\hat{\beta}_1 x_{1,i,t-1})/\Phi(\hat{\beta}_1 x_{1,i,t-1})$ is the inverse Mills ratio, ϕ is the normal density and $\hat{\beta}_1$ is obtained by estimating the first-stage probit model

$$\begin{aligned}\Pr(P_{it-1}^{(f)} = 1 | x_{1,i,t-1}) &= \Pr(\beta_1 x_{1,i,t-1} + e_{1,i,t-1}^{(f)} > 0)\\ &= \Phi(\hat{\beta}_1 x_{1,i,t-1})\end{aligned}$$

where $P^{(f)}_{it-1}$ is a dummy variable equal to 1 if HH i holds asset category f in $t-1$, 0 otherwise. The error terms are both normal, $e^{(f)}_{1,it-1} \sim \mathcal{N}(0,\,1)$ and $e^{(f)}_{2,it} \sim \mathcal{N}(0,\,\sigma)$, and $\rho_{12} = Cov(e_{1,t-1}, e_{2,t})$.

Thus, I employ the following two-step model:

$$\begin{aligned} as^{(f)}_{i,t} &= c^{(f)} + \alpha \text{HHs Controls}^{(f)}_{i,t-1} + \beta_1 UMP^{\text{gov}}_{t,average} \times \text{gov bonds}^{(f)}_{i,t-1} \\ &\quad + \beta_2 UMP^{\text{risky}}_{t,average} \times \text{risky}^{(f)}_{i,t-1} + \delta_t + \eta^{(f)}_{i,t} \\ P^{(f)}_{i,t-1} &= c^{(f)} + \alpha \text{HHs Controls}^{(f)}_{i,t-1} + +\gamma R^{(f)}_{i,t-1} + \delta_{t-1} + \mu^{(f)}_{i,t-1} \end{aligned} \tag{2.6}$$

The first equation is identical to the linear model in Equation (2.4) in Section 2.4.2.1. $P^{j}_{i,t}$ is a dummy equal to zero if the household has not rebalanced between $t-1$ and t conditional that it did not hold asset j in $t-1$ or 1 otherwise; $Z_{i,t-1}$ contains households' financial and demographic characteristics; $R_{i,t-1}$ is the vector that includes the exclusion restrictions needed to identify the model, i.e. variables that impact only the binary decision of rebalancing. I employ the use of on-line banking, the ownership of a brokerage account, and the fact that at least one member of the household works in the financial industry. A well established literature finds that entry costs, trading costs, and information costs have a primary role in explaining the (low) rate of risky asset ownership (see, among others, Vissing-Jorgensen (2002) and Bertaut (1998)). I argue that these three variables have a positive impact on the aforementioned costs, by lowering the fixed (brokerage account), trading (on-line banking), and information costs (one member of the household works in the financial industry).

Results are shown in Table 2.17. Column (1) and (3) report the second stage for the risky asset and government bonds rebalancing. The first stage regressions are reported in Column (2) and (4). The significance of the inverse Mills ratio indicates that for both asset categories, the first and the second stages are not independent. Nonetheless, comparing these estimates with the baseline regressions, it is possible to notice that they are quantitatively very similar, implying that the bias introduced by not taking into account the sample selection is negligible.

2.5.2 Robustness checks to the identification

The first robustness check that I perform to the identification regards the construction of the two UMP measures. As described in Section 2.4.1, I construct the monetary policy measures by aggregating the estimates of the β coefficients (Equation 2.3), disregarding their statistical significance. Thus, I repeat the procedure described in Section 2.4.1, but this time only including in the construction of the UMP measures only estimates that are significant at least at the 10% level. The two newly created UMP measures, $\text{UMP}^{\text{risky}}_{\text{t, only sign}}$ and $\text{UMP}^{\text{gov}}_{\text{t, only sign}}$ are plotted in the top left panel of Figure 2.11 (in the Appendix).

Secondly, to assess whether the results are driven by the model used to estimate the effects of ECB's unconventional announcements on the financial indexes, I employ a panel technique. Thus, the empirical model is as follows:

$$r^{(f)}_{j,t} = \alpha^{(f)}_j + \sum_{a=1}^{A} \beta_a D_{a,t} + \gamma\Delta \text{ Eonia}_t + \sum_{n=1}^{N} \delta_n z_{n,t} + \eta^{(f)}_{j,t} \tag{2.7}$$

where $r_{j,t}$ is the daily return in the financial index j at time t; f={risky, gov} and j= bot index, btp index, ctz index, cct index if f=gov and j= equity index, corporate bond index, liquidity funds index, mixed funds index, bond funds index, and equity funds index if f=risky. The two UMPs impact measures, $UMP^{risky}_{t,panel}$ and $UMP^{gov}_{t,panel}$ (top right panel of Figure 2.11) are then derived applying points 1) and 2) of the procedure described in Section 2.4.1.

Furthermore, a general limitation associated with event-study analyses is that the choice of the event window around the announcement is crucial. It involves a trade-off between keeping the interval narrow enough to make sure it only captures the impact of the monetary policy news and choosing a window wide enough to fully account the reaction of market participants. Thus, I test the robustness of the baseline results by extending the event window to two-day and, following Hachula et al. (2020), I employ the following univariate model

$$r_{j,t} = \alpha_j + \sum_{a=1}^{A} (\beta^1_a D_{a,t} + \beta^2_a D_{a-1,t}) + \gamma\Delta\text{Eonia}_t + \sum_{n=1}^{N} \delta_n z_{n,t} + \eta_{j,t} \tag{2.8}$$

where the dummy $D_{a-1,t}$ is equal to 1 if announcement a happens in $t-1$ and 0 otherwise. The coefficient of interest is $\beta_{a,j} = \beta^1_a + \beta^2_a$. The two UMP impact measures, $UMP^{risky}_{t,two-day}$ and $UMP^{gov}_{t,two-day}$ (bottom panel of Figure 2.11), are then derived applying the procedure described in Section 2.4.1.

Table 2.18 in the Appendix shows that results of all three robustness checks are quantitatively and qualitatively similar to the baseline outcome.

2.5.3 Robustness checks to the estimation sample

Finally, the sensitivity of the baseline results is tested across different estimation samples.[12] First, the analysis is repeated including sampling weights. Sampling weights denote the inverse of the probability that each observation is included in the sample,

[12]Sampling weights are provided for all SHIW waves directly by the Bank of Italy.

in order to correct for the sampling design, nonresponse or sample selection. In fact, if, according to the selection process, all the elements of the population have an unequal chance to be included in the sample, then not correcting for sampling weights might lead to biased estimates. On the other hand, using weights increase the sampling variance, especially in small samples (for an exhaustive discussion about this trade-off, see Solon et al. 2015). Second, I control for the quality of the survey data. For each household, the interviewer is asked to rate the reliability of the information on saving and financial investments provided by the respondent on a scale from zero to ten. All households with a score below eight are eliminated from the sample. Column (1) and (2) of Table 2.19 report the results when using the weighted sample, while Column (3) and (4) show results when cutting the sample. In both cases the findings are similar to the baseline results.

2.6 The impact of UMP-induced portfolio rebalancing on households' financial portfolios

In this Section, I simulate the impact of unconventional monetary policy on households' financial portfolios across the income distribution. In order to do so, I use the information on investment in government bonds and risky assets that refer to December 31, 2006 (Table 2.20).[13] This allows me to evaluate unconventional monetary policy effects across all years included in previous empirical analysis and to avoid the possibility that financial wealth distribution in later waives might already embed unconventional monetary policy effects. Then, I impose the valuation and rebalancing effect induced by unconventional monetary policy between 2006 and the end of 2016. Finally, I compare the value of the initial and the final portfolio, so to calculate the UMPs-induced financial return.

The empirical analysis in Section 2.4.2.1 finds that 1) only households in the top quartile rebalance their portfolio following unconventional stimulus (while the UMP's valuation the effect is common for all); and 2) the active rebalancing is expressed as a fraction of household's previous year investment in risky asset: an accomodative announcement that increases risky asset return by 1% leads to a $6.62\% * \text{risky}_{i,t-1}$ and to a $3.16\% * \text{risky}_{i,t-1}$ positive rebalancing of, respectively, risky assets and government bonds. Thus, the value of total the portfolio on December 31 2016 is, calculated as

[13] I assume that househols'd potfolio is composed only by government bonds and risky assets.

follows:

$$\text{Tot Port}_{2016,1} = \begin{cases} \sum_{t=2008}^{2016} \text{gov}_{t-1,i}(1+\text{UMP}_t^{gov}) + \text{risky}_{t-1,i}(1+\text{UMP}_t^{risky}), & \text{if Perc} \leq 75\% \\ \begin{array}{l} \displaystyle\sum_{t=2008}^{2016} (\text{gov}_{t-1,i} + 3.16\% * \text{risky}_{t-1,i})(1+\text{UMP}_t^{gov}) + \\ \text{risky}_{t-1,i}(1+6.62\%)(1+\text{UMP}_t^{risky}) \end{array}, & \text{if Perc} > 75\% \end{cases} \tag{2.9}$$

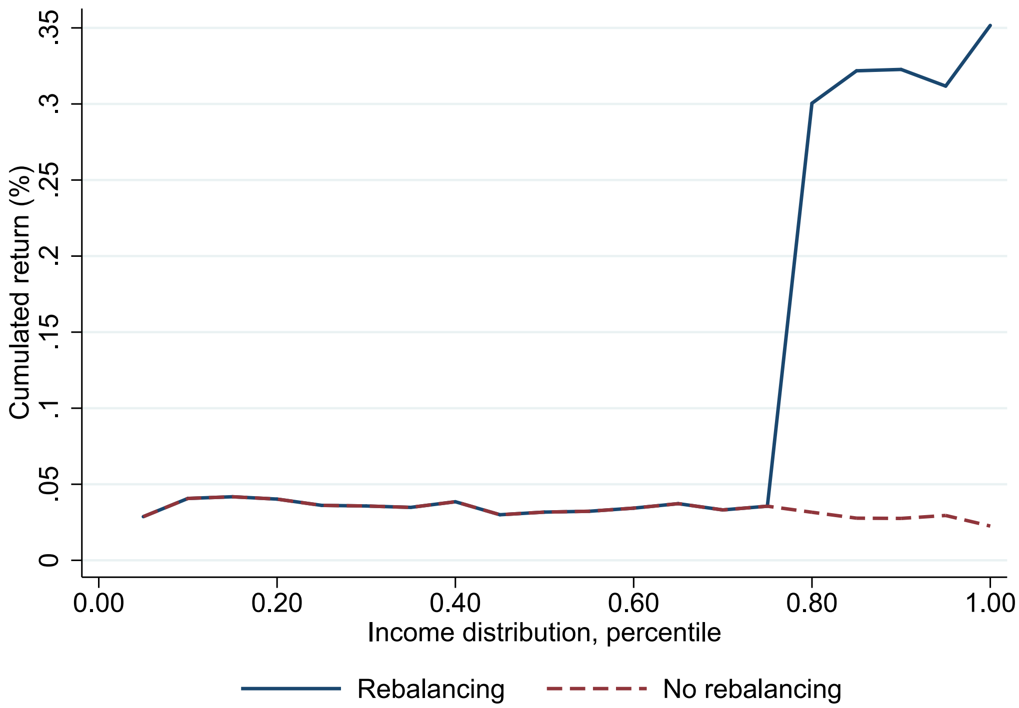
where $\text{Port}_{16,i}$ is the value of the total portfolio of family i at the end of 2016; $\text{gov}_{t-1,i}$ and $\text{risky}_{t-1,i}$ are the stocks of investment in government bonds and risky assets; UMP_t^{gov} and UMP_t^{risky} are the biennial vectors containing the UMPs-induced valuation effect estimated in Section 2.4.1; if household i belongs to the top quartile of the income distribution, the equation contains also the rebalancing effect ($3.16\% * \text{risky}_{t-1,i}$ for the government bonds and $6.62\% * \text{risky}_{t-1,i}$ for the risky assets). Finally, I calculate the portfolio return of household i over the 10 year period under analysis, $\text{Tot Port Ret}_{2016,i} = \frac{\text{Tot Port}_{2016,i} - \text{Tot Port}_{2006,i}}{\text{Tot Port}_{2006,i}}$. Thus, the total portfolio return is a function of the initial portfolio composition in terms of percentage investment in the two categories, valuation effect and rebalancing.

Figure 3.5 shows the average total portfolio return in 2016 for twenty different groups of households, ordered from left to right according to the quintile of income distribution they belong to. The graph contains two different scenarios. The solid blue line is the graphical representation of Equation (2.9), i.e. it depicts the portfolio rate of return when including the active portfolio rebalancing. The red dashed line is its counterfactual and considers the portfolio rate of return when the portfolio rebalancing component is shut down. By construction, the two lines differ only for households in the top quartile. From Figures 3.5, it emerges the portfolio return between the lower three quartiles and the upper ends of the income distribution varies considerably, but that this difference has to be attributed only to the rebalancing effect. Even more striking, the dashed red line shows that, without considering the rebalancing, the portfolio return for the top 5% households would be lower than the bottom 5%. This result is rooted in the very poor performance of the Italian stock market in the crisis years (left panel of Figure 2.7, red line) and in the ECB's inability in sustaining that specific asset class between 2006 and 2008. Figure 2.2 shows that the UMPs had a negative effect on risky assets in the first two years under analysis. This result has to be attributed entirely to the strong negative effects of unconventional tools announcements on the FTSE MIB index.[14] There is a positive correlation between percentage of the portfolio invested in equity and income (Table 2.20). This explains why rich households' portfolio have

[14]These results are not reported here, but are available upon request.

experienced stronger capital losses. On the contrary, including the portfolio rebalancing boosts portfolios' return, with the top quartile gaining between 30-35% of the initial value of their portfolio thanks to unconventional monetary policy valuation effects and their UMPs-induced investment decisions. In contrast, the bottom three quartiles increase the value of their portfolio by 3-4%. Thus, it is possible to conclude that households' heterogeneous portfolio rebalancing driven by UMPs is the major cause for financial wealth polarization among households.

Figure 2.5: The impact of UMP on Italian households' financial portfolio, 2006-2016



Note: The graph shows the average cumulated total portfolio rate of return between 2006 and 2016 for twenty different groups of households, ordered from left to right according to the quintile of income distribution they belong to. Two different scenarios are represented. The solid blue line depicts the cumulated portfolio rate of return when including both the UMP-induced valuation effect and active portfolio rebalancing. The red dashed line is its counterfactual and considers the cumulated portfolio return when the portfolio rebalancing component is shut down. By construction, the two lines only differ for households in the top quartile, as the empirical analysis finds that they are the only one rebalancing their financial portfolio following an accommodative unconventional monetary policy shock. For the initial value of total portfolios on 2006, see Table 2.20. The value of financial portfolios in 2016 is calculated using Equation (2.9). For each household i, the portfolio return is calculated as $\text{Tot Port Ret}_{2016,i} = \frac{\text{Tot Port}_{2016,i} - \text{Tot Port}_{2006,i}}{\text{Tot Port}_{2006,i}}$.

It is important to highlight some caveats of this analysis. This simulation aims at gauging the direct distributive implications of unconventional monetary policy on the basis of the *ex-ante* distribution of financial wealth, considering only some of the asset categories Italian households are invested in. Thus, it only measures what group of households gains/losses the most *ex-post* conditioning to the two asset categories under analysis, but it is mute with respect to the optimal portfolio rebalancing response to unconventional monetary policy. In principle, there could be a scenario where households in the bottom three quartiles show inertia in rebalancing their government bond or risky asset positions, but are active in other financial investments more heavily affected by UMPs. This could boost the value of their portfolio but it would not be captured by this analysis.

2.7 Conclusion

Unconventional monetary policy is expected to affect investors' portfolio choices, yet in the current academic debate on the impact of unconventional tools on households' wealth redistribution it is always assumed that households do not adjust their portfolios in response to monetary policy. This Chapter rejects this claim and shows that UMP does indeed matter for households' portfolio decisions, at least for households at the top of the income distribution.

To understand the impact of unconventional tools, I combine several waves of the Bank of Italy's Survey on Household Income and Wealth with financial indexes. This allows me to construct a novel dataset that contains granular information on the active portfolio rebalancing of a representative sample of Italian households from 2006 to 2016. Then, I identify monetary policy by isolating the unexpected change in households' financial wealth due to ECB's unconventional announcements. This allows me to estimate the UMP wealth effect on portfolio rebalancing.

The analysis shows that in a period of financial turbulence in which households have drastically reduced their financial investment, accommodative unconventional tools have induced a positive shift toward both Italian government bonds and Italian risky assets. These findings point toward the effectiveness of ECB's unconventional actions in restoring trust in the financial system and are consistent with the so-called confidence channel of monetary policy. Moreover, the analysis also shows that households' risk aversion is negatively affecte by accomodative UMP. Finally, I conduct a simulation exercise that evaluates the impact of unconventional tools on households' financial wealth considering both UMP-induced valuation and rebalancing effects. It appears that the difference in households' portfolio performance across the income distribution has to be attributed almost only to the rebalancing component, suggesting that rebalancing decision might play a prominent role in the redistributional effects of monetary policy.

These findings have important policy implications. Understanding how households are affected and heterogeneously respond to monetary policy changes can have important implications for the transmission mechanism of monetary stimulus to households and, through them, to the real economy. Moreover, although distributional politics are not one of the the objectives of central banks, uncovering the unintended and negative consequences of monetary intervention should be of primary interest, in order to understand where the risks are concentrated in the household sector.

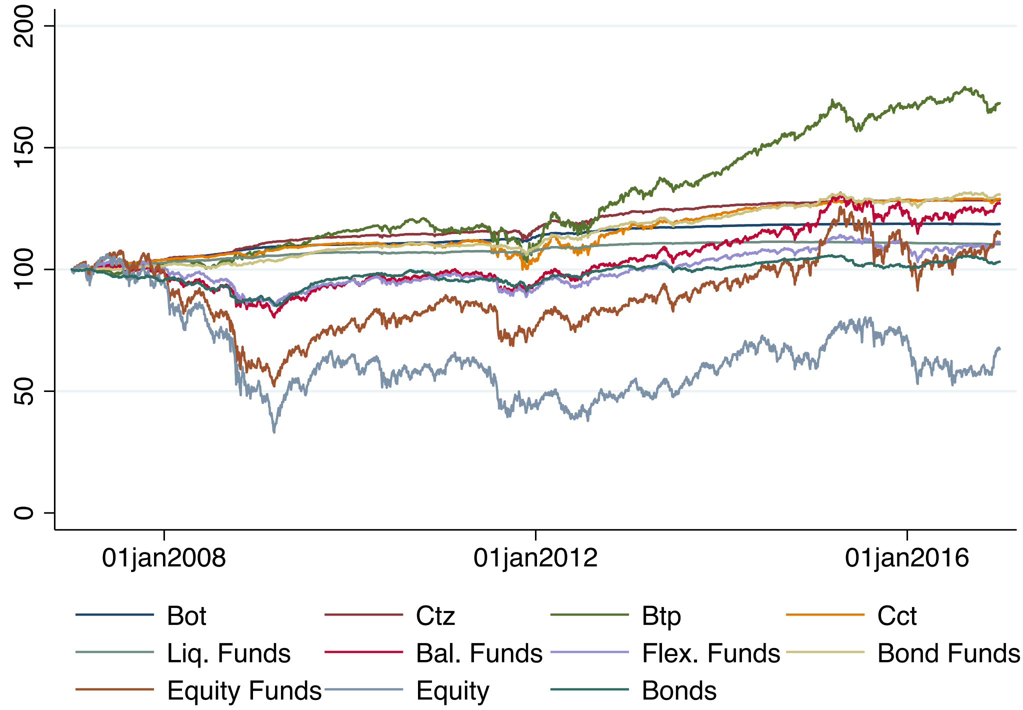
Finally, I acknowledge that this Chapter leaves some open issues. Due to data limitation, I am unable to uncover the role that households' behavioral traits play on the monetary policy/rebalancing decisions relationship. Moreover, comparing the impact of conventional and unconventional monetary policy on rebalancing choices could also be important. Addressing these questions is beyond the scope of this Chapter and represents an interesting and exciting avenue for future research.

2.A Figures

Figure 2.6: Italian households' portfolio composition, 2006-2016

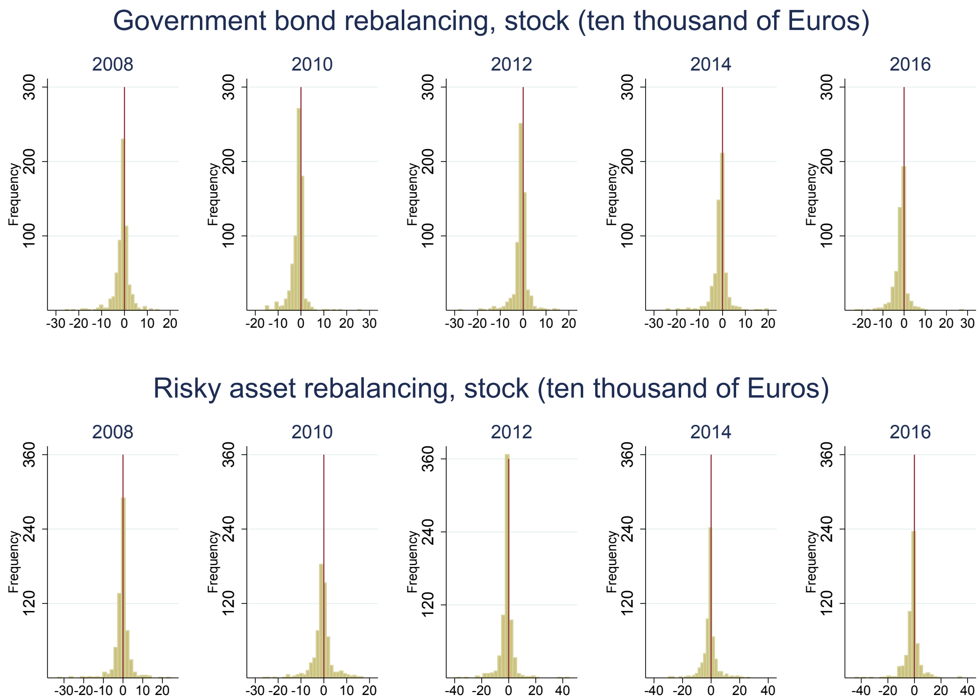
Note: The figure shows the evolution of Italian households' portfolio composition in the period 2006-2016, for three groups of households: Group 1 (bottom two quartiles of the income distribution), Group 2 (third quartile) and Group 3 (top quartile). Government bonds include: bot, btp, cct and ctz. Risky assets include: equity, corporate bonds, and mutual funds. Foreign assets include all assets issued by non-residents. Others include: postal bonds, certificates of deposits, repos, unlisted equity, and managed portfolio. Deposits are excluded. Averages are calculated using sample weights.
Source: Bank of Italy's Survey on Household Income and Wealth, 2006-2016.

Figure 2.7: Evolution of the financial indexes included in the analysis, 2007-2016



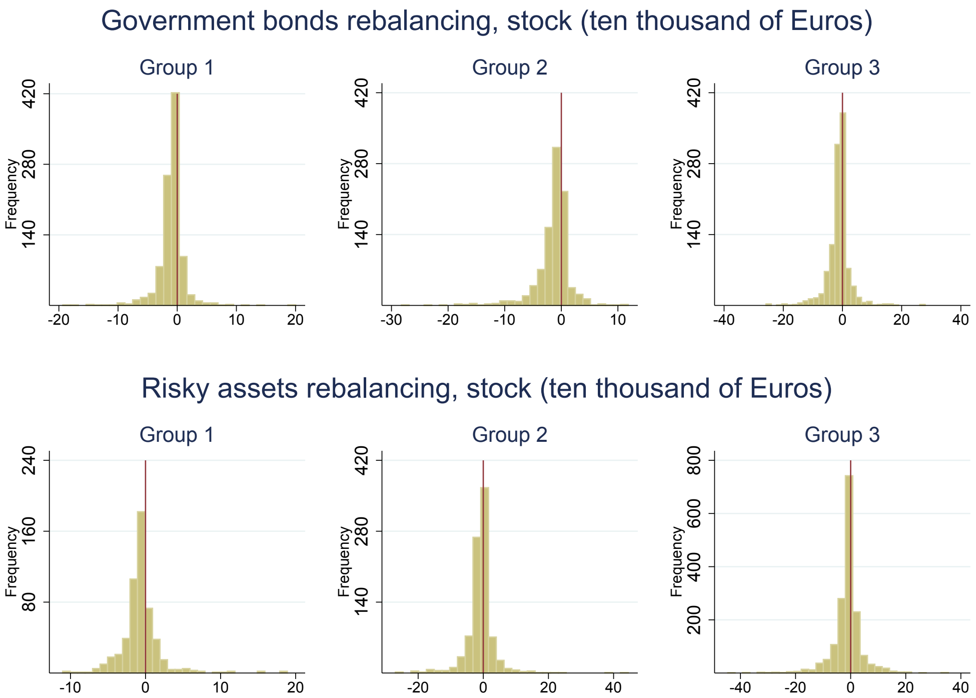
Note: The figure shows the eleven financial indexes used to approximate the return of the ten asset classes included in the analysis (Section 2.3.2) and to estimate the impact of ECB's UMP announcements on the Italian financial asset classes included in the empirical analysis (Section 2.4.1). All data are indexed to 100 on January 1, 2007.
Source: Bloomberg and Datastream.

Figure 2.8: Italian households' active rebalancing by year, 2006-2016



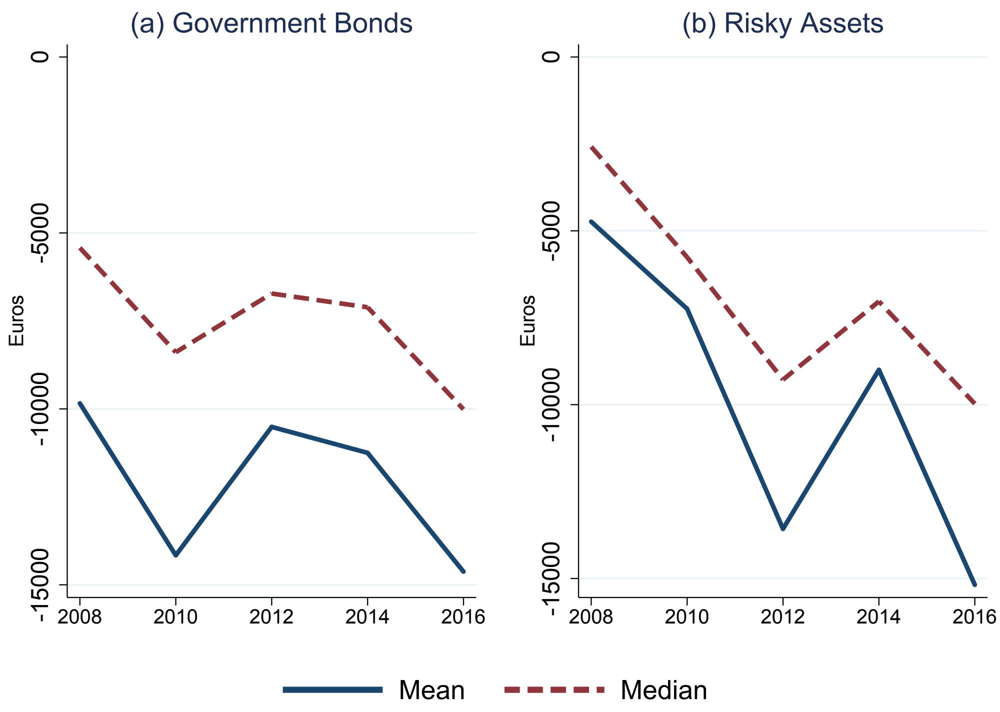
Note: The figure shows the distribution of Italian households' active rebalancing over the 2006-2016 period for the two asset categories included in the empirical analysis: Italian government bonds (bot, ctz, btp, cct) in the top panel; Italian risky assets (liquidity funds, mixed funds, bond funds, equity funds, equity and corporate bonds) in the bottom panel. For a detailed explanation of the construction of the two active rebalancing categories, please refer to Section 2.3.2. Only data included in the estimation sample is used for the calculations. Thus, rebalancing equal to zero as well as positive rebalancing in t conditioning on not being invested in the asset category in $t-1$ have been excluded from the table. The unit is ten thousand euro.
Source: Bank of Italy's Survey on Household Income and Wealth, Bloomberg, Datastream and own calculation.

Figure 2.9: Italian households' active rebalancing by household groups, 2008-2016



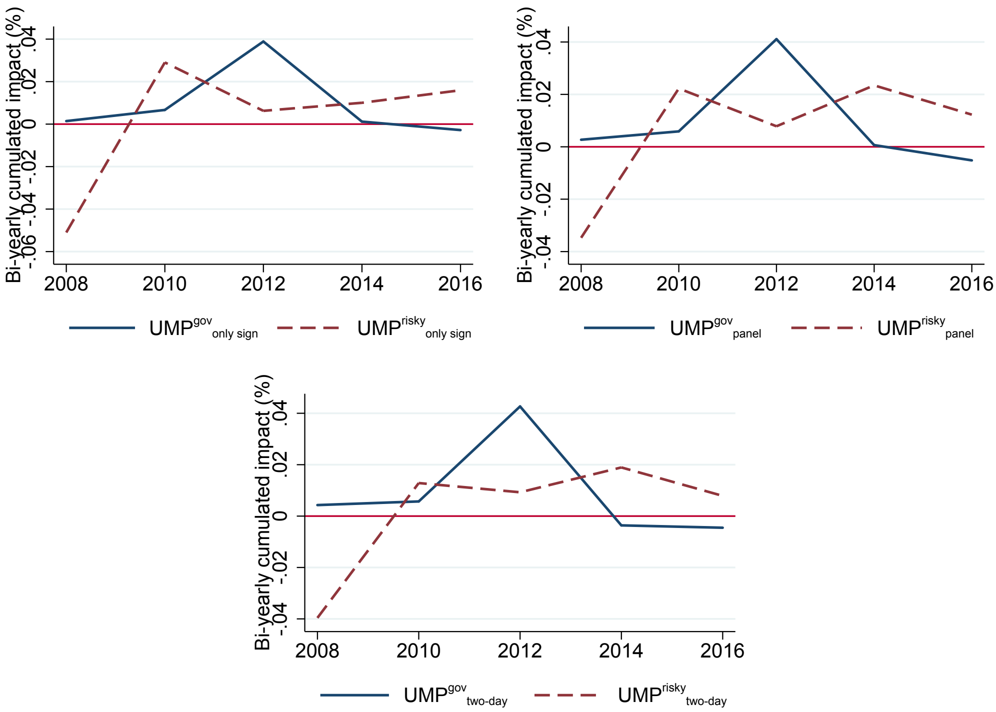
Note: The graph shows the distribution of Italian households' active rebalancing over the 2006-2016 period for the two asset categories included in the empirical analysis: Italian government bonds (bot, ctz, btp, cct) in the top panel; Italian risky assets (liquidity funds, mixed funds, bond funds, equity funds, equity and corporate bonds) in the bottom panel. Three groups of households are included in the graph: Group 1 (households in the bottom two quartiles of the income distribution), Group 2 (households in the third quartile) and Group 3 (households in the top quartile). For a detailed explanation of the construction of the two active rebalancing categories, please refer to Section 2.3.2. Only data included in the estimation sample is used for the calculations. Thus, rebalancing equal to zero as well as positive rebalancing in t conditioning on not being invested in the asset category in $t-1$ have been excluded from the table. The unit is ten thousand euro.
Source: Bank of Italy's Survey on Household Income and Wealth, Bloomberg, Datastream and own calculation.

Figure 2.10: Evolution of Italian households' active rebalancing, 2006-2016



Note: The figure shows the evolution over time of the mean and the median value of households' active rebalancing of the two categories included in the analysis, Italian government bonds (Bot, Ctz, Btp, Cct) and Italian risky assets (Liquidity funds, mixed funds, bond funds, equity funds, equity and corporate bonds) for the year 2006-2016. The unit is thousands of euro. For a detailed explanation of the construction of the two active rebalancing categories, please refer to Section 2.3.2. Only data included in the estimation sample is reported here. Thus, rebalancing equal to zero as well as positive rebalancing in t conditioning on not being invested in the asset category in $t-1$ have been excluded from the picture.
Source: Bank of Italy's Survey on Household Income and Wealth, Bloomberg, Datastream and own calculation.

Figure 2.11: Alternative biennial unconventional monetary policy measures



Note: The graph depicts the biennial cumulated impact of UMP announcements' unexpected component on government bonds and risky asset categories. The blue and red lines are obtained by estimating at the daily frequency the effect of ECB's UMP announcements on the returns of several financial indexes (bot, btp, ctz, cct in the case of UMP^{gov} and equity, corporate bond, liquidity funds, mixed funds, bond funds and equity funds in the case of UMP^{risky}), employing Equation (2.3) and then only adding up significant estmates in the top left panel, Equation (3.6) for the top panel and Equation (2.8) for the bottom panel. The daily impact is then aggregated into a biennial series by summing within two year window, as explained in Section 2.4.1. For a complete list of the ECB's announcements included in the analysis, see Table 2.10.

2.B Tables

Table 2.5: List of asset classes included in the SHIW, 2006-2016

2006	*2008*	*2010*	*2012*	*2014*	*2016*
Current account	Current account	Current account	Current account	Current account	Current account
Saving account	Saving account	Saving account	Saving account	Saving account	Saving account
Certif. of dep	Certif. of dep	Certif. of dep	Certif. of dep	Certif. of dep	Certif. of dep
Repos	Repos	Repos	Repos	Repos	Repos
PO savings certificates	PO savings certificates	PO savings certificates	PO savings certificates	PO savings certificates	PO savings certificates
BOTs (T-bills)	BOTs (T-bills)	BOTs (T-bills)	BOTs (T-bills)	BOTs (T-bills)	(T-bills)
CCTs (T-certificates)	CCTs (T-certificates)	CCTs (T-certificates)	CCTs (T-certificates)	CCTs (T-certificates)	CCTs (T-certificates)
.	.	.	Infl-indexed BTPs	Infl-indexed BTPs	Infl-indexed BTPs
BTPs (T-bonds)	BTPs (T-bonds)	BTPs (T-bonds)	BTPs (T-bonds)	BTPs (T-bonds)	BTPs (T-bonds)
CTZs (zero coupon)	CTZs (zero coupon)	CTZs (zero coupon)	CTZs (zero coupon)	CTZs (zero coupon)	CTZs (zero coupon)
Other Gov. Bonds	Other Gov. Bonds	Other Gov. Bonds	Other Other Gov. Bonds	Other Gov. Bonds	Other Gov. Bonds
Bonds	Bonds	.	.	.	.
.	.	Bonds iss. by Italian firms	Bonds iss. by Italian firms	Bonds iss. by Italian firms	Bonds iss. by Italian firms
.	.	Bonds iss. by Italian banks	Bonds iss. by Italian banks	Bonds iss. by Italian banks	Bonds iss. by Italian banks
Equity funds	Equity funds	Equity funds	Equity funds	Equity funds	Equity funds
Balanced equity funds	Balanced equity funds	.	.	.	.
Balanced bond funds	Balanced bond funds	.	.	.	
Balanced funds	Balanced funds	.	.	.	.
Bond funds	Bond funds	Bond funds	Bond funds	Bond funds	Bond funds
Money market funds	Money market funds	Money market funds	.	.	.
Flexible funds	Flexible funds	Flexible funds			
.	.	Flexible&balanced funds	Flexible&balanced funds	Flexible&balanced funds	Flexible&balanced funds
.	.	Non-harmonized funds	.	.	.
Indexed funds	Indexed funds	Indexed funds	.	.	.
.	.	.	Funds/ETFs in foreign cu.	Funds/ETFs in foreign cu.	Funds/ETFs in foreign cu.
Shares in listed c.	Shares in listed c.	Shares in listed c.	Shares in listed c.	Shares in listed c.	
- of which in privatized c.	- of which in privatized c.	.	.	.	.
Shares in unlisted c.	Shares in unlisted c.	Shares in unlisted c.	Shares in unlisted c.	Shares in unlisted c.	Shares in unlisted c.
Shares in private c.	Shares in private c.	Shares in private c.	Shares in private c.	Shares in private c.	
Shares in partnerships	Shares in partnerships	Shares in partnerships	Shares in partnerships	Shares in partnerships	Shares in partnerships
Managed portfolios	Managed portfolios	Managed portfolios	Managed portfolios	Managed portfolios	Managed portfolios
Bonds and inv funds	Bonds and inv funds	.	.	.	.
(foreign)	(foreign)	.	.	.	.
.	.	.	Gov bonds (foreign)	Gov bonds (foreign)	Gov bonds (foreign)
.	.	.	Bonds (foreign)	Bonds (foreign)	Bonds (foreign)
Shares (foreign)	Shares (foreign)	Shares (foreign)	Shares (foreign)	Shares (foreign)	Shares (foreign)
.	.	Funds (foreign)	.	.	.
Other (foreign)	Other (foreign)	Other (foreign)	Other (foreign)	Other (foreign)	Other (foreign)
Loans to coop.	Loans to coop.	Loans to coop.	Loans to coop.	Loans to coop.	Loans to coop.
			Other fin. assets	Other fin. assets	Other fin. assets

Note: The table shows all asset classes included in the SHIW in the years 2006-2016. In case the asset class is not included in the survey, the symbol . is used.

Source: Survey on Households Income and Wealth, Banca d'Italia,

Table 2.6: Description of the asset classes included in the analysis

Asset classes	Description	Classification
Bot	Government bills up to 1 year	Gov bonds
Ctz	Government bills up to 2 year	Gov bonds
Btp	Government bonds	Gov bonds
Btpi	Inflation linked bonds	Gov bonds
Cct	Government floating rate notes	Gov bonds
Bonds issued by Italian banks	Bank bonds are bonds that are issued by banks. As with any type of bond, bank bonds are a debt instrument	Corp bonds
Bonds issued by Italian firms	Firms bonds that are issued by firms. As with any type of bond, firms bonds are a debt instrument	Corp bonds
Shares in listed Italian companies	Stocks of publicly-traded Italian companies traded on the Milan Stock Exchange	Equities
Italian Liquidity Fund	A liquidity fund portfolio is comprised of short-term, or less than one year, securities representing high-quality, liquid debt and monetary instruments	Funds
Italian Balanced Funds	A balanced fund is a mutual fund that generally keeps to a 50-50 mix of stock and bond investments	Funds
Italian Flexible Funds	Mutual fund that allows capital to be invested as the financial professional sees fit. Flexible mutual funds do not have any restrictions on where the money is to be invested or how much money is allowed to be used	Funds
Italian Balance Equity Funds	Balanced fund with a higher percentage of equity	Funds
Italian Balanced Bond Funds	Balanced fund with a higher percentage of bonds	Funds
Italian Bond Funds	A bond fund is a fund invested primarily in bonds and other debt instruments. The exact type of debt the fund invests in will depend on its focus, but investments may include government, corporate, municipal and convertible bonds, along with other debt securities like mortgage-backed securities	Funds
Italian Equity Funds	Fund that invest primarily in stocks represent the largest category of mutual funds. Generally, the investment objective of this class of funds is long-term capital growth	Funds
Italian Mixed Funds	Given by the sum of Italian Balanced Funds, Italian Balanced Bond Funds, Italian Balance Equity Funds, Italian Flexible Funds	Funds

Note: The table lists all asset classes included in the analysis (Column 1), together with a brief description (Column 2) and their final classification (Column 3).

Table 2.7: Italian households' average holding of financial assets

Assets classes	2006	2008	2010	2012	2014	2016
Included in the analysis						
Bot[a]	0.064	0.083	0.071	0.045	0.048	0.047
Btp[b]	0.019	0.019	0.016	0.026	0.022	0.018
Cct[c]	0.021	0.018	0.017	0.010	0.017	0.013
Ctz[d]	0.002	0.000	0.003	0.000	0.001	0.000
Equities	0.066	0.060	0.041	0.046	0.037	0.033
Corporate bonds[e]	0.064	0.080	0.071	0.086	0.078	0.052
Mutual funds	0.072	0.061	0.072	0.057	0.059	0.064
Excluded from the analysis						
Deposits[f]	0.894	0.898	0.916	0.942	0.930	0.938
Other goverment papers	0.005	0.002	0.005	0.003	0.005	0.003
Postal bonds	0.071	0.064	0.058	0.064	0.050	0.049
Certificates of deposits	0.015	0.021	0.018	0.015	0.018	0.026
Repos	0.005	0.015	0.010	0.010	0.012	0.014
Other mutual funds	0.035	0.029	0.005	0.004	0.002	0.003
Assets issued by non-nationals	0.010	0.010	0.010	0.009	0.006	0.007
Unlisted equity	0.008	0.010	0.010	0.008	0.008	0.008
Managed portfolios	0.013	0.011	0.012	0.021	0.012	0.010
Loans to cooperatives	0.019	0.019	0.018	0.015	0.012	0.011

Note: The table includes the percentage holding of all financial asset classes contained in the Bank of Italy's SHIW from 2006 to 2016. The top panel reports the asset classes included in the analysis; the excluded ones are reported in the bottom panel. Averages are calculated using sample weights.
Source: Bank of Italy's Survey on Household Income and Wealth, 2006-2016.
[a] Treasury bills up to one year maturity.
[b] Long-term government bond.
[c] Floating-rate Treasury credit certificates, 2-4 years in maturity indexed to BOT.
[d] Treasury bills up to two year maturity.
[e] Bonds issued by Italian banks and firms.
[f] Current accounts, saving accounts, post office current accounts, post office saving accounts.

Table 2.8: Description of the financial indexes included in the analysis

Asset classes	Index	Index description
Bot	FTSE MTS Italy BOT Ex-Bank of Italy	It measures the performance of short-term Italian government debt securities, or BOTs. The FTSE MTS Ex-Bank of Italy BOT Index includes all the BOTs listed on MTS
Btp/Btpi	FTSE MTS Italy BTP Ex-Bank of Italy	It measures the performance of short-term Italian government debt securities, or BOTs. The FTSE MTS Ex-Bank of Italy BTP Index includes all the BTPs listed on MTS
Ctz	FTSE MTS Italy CTZ Ex Bank of Italy Index	It measures the performance of short-term Italian government debt securities, or BOTs. The FTSE MTS Ex-Bank of Italy CTZ Index includes all the CTZs listed on MTS
Cct	FTSE MTS Italy CCT Ex-Bank of Italy	It measures the performance of short-term Italian government debt securities, or BOTs. The FTSE MTS Ex-Bank of Italy CCT Index includes all the CCTs listed on MTS
Bonds	Italian constituents of BofA Merril Lynch Euro Corporate Index	It tracks the performance of EUR denominated investment grade corporate debt publicly issued in the eurobond or Euro member domestic markets
Equity	FTSE MIB Gross Total Return	It is the primary benchmark index for the Italian equity market. It captures approximately 80% of the domestic market capitalization and it measures the performance of the 40 most liquid and capitalized Italian shares
Liquidity Fund	Banca Fideuram Indice Fondi di Mercato Monetario	It measures the performance of all Italian liquidity funds. The index is calculated as the weighted average of the daily net asset value of each included fund
Balanced Funds	Italy Fideuram Balanced Group	It measures the performance of all Italian balanced funds. The index is calculated as the weighted average of the daily net asset value of each included fund
Flexible Fund	Italy Fideuram Flexible	It measures the performance of all Italian flexible funds. The index is calculated as the weighted average of the daily net asset value of each included fund
Bond Fund	Italy Fideuram Bond Funds	It measures the performance of all Italian bond funds. The index is calculated as the weighted average of the daily net asset value of each included fund
Equity Fund	Italy Fideuram Equity Funds	It measures the performance of all Italian equity funds. The index is calculated as the weighted average of the daily net asset value of each included fund

Note: The table reports the name and the description of the financial indexes used to approximate the asset classes included in the empirical analysis. The first column lists the asset classes; the second column contains the name and the provider of the indexes; the third column provides a brief description of the indexes.
Source: Bloomberg and Datastream.

Table 2.9: List of the macroeconomic news included in the analysis

Euro area	EC Bus. Climate Ind.; Current Account Net WDA SA; EC Cons. Conf. Ind; CPI YoY; CPI MoM; BOP CA Net NSA; New Orders (Manu.); YoY GFCF QoQ; EC Serv. Conf. Ind.; Markit Comp. PMI SA; Markit Serv. PMI SA; Retail Sales Vol. YoY WDA; Retail Sales Vol. MoM SA; ZEW Exp. of Econ. Growth; Trade Bal. with non EZ; M3 Money Supply 3 M. MA; PPI Industry Ex Constr.YoY; PPI Industry Ex Constr. MoM; Unem. Rate; GDP SA QoQ (real SA)
Italy	CPI NIC Incl Tbc. YoY; NSA CPI NIC Incl Tbc. MoM NSA; Cons. Conf. Ind. SA; Bus. Conf. Manu. Sector; Hourly Wages MoM SA; Ind. Orders YoY NSA; Ind. Orders MoM SA; Ind. Prod. YoY WDA; Ind. Prod. MoM SA; Ind. Prod. YoY; Ind. Sales YoY; Ind. Sales MoM SA; Manu. PMI SA; Serv. PMI SA; PPI Manu. MoM; PPI Manu. YoY; PPI Manu. YoY; Priv. Cons. QoQ SA WDA; Retail Sales MoM SA; Retail Sales YoY; Trade Balance Total; Unem. Rate SA; Real GDP YoY SA WDA; Trade Balance Non EU NSA
US	C PI YoY NSA; CPI MoM SA; CPI Ex. Fd. & En. YoY NSA; UM Cons. Conf. Ind; Pers. Cons. Exp. CPI YoY SA; Gov. Budget Balance; Cons. Spend. GR MoM SA; Core PPI; Housing Starts/Permits; PPI Fin. Goods SA; MoM% Avg. H Earnings YoY% SA; Dur. Goods Orders MoM SA; Markit Manu. PMI SA; PPI - Fin. Goods; Diff. between Exp. and Imp.; Cap. Util.n % of Tot. Cap.; Avg. H Earnings MoM% SA; CB Leading Ind. MoM; Ind. Prod. MoM SA; In. Jobless Claims SA; GDP QoQ SAAR; Bus. Inventories MoM SA; Constr. Spend. MoM SA; Production Nonfarm QoQ SA

Note: The table lists all economic data releases included in Equation (2.3).
Source: Bloomberg

Table 2.10: ECB's unconventional monetary policy announcements

22.08.2007	Supplementary liquidity-providing longer-term ref oper (LTRO) with a maturity of 3 m
28.03.2008	LTROs with a maturity of six months
29.09.2008	Special term refinancing operation
08.10.2008	Fixed rate tender procedure with full allotment on the main refinancing operation(MROs)
15.10.2008	List of assets eligible as collateral in Eurosystem credit operations extended
07.05.2009	LTROs with a maturity of one year
04.06.2009	Details on Purchase program for covered bonds (CBPP)
03.12.2009	Phasing out of 6-month LTROs, indexation of new one year LTROs
04.03.2010	Phasing out of 3-month LTROs, indexation of six month LTROs
10.05.2010	Securities Markets Program (SMP)
28.07.2010	Risk control measures in collateral framework reviewed
03.03.2011	Further LTROs
09.06.2011	MROs as fixed rate tender procedures with full allotment,at least until October 2011
04.08.2011	Further LTROs with a maturity of three and six months
08.08.2011	ECB will actively implement its Securities Market Program
06.10.2011	New covered bond purchase program (CBPP2)
08.12.2011	Two additional LTROs with a maturity of three years
21.12.2011	Results of first three year LTRO
09.02.2012	ECB's Governing Council approves eligibility criteria for additional credit claims
28.02.2012	Results of second three year LTRO
06.06.2012	FRFA on MROs as long as necessary, and at least until January 2013
26.07.2012	'Whatever it takes...' speech by ECB President Mario Draghi in London
02.08.2012	Outright Monetary Transactions program (OMT)
06.09.2012	Technical features of OMT
06.12.2012	FRFA on MROs as long as necessary, and at least until July 2013
22.03.2013	Collateral rule changes for some uncovered government guaranteed bank bonds
02.05.2013	FRFA on MROs as long as necessary, and at least until July 2014
04.07.2013	Gov Council expects the key ECB interest rates to remain at present or lower levels for an extended period of time (open-ended forward guidance)
08.11.2013	FRFA on MROs as long as necessary, and at least until July 2015
05.06.2014	Targeted longer-term refinancing operations (TLTROs)
03.07.2014	Details on TLTROs published
04.09.2014	Deposit rate -0.2
22.01.2015	Announcement of expanded asset purchase programme (APP)
16.07.2015	Reaffirmation that purchases are intended to run until end of September 2016
31.08.2015	New category of assets added as eligible collateral
03.09.2015	Increase in PSPP issue share limit
23.09.2015	Eurosystem adjust purchase process in ABS programme
22.10.2015	Questions on requirements for APP extension answered
09.11.2015	Increase in PSPP issue share limit enlarges purchasable universe
03.12.2015	APP extended until March 2017, deposit rate -0.3
21.01.2016	Review and possibly reconsider monetary policy stance at next meeting
10.03.2016	New targeted longer-term refinancing operations (TLTRO II), APP expanded, corporate bonds added to APP, deposit rate -0.4
21.04.2016	Details on implementation of APP expansion
03.05.2016	Legal acts relating to TLTRO II is published
02.06.2016	Details on corporate sector purchase programme (CSPP) published
21.07.2016	Confirmation that APP at 80 billion per month to run at least until March 2017
08.09.2016	Council meeting confirming continuation of APP
05.10.2016	Changes to collateral eligibility criteria and risk control measures for unsecured bank bonds
20.10.2016	Council meeting confirming continuation of APP

Note: The table lists all ECB unconventional announcements included in Equation (2.3)

Table 2.11: UMP and active rebalancing - baseline results, extended table

	Risky assets (1)	Government bonds (2)
Capital gains	-1.352***	-0.310
	(0.307)	(0.299)
Risky, lag × $UMP^{average}_{risky}$	6.283***	4.599***
	(2.088)	(1.712)
Gov, lag × $UMP^{average}_{gov\ bond}$	-2.443	-1.782
	(2.724)	(2.567)
Risky, lag	-0.504***	0.0713*
	(0.0415)	(0.0398)
Gov, lag	0.223***	-0.693***
	(0.0680)	(0.0502)
Δ Income	0.303***	0.0870
	(0.0775)	(0.0602)
Δ Net wealth	0.0672***	0.0418***
	(0.00700)	(0.00719)
Net wealth	0.0182***	0.0111**
	(0.00435)	(0.00434)
Income	0.139**	0.145***
	(0.0686)	(0.0456)
Age	91.86	32.91
	(72.09)	(40.49)
Married	-5023.1*	3100.3
	(2747.8)	(2290.3)
Divorced	180.9	-4756.7***
	(3340.6)	(1565.1)
Famiy size	-2964.9***	-2390.5***
	(794.3)	(544.9)
College+	6864.0***	214.1
	(2477.4)	(1876.3)
Mortgage	-3344.3*	-1586.2
	(1940.0)	(1677.9)
Woman	-317.0	-1108.3
	(1555.2)	(1120.4)
High ret/high risk	18967.5*	-13023.6***
	(10600.8)	(3521.0)
Good ret/fair risk	4689.7**	-2934.5
	(2380.7)	(1803.1)
Fair ret/low risk	3488.9**	-1125.9
	(1547.5)	(1200.7)
Constant	yes	yes
Time FE	yes	yes
Observations	3,093	3,023
R^2	0.338	0.370

Note: The table reports the estimates from Equation (2.4) including data from 2008 to 2016. The dependent variables are the stock (in euro) of active rebalancing of risky assets (Column 1) and of government bonds (Column 2). The variables $UMP^{risky}_{average}$ and $UMP^{gov}_{average}$ are constructed estimating Equation (2.3) with daily data and then following the procedure explained in Section 2.4.1 to construct a biennial series. Standard errors (in parenthesis) are robust to heteroskedasticity. Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 2.12: UMP and risky assets active rebalancing - excluding 2016

	Full Sample	Group 1 1-50%	Group 2 51-75%	Group3 76-100%
	(1)	(2)	(3)	(4)
Risky, lag × $UMP^{risky}_{average}$	6.876***	0.473	5.424	7.843***
	(2.280)	(5.399)	(5.574)	(2.635)
Gov, lag × $UMP^{gov}_{average}$	-1.056	4.540	-3.966	2.690
	(2.601)	(10.07)	(3.602)	(3.461)
Risky, lag	-0.473***	-0.363***	-0.500***	-0.464***
	(0.0490)	(0.117)	(0.116)	(0.0568)
Gov, lag	0.168***	0.109	0.131	0.162**
	(0.0652)	(0.117)	(0.107)	(0.0820)
Δ Income	0.308***	0.315*	0.300	0.330***
	(0.0844)	(0.167)	(0.194)	(0.103)
Δ Net wealth	0.0668***	0.0546***	0.0870***	0.0593***
	(0.00751)	(0.0125)	(0.0156)	(0.00914)
Net wealth, lag	0.0197***	0.0216**	0.0136*	0.0194***
	(0.00424)	(0.0104)	(0.00729)	(0.00536)
Income, lag	0.120*	0.201	0.0697	0.0846
	(0.0727)	(0.256)	(0.383)	(0.116)
Constant	yes	yes	yes	yes
Demographics, lag	yes	yes	yes	yes
Risk aversion	yes	yes	yes	yes
Time FE	yes	yes	yes	yes
Observations	2,566	446	755	1365
R^2	0.327	0.220	0.358	0.346

Note: The table reports the estimates from Equation (2.4) including data from 2008 to 2014. In all four columns the dependent variable is the stock (in euro) of Italian risky assets active rebalancing. The first column shows the results when considering the full sample. In the following three columns households are split according to the value of their disposable income: Group 1 includes households in the bottom two quartiles of the income distribution, Group 2 contains households in the third quartile, and Group 3 comprises households in the top quartile. The variables $UMP^{risky}_{average}$ and $UMP^{gov}_{average}$ are constructed estimating Equation (2.3) with daily data and then following the procedure explained in Section 2.4.1 to construct a biennial series. Only the coefficients of interest and the households' financial controls are reported here. Standard errors (in parenthesis) are robust to heteroskedasticity. Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 2.13: UMP and government bonds active rebalancing - excluding 2016

	Full Sample	Group 1 1-50%	Group 2 51-75%	Group3 76-100%
	(1)	(2)	(3)	(4)
Risky, lag × $UMP^{risky}_{average}$	5.093***	11.91	-1.915	5.510**
	(1.928)	(7.785)	(4.692)	(2.234)
Gov, lag × $UMP^{gov}_{average}$	-3.040	7.206	-4.099	-4.728
	(2.813)	(7.144)	(5.262)	(3.078)
Risky, lag	0.0898*	-0.0228	0.0245	0.0950*
	(0.0470)	(0.100)	(0.0887)	(0.0551)
Gov, lag	-0.648***	-0.619***	-0.669***	-0.640***
	(0.0574)	(0.118)	(0.0897)	(0.0932)
Δ Income	0.122*	0.423***	0.224**	0.0358
	(0.0666)	(0.133)	(0.0966)	(0.0939)
Δ Net wealth	0.0380***	0.0244***	0.0265***	0.0437***
	(0.00556)	(0.00845)	(0.00803)	(0.00851)
Net wealth , lag	0.00750**	0.00173	-0.000383	0.0107**
	(0.00341)	(0.00544)	(0.00507)	(0.00497)
Income , lag	0.153***	0.566***	-0.133	0.0876
	(0.0492)	(0.146)	(0.192)	(0.0943)
Constant	yes	yes	yes	yes
Demographics, lag	yes	yes	yes	yes
Risk aversion	yes	yes	yes	yes
Time FE	yes	yes	yes	yes
Observations	2,533	805	789	939
R^2	0.433	0.342	0.563	0.417

Note: The table reports the estimates from Equation (2.4) including data from 2008 to 2014. In all four columns the dependent variable is the stock (in euro) of Italian government bonds active rebalancing. The first column shows the results when considering the full sample. In the following three columns households are split according to the value of their disposable income: Group 1 includes households in the bottom two quartiles of the income distribution, Group 2 contains households in the third quartile, and Group 3 comprises households in the top quartile. The variables $UMP^{risky}_{average}$ and $UMP^{gov}_{average}$ are constructed estimating Equation (2.3) with daily data and then following the procedure explained in Section 2.4.1 to construct a biennial series. Only the coefficients of interest and the households' financial controls are reported here. Standard errors (in parenthesis) are robust to heteroskedasticity. Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 2.14: UMP and risky assets active rebalancing - controlling for UMP second order effects

	Always Employed				No Mortgage				Home Owner			
	Full Sample	Group 1 1-50%	Group 2 51-75%	Group 3 76-100%	Full Sample	Group 1 1-50%	Group 2 51-75%	Group 3 76-100%	Full Sample	Group 1 1-50%	Group 2 51-75%	Group 3 76-100%
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Risky, lag × $\text{UMP}^{\text{risky}}_{\text{average}}$	5.781***	1.721	5.714	6.147**	6.505***	2.257	4.666	7.399***	6.073***	1.088	3.229	7.007***
	(0.008)	(0.734)	(0.269)	(0.017)	(0.003)	(0.651)	(0.354)	(0.004)	(0.003)	(0.861)	(0.383)	(0.005)
Gov, lag × $\text{UMP}^{\text{gov}}_{\text{average}}$	-2.772	-2.422	-3.413	-0.108	-2.115	0.399	-3.692	0.690	-2.446	2.415	-2.689	-0.711
	(0.326)	(0.822)	(0.266)	(0.978)	(0.466)	(0.971)	(0.261)	(0.861)	(0.385)	(0.828)	(0.342)	(0.850)
Risky, lag	-0.501***	-0.361***	-0.490***	-0.512***	-0.498***	-0.288***	-0.513***	-0.504***	-0.497***	-0.297***	-0.580***	-0.489***
	(0.000)	(0.002)	(0.000)	(0.000)	(0.000)	(0.006)	(0.000)	(0.000)	(0.000)	(0.010)	(0.000)	(0.000)
Gov, lag	0.224***	0.276*	0.117	0.247***	0.225***	0.175	0.133*	0.248***	0.243***	0.100	0.122	0.270***
	(0.002)	(0.068)	(0.128)	(0.009)	(0.001)	(0.106)	(0.083)	(0.008)	(0.001)	(0.607)	(0.121)	(0.003)
Δ Income	0.299***	0.508***	0.296	0.296***	0.304***	0.415**	0.437**	0.283***	0.307***	0.405*	0.406**	0.287***
	(0.000)	(0.009)	(0.111)	(0.003)	(0.001)	(0.025)	(0.028)	(0.010)	(0.000)	(0.052)	(0.024)	(0.003)
Δ Net wealth	0.0690***	0.0695***	0.0872***	0.0624***	0.0697***	0.0685***	0.0883***	0.0623***	0.0684***	0.0763***	0.0755***	0.0633***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Net wealth	0.0179***	0.00978	0.0157**	0.0180***	0.0189***	0.0181*	0.0132*	0.0191***	0.0164***	0.0214*	0.00846	0.0161***
	(0.000)	(0.305)	(0.049)	(0.002)	(0.000)	(0.074)	(0.082)	(0.002)	(0.001)	(0.086)	(0.292)	(0.008)
Income	0.147**	0.626**	-0.00232	0.142	0.131*	0.297	0.167	0.151	0.139*	0.589*	-0.0215	0.130
	(0.044)	(0.013)	(0.995)	(0.217)	(0.090)	(0.240)	(0.667)	(0.252)	(0.060)	(0.095)	(0.956)	(0.278)
Constant	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
Demographics, lag	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
Risk aversion	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
Time FE	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
Observations	2,882	483	817	1582	2,703	511	791	1,401	2,655	372	761	1,522
R^2	0.335	0.227	0.350	0.353	0.338	0.210	0.379	0.354	0.339	0.243	0.398	0.347

Note: The table reports the estimates from Equation (2.4) for different subsamples of households, i.e households that are employed both in $t-1$ and t (Column 1-4), households that have no mortgage both in $t-1$ and t (Column 5-8) and households that are home owners both in $t-1$ and t (Column 9-12). For each subsample, four columns are reported. The first column shows the results when considering the full sample. In the following three columns households are split according to the value of their disposable income: Group 1 includes households in the bottom two quartiles of the income distribution, Group 2 contains households in the third quartile, and Group 3 comprises households in the top quartile. In all columns the dependent variable is the stock (in euro) of Italian risky asset active rebalancing. The variables $\text{UMP}^{\text{risky}}_{\text{average}}$ and $\text{UMP}^{\text{gov}}_{\text{average}}$ are constructed estimating Equation (2.3) with daily data and then following the procedure explained in Section 2.4.1 to construct a biennial series. Only the coefficients of interest and the households' financial controls are reported here. Standard errors (in parenthesis) are robust to heteroskedasticity. Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 2.15: UMP and government bonds active rebalancing - controlling for UMP second order effects

	Always Employed				No Mortgage				Home Owner			
	Full Sample	Group 1 1-50%	Group 2 51-75%	Group 3 76-100%	Full Sample	Group 1 1-50%	Group 2 51-75%	Group 3 76-100%	Full Sample	Group 1 1-50%	Group 2 51-75%	Group 3 76-100%
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Risky, lag $\times$ $\text{UMP}^{\text{risky}}_{\text{average}}$	4.799***	6.699	0.807	4.946**	4.748***	10.99	-0.987	4.998**	3.966**	18.27**	-0.310	4.016**
	(0.006)	(0.352)	(0.874)	(0.015)	(0.007)	(0.131)	(0.862)	(0.015)	(0.023)	(0.029)	(0.955)	(0.048)
Gov, lag $\times$ $\text{UMP}^{\text{gov}}_{\text{average}}$	-3.189	8.781	-4.955	-4.593*	-1.573	9.016	-3.320	-3.099	-1.986	11.89**	-3.272	-4.084
	(0.232)	(0.148)	(0.305)	(0.098)	(0.563)	(0.131)	(0.520)	(0.283)	(0.477)	(0.039)	(0.538)	(0.151)
Risky, lag	0.0803**	0.0773	0.0632	0.0849*	0.0877**	-0.00624	0.0774	0.0924*	0.0757*	0.0383	0.0864	0.0705
	(0.048)	(0.455)	(0.394)	(0.083)	(0.033)	(0.944)	(0.311)	(0.063)	(0.068)	(0.734)	(0.263)	(0.164)
Safe, lag	-0.647***	-0.615***	-0.627***	-0.660***	-0.685***	-0.658***	-0.683***	-0.687***	-0.680***	-0.676***	-0.671***	-0.667***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Δ Income	0.0689	0.393**	0.122	-0.000351	0.108	0.398***	0.167*	0.0299	0.0980	0.428***	0.181*	0.0394
	(0.298)	(0.011)	(0.192)	(0.997)	(0.104)	(0.002)	(0.082)	(0.756)	(0.140)	(0.002)	(0.057)	(0.666)
Δ Net wealth	0.0399***	0.0365***	0.0292***	0.0422***	0.0375***	0.0304***	0.0251***	0.0419***	0.0362***	0.0304***	0.0226***	0.0396***
	(0.000)	(0.000)	(0.001)	(0.000)	(0.000)	(0.001)	(0.004)	(0.000)	(0.000)	(0.004)	(0.007)	(0.000)
Net wealth	0.00942***	0.00920	0.00317	0.0110**	0.00719**	0.00583	-0.000778	0.00939*	0.00524	0.00391	-0.00499	0.00747
	(0.007)	(0.127)	(0.540)	(0.025)	(0.037)	(0.287)	(0.881)	(0.064)	(0.151)	(0.565)	(0.390)	(0.136)
Income	0.122**	0.459***	-0.00177	0.0361	0.144***	0.528***	0.112	0.0564	0.169***	0.545***	0.0968	0.0813
	(0.016)	(0.003)	(0.993)	(0.691)	(0.004)	(0.000)	(0.592)	(0.569)	(0.001)	(0.000)	(0.646)	(0.393)
Constant	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
Demographics, lag	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
Risk aversion	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
Time FE	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
Observations	2,755	834	861	1,060	2,772	931	856	985	2,530	668	829	1,033
R^2	0.415	0.357	0.540	0.390	0.437	0.386	0.554	0.407	0.421	0.439	0.533	0.381

Note: The table reports the estimates from Equation (2.4) for different subsamples of households, i.e households that are employed both in $t-1$ and t (Column 1-4), households that have no mortgage both in $t-1$ and t (Column 5-8) and households that are home owners both in $t-1$ and t (Column 9-12). For each subsample, four columns are reported. The first column shows the results when considering the full sample. In the following three columns households are split according to the value of their disposable income: Group 1 includes households in the bottom two quartiles of the income distribution, Group 2 contains households in the third quartile, and Group 3 comprises households in the top quartile. In all columns the dependent variable is the stock (in euro) of Italian government bonds active rebalancing. The variables $\text{UMP}^{\text{risky}}_{\text{average}}$ and $\text{UMP}^{\text{gov}}_{\text{average}}$ are constructed estimating Equation (2.3) with daily data and then following the procedure explained in Section 2.4.1 to construct a biennial series. Only the coefficients of interest and the households' financial controls are reported here. Standard errors (in parenthesis) are robust to heteroskedasticity. Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 2.16: The effects of UMP on households' change in risk aversion

	HHs investing in risky assets		HHs investing in gov bonds	
	Full Sample	No Inconsistent Answers	Full Sample	No Inconsistent Answers
	(1)	(2)	(3)	(4)
Risky, lag × UMP^{risky}	-0.0000721**	-0.0000638**	-0.000104**	-0.000105**
	(0.0000288)	(0.0000292)	(0.0000463)	(0.0000463)
Safe, lag × UMP^{safe}	0.0000124	0.00000338	0.0000162	0.0000158
	(0.0000487)	(0.0000472)	(0.0000457)	(0.0000458)
Risky, lag	-0.00000125**	-0.00000115**	-0.00000334***	-0.00000338***
	(0.000000508)	(0.000000529)	(0.000000940)	(0.000000942)
Safe, lag	-0.000000733	-0.00000135	0.000000625	0.000000668
	(0.00000128)	(0.00000129)	(0.000000951)	(0.000000952)
Δ Income	-0.00000585***	-0.00000527**	-0.00000479*	-0.00000474*
	(0.00000216)	(0.00000244)	(0.00000267)	(0.00000267)
Δ Net wealth	-0.000000312**	-0.000000415***	-0.000000317*	-0.000000317*
	(0.000000138)	(0.000000154)	(0.000000179)	(0.000000180)
Net wealth	-7.37e-08	5.02e-08	-0.000000222*	-0.000000227*
	(0.000000102)	(0.000000114)	(0.000000131)	(0.000000132)
Income	-0.00000562***	-0.00000644***	-0.00000845***	-0.00000845***
	(0.00000167)	(0.00000182)	(0.00000202)	(0.00000202)
Cuts	yes	yes	-yes	yes
Demographics, lag	yes	yes	-yes	yes
Risk aversion	yes	yes	-yes	yes
Time FE	yes	yes	-yes	yes
Observations	3,093	2,350	3,023	3,014
Pseudo R^2	0.174	0.123	0.223	0.221

Note: The table reports the estimates from Equation (3.4) including data from 2008 to 2016. In all four columns the dependent variable is households' change in risk aversion, a variable equal to -1 if the household has experienced a decrease in risk aversion, is equal to 0 if household's risk aversion has remained unchanged and is equal to 1 if risk aversion has increased between $t-1$ and t. Column 1 shows the results when considering the full sample of households investing in Italian risky assets. Column 2 reports results when eliminating inconsistent answers. Column 3 shows the results when considering the full sample of households investing in Italian goverment bonds. Column 4 reports results when eliminating inconsistent answers. The variables $\text{UMP}^{\text{risky}}_{\text{average}}$ and $\text{UMP}^{\text{gov}}_{\text{average}}$ are constructed estimating Equation (2.3) with daily data and then following the procedure explained in Section 2.4.1 to construct a biennial series. Only the coefficients of interest and the households' financial controls are reported here. Standard errors (in parenthesis) are robust to heteroskedasticity. Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 2.17: Robusness checks - Heckman selection model

	Risky assets 2^{nd} stage (1)	Risky assets 1^{st} stage (2)	Gov Bonds - 2^{nd} stage (3)	Gov Bonds - 1^{st} stage (4)
Risky, lag × $\text{UMP}^{\text{risky}}_{\text{average}}$	6.239***		4.589***	
	(0.948)		(1.100)	
Gov, lag × $\text{UMP}^{\text{gov}}_{\text{average}}$	-2.501		-1.851	
	(1.910)		(1.141)	
Risky, lag	-0.507***		0.0665***	
	(0.0171)		(0.0211)	
Gov, lag	0.221***		-0.693***	
	(0.0435)		(0.0232)	
Δ Income	0.294***		0.0785	
	(0.0723)		(0.0607)	
Δ Net wealth	0.0673***		0.0417***	
	(0.00455)		(0.00397)	
Net wealth, lag	0.0168***	5.04e-07***	0.0108***	5.81e-08
	(0.00350)	(5.62e-08)	(0.00303)	(5.41e-08)
Income, lag	0.0745	1.51e-05***	0.0910*	1.02e-05***
	(0.0615)	(9.50e-07)	(0.0552)	(8.88e-07)
Inverse Mills ratio	-4415.6***		-4484.8*	
	(1588.3)		(2720.2)	
Constant	yes	yes	yes	yes
Demographics, lag	yes	yes	yes	yes
Risk aversion	yes	yes	yes	yes
Time FE	yes	yes	yes	yes
Observations		19,675		19,730

Note: The table reports the estimates from an Heckman selection model (Equation 2.6) including data from 2008 to 2016. Column 2 and 4 report the estimates of the first stage of the model, i.e. a probit model for the binary choice of rebalancing conditioning on not being invested in the asset category in $t-1$ (extensive margin decision); The dependent variables are dummy variables equal to 1 if a rebalancing of risky assets (Column 2) and government bonds (Column 4) has occurred between $t-1$ and t, zero otherwise. Column 1 and 3 include the estimates of the rebalancing equation conditioning to participation. The dependent variables are the (stock of) active rebalancing of risky assets and government bonds, respectively. The variables $\text{UMP}^{\text{risky}}_{\text{average}}$ and $\text{UMP}^{\text{gov}}_{\text{average}}$ are constructed estimating Equation (2.3) with daily data and then following the procedure explained in Section 2.4.1 to construct a biennial series. Only the coefficients of interest and the households' financial controls are reported here. Standard errors (in parenthesis) are robust to heteroskedasticity. Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 2.18: Robusness checks - UMP identification

	Risky assets (1)	Gov bonds (2)	Risky assets (3)	Gov bonds (4)	Risky assets (5)	Gov bonds (6)
Risky, lag × $\text{UMP}^{\text{risky}}_{\text{only sign}}$	4.535***	3.750**				
	(1.608)	(1.506)				
Safe, lag × $\text{UMP}^{\text{gov}}_{\text{only sign}}$	-2.484	-2.343				
	(2.858)	(2.746)				
Risky, lag × $\text{UMP}^{\text{risky}}_{\text{panel}}$			7.174***	4.710**		
			(2.260)	(1.838)		
Safe, lag × $\text{UMP}^{\text{gov}}_{\text{panel}}$			-2.258	-1.962		
			(2.689)	(2.532)		
Risky, lag × $\text{UMP}^{\text{risky}}_{\text{two-day}}$					5.801**	4.597**
					(2.252)	(1.785)
Safe, lag × $\text{UMP}^{\text{gov}}_{\text{two-day}}$					-2.352	-2.086
					(2.495)	(2.399)
Risky, lag	-0.496***	0.0873**	-0.525***	0.0614*	-0.501***	0.0798**
	(0.0427)	(0.0428)	(0.0399)	(0.0353)	(0.0420)	(0.0394)
Safe, lag	0.218***	-0.678***	0.225***	-0.680***	0.212***	-0.681***
	(0.0683)	(0.0497)	(0.0678)	(0.0485)	(0.0675)	(0.0480)
Δ Income	0.307***	0.0913	0.305***	0.0926	0.298***	0.0899
	(0.0775)	(0.0598)	(0.0774)	(0.0598)	(0.0776)	(0.0596)
Δ Net wealth	0.0672***	0.0369***	0.0671***	0.0369***	0.0671***	0.0370***
	(0.00703)	(0.00543)	(0.00699)	(0.00545)	(0.00701)	(0.00545)
Net wealth	0.0184***	0.00793**	0.0183***	0.00800**	0.0185***	0.00799**
	(0.00436)	(0.00323)	(0.00434)	(0.00321)	(0.00434)	(0.00320)
Income	0.140**	0.142***	0.139**	0.142***	0.137**	0.141***
	(0.0690)	(0.0457)	(0.0685)	(0.0456)	(0.0687)	(0.0456)
Constant	yes	yes	yes	yes	yes	yes
Demographics, lag	yes	yes	yes	yes	yes	yes
Risk aversion	yes	yes	yes	yes	yes	yes
Time FE	yes	yes	yes	yes	yes	yes
Observations	3,093	3,023	3,093	3,023	3,093	3,023
R^2	0.335	0.431	0.339	0.431	0.336	0.431

Note: The table reports the estimates from Equation (2.4) including data from 2008 to 2016. The dependent variables are the stock (in euro) of active rebalancing of risky assets (Column 1 and 3) and of government bonds (Column 2 and 4). The variables $\text{UMP}^{\text{risky}}_{\text{only sign}}$ and $\text{UMP}^{\text{gov}}_{\text{only sign}}$ are constructed estimating Equation 2.3 and theb following the procedure expalined in Section 2.4.1 but aggregating only significant *beta*s. The variables $\text{UMP}^{\text{risky}}_{\text{panel}}$ and $\text{UMP}^{\text{gov}}_{\text{panel}}$ are constructed using panel techniques to estimate the daily impact of ECB's unconventional announcements (Equation 3.6) and then following the procedure explained in Section 2.5.2. Variables $\text{UMP}^{\text{two-day}}_{\text{risky}}$ and $\text{UMP}^{\text{gov}}_{\text{two-day}}$ are constructed using a two-day window to capture the impact of ECB's unconventional announcements (Equation 2.8) and then following the procedure explained in Section 2.4.1. Only the coefficients of interest and the households' financial controls are reported here. Standard errors (in parenthesis) are robust to heteroskedasticity. Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 2.19: Robusness checks - estimation sample

	Risky Assets weighted (1)	Gov Bonds weighted (2)	Risky Assets quality check (3)	Gov Bonds quality check (4)
Risky, lag × $\text{UMP}^{\text{risky}}_{\text{average}}$	6.876***	2.666*	7.576***	4.363**
	(2.549)	(1.512)	(2.565)	(2.100)
Gov, lag × $\text{UMP}^{\text{gov}}_{\text{average}}$	-3.717	-0.563	-2.397	-0.735
	(3.565)	(2.532)	(2.802)	(3.590)
Risky, lag	-0.534***	0.0353	-0.491***	0.0750
	(0.0549)	(0.0339)	(0.0500)	(0.0487)
Gov, lag	0.257***	-0.772***	0.229***	-0.681***
	(0.0886)	(0.0490)	(0.0816)	(0.0591)
Δ Income	0.168	0.0689	0.327***	0.145*
	(0.109)	(0.0649)	(0.106)	(0.0846)
Δ Net wealth	0.0640***	0.0307***	0.0741***	0.0405***
	(0.00834)	(0.00652)	(0.00925)	(0.00805)
Net wealth, lag	0.0133**	0.00914**	0.0160***	0.0119**
	(0.00528)	(0.00382)	(0.00593)	(0.00472)
Income , lag	0.147*	0.174***	0.150*	0.106*
	(0.0778)	(0.0470)	(0.0908)	(0.0620)
Constant	yes	yes	yes	yes
Demographics, lag	yes	yes	yes	yes
Risk aversion	yes	yes	yes	yes
Time FE	yes	yes	yes	yes
Observations	3,093	3,023	1,877	1,619
R^2	0.388	0.470	0.348	0.423

Note: The table reports the estimates from Equation (2.4) including data from 2008 to 2016. The dependent variables are the stock (in euro) of active rebalancing of risky assets (Column 1 and 3) and of government bonds (Column 2 and 4). Column (1) and (2) of Table 2.19 report the results when using the weighted sample, while Column (3) and (4) show results when cutting the sample to control for the quality of the survey data. The variables $\text{UMP}^{\text{risky}}_{\text{average}}$ and $\text{UMP}^{\text{gov}}_{\text{average}}$ are constructed estimating Equation (2.3) and then following the procedure explained in Section 2.4.1. Only the coefficients of interest and the households' financial controls are reported here. Standard errors (in parenthesis) are robust to heteroskedasticity. Significance levels: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 2.20: Composition of Italian households' financial portfolio

Perc	Risky Assets	Gov Bonds	Tot Port	Perc	Risky Assets	Gov Bonds	Tot Port
1-5%	282.05	243.32	525.37	*51-55%*	2708.66	3709.46	6418.12
6-10%	102.74	600.40	703.14	56-60%	5330.80	9706.12	15036.92
11-15%	193.73	1561.69	1755.42	*61-65%*	3227.68	9259.44	12487.12
16-20%	279.19	1460.13	1739.31	*66-70%*	4535.45	7019.62	11555.06
21-25%	848.93	2025.16	2874.09	*71-75%*	6031.49	13256.72	19288.21
26-30%	893.31	2001.81	2895.11	*76-80%*	11004.80	13590.84	24592.32
31-35%	1285.07	2512.35	3797.41	*81-85%*	10227.97	7731.53	17959.49
36-40%	1282,79	4631.78	5914.57	*86-90%*	11284.95	8359.34	19644.29
41-45%	3661.5	2707.35	5368.85	*91-95%*	14073.27	13393.51	27466.78
46-50%	3706.02	4726.50	8432.52	*96-100%*	40792.87	14446.77	55239.64

Note: The table shows the average holding of risky assets, government bonds and the value of total portfolio along the income distribution. Averages are computed using sample weights provided in the SHIW.

Source: Bank of Italy's Survey on Household Income and Wealth, 2006.

CHAPTER 3

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

3.1 Introduction

The primary mandate of major central banks is to maintain price stability, which is the reason why central bankers have traditionally payed less attention to the distributional impact of their policy measures on inequality. However, in the aftermath of the Global Financial Crisis of 2008, economic inequality in industrialized countries increased drastically and the public raised concerns that the long-enduring low interest rate environment was exacerbating this problem, since low interest rates might only benefit certain groups of households (Bivens, 2015). One key aspect of the debate on possible redistributional effects of central banks' actions is to understand how financial and demographic characteristics of households may interact with monetary policy. Several papers document monetary policy's heterogeneous effects along income, wealth, house ownership, and employment status of households (see, among the others, Adam and Tzamourani, 2016; Ampudia et al., 2018; Wong, 2019).

In this Chapter we take a different perspective and evaluate the impact of gender and marital status on monetary policy-driven financial portfolio decisions of different household groups, focusing in particular 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,

[1]This chapter is based on joint work with Chi Hyun Kim. 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.

2014a,b; Young, 2018). In particular, in the financial literature the link among gender, marital status, risk preferences, and investment decisions is well established. The empirical evidence shows that men invest significantly more in financial assets than women. Moreover, women are more risk averse (Jianakoplos and Bernasek, 1998; Sunden and Surette, 1998; Fisher and Yao, 2017) and less confident (Barber and Odean, 2001; Croson and Gneezy, 2009). The few works that consider gender and marital status *jointly* when analyzing their implications on financial decisions find similar results, with single female-headed households being the most fragile. Married women have a higher propensity to invest in risky assets than single ones, 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). Moreover, 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. In a prolonged period of low interest rates - and, thus, high asset prices - the single female-headed households' low propensity to bear financial risks and invest in financial assets may lead to a distributional divergence between this household group and the rest of the population.

We investigate how monetary policy influences financial investment decisions focusing on the portfolio choices of US households in the stock market between 2001 and 2017. The choice to concentrate on US households is twofold. First, 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). And the data confirm it. According to *Statista* - a data portal of households - in the year 2018 there are about 15 million US single female-headed households between 25 and 65 years old, representing almost 12% of the US total households. Nonetheless, the distribution of poverty across the US society is skewed towards single women, as 24.9% of them have 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 over the past decade, understanding how it influences single women's financial wealth accumulation is crucial. Second, it allows us to use the household survey data from the Panel Study of Income Dynamics (PSID). The PSID is a nationally representative longitudinal study of US families and it is a rich source of disaggregated and detailed information on households' portfolio composition, wealth, and demographics.

As a first step, we investigate the effect of monetary policy on households' stock market participation, focusing on the binary entry and exit choices of single female- and both married and single male-headed households separately. Our results show that only entry decisions of single women are affected by monetary policy. After a contractionary monetary policy shock, single female-headed households are 11% less likely to enter the stock market. Male-headed households's entry decisions and the exit decisions of both household groups are not affected. This result is interesting because we observe household groups-specific effects of monetary policy only in the entry decisions, even if we control for characteristics that are correlated with gender, marital status and financial wealth, as position in their life cycle, education and income. To the contrary, monetary policy does not affect different household groups' probability to exit the stock market, suggesting that female- and male-headed households do not behave differently once they both participate in the stock market. This result is in line with the findings in the literature that stock market nonparticipant women are more risk averse than men (both participant and nonparticipant) and participant women, and that women 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, we conduct a static simulation exercise and calculate how single female-headed households' entry rate is affected between 2001 and 2017 and how much financial wealth they potentially missed out or gained because of stock market participation choices driven by monetary policy. Our exercise suggests that single female-headed households lost more that \$2000 million over the entire period, suggesting that monetary policy has a sizeable impact.

As a second step, we concentrate on stock market participants and we analyze how single female- and both married and single male-headed households adjust their stock market investment after a (contractionary) monetary policy shock. We do not find any group-specific response to monetary shocks, confirming that households with different gender/marital status do not behave differently once they both participate in financial markets.

One empirical challenge lies on the identification of US monetary policy shocks. We identify monetary policy shocks of the Federal Reserve (Fed) at a daily frequency following the method proposed by Nakamura and Steinsson (2018). With this method, we capture exogenous variations in interest rate futures within a narrow time window around the Federal Open Market Committee (FOMC) meetings. Subsequently, in order to match the frequency of the household survey data, we aggregate the daily monetary policy shocks into a series with biennial frequency. In doing so, we take into account the month of the year in which each household answers the survey questions, thus allowing us to construct an idiosyncratic biennial monetary policy shock series for each household. Afterwards, we improve our identification by exploiting households' heterogeneity in financial wealth and gender, which influences their exposure to monetary policy shocks.

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.

So far the literature investigating how gender and marital status affect portfolio decisions is limited. Several papers focus on the US. Sunden and Surette (1998) highlight the interaction between gender and marital status in determining the allocation of assets in retirement savings plans. Agnew et al. (2003) find that men invest more in equities and trade more frequently than women and married investors invest more aggressively than their single counterparts. Barber and Odean (2001) report that the 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. Few studies consider other countries. In particular, Bertocchi et al. (2011) for Italy and Christiansen et al. (2010) for Denmark gauge the relevance of gender and marital status on stock market participation decisions. Compared to these studies, our paper compare different household groups' portfolio choices conditioning on a specific shock, a monetary policy shock.

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 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 for stocks the PSID additionally asks the households about their purchases or sales, 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 (the last available survey wave). One important feature of the PSID data is that the interviews happen every other year (in odd years) between March and December, and the answers to questions regarding wealth refer 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, but for our analysis this does not constitute an issue, as we mainly focus on wealth variables which are all measures at the same point in

[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).

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.

3.2.1.1 The 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 anaysis, instead of only the household head, is driven by the impossibility to recover the family member in charge of financial decisions in the married families. In fact, in the PSID the term head refers by default 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 recover 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 *single female-headed households* (SFHHs), *male-headed households* (MHHs), which include both single and married households, and *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 as *stock market participants* 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 a zero stock market investment in $t-1$, but a positive one in t, and zero if its stock market 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 consider only stock market participants to analyze their stock market portfolio choices. In order to do so, we need to decompose

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

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 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 this asset classe purchases or sells, we proceed as follows. First, we approximate the capital gain/loss on this asset class between $t-1$ and t with the return 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*, that is the amount of riskless assets sold or bought between $t-1$ and t.[5]

Finally, we define liquid assets as the sum of holdings of stocks, riskless assets and other assets.

3.2.1.3 Sample selection

We include only households that participate in the survey for at least three consecutive waves.[6] We exclude households where the age of the head is younger than 25 years or older than 65 years. Also, 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; and (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).[7]

[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] Please refer to Forti Grazzini (2020) for further explanations of the methodology.

[6] This choice is driven by the large number of households appearing in the PSID for only one wave. Moreover, in the robustness check section in the appendix, we relax this constraint and show that results do not change.

[7] 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. This would dramatically reduce the accuracy of the weighted data in representing the target population.

Although the sample size of each wave ranges over 5000 to 7000, the size or our final data set is significantly reduced due to the data requirements that we impose.

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 statistics of the 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.[8]

With regard to all households of our sample (Panel *A*), on average, 29% of the male-headed households participate in the stock market during 2001-2017. 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 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 prefer riskier financial investments (with a 17% of their portfolio invested in risky assets, compared to the female 9%). Thus, female-lead households appear to hold less wealth and to invest a lower percentage of their net worth and financial portfolio in risky assets. This stock investment rate differential might be explained by several factors and household charachteristics, but is also in line with the literature documenting that women are more risk averse (and thus cautious) in their investment behavior.

The picture is slightly different when considering participating households (Panel *B*). Although female-headed households 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). Stock market participants tend to be wealthier with higher financial wealth and with higher education level. Moreover, participants invest a higher share of their financial portfolio in stocks. However, home ownership and employment are comparable. 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, suggesting that the former might require more wealth to bear the risk to invest in risky assets being, in turn, more risk averse. The ratios between

[8]In particular, each household is weighted using the sample weight provided by the PSID for the first wave the household appears in our data set. For an overview of the unweighted summary statistics, see Table 3.7 in the Appendix.

SFHHs and MHHs that hold stocks in Panel A and B (7.6 and 4.4, respectively) point in the same direction.

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	

Note: 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 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.

3.3 The identification of monetary policy

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 to them. Fourth, we need to take into account the household head's gender/marital status.

3.3.1 Monetary policy shocks

In order to measure monetary shocks we use a high frequency identification technique (see, among the others, Kuttner, 2001; Gürkaynak et al., 2005; Gertler and Karadi, 2015; 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 itself. In order to ensure the exogeneity of the shock measure, it is crucial that the time span around the FOMC meeting is short enough, so to make it highly likely 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.[9] We use daily data from January 1, 2001, to December 31, 2017, and we include all FOMC scheduled meetings that happened throughout this 17 year period. Figure 3.1 depicts the time series of the policy news shock.[10]

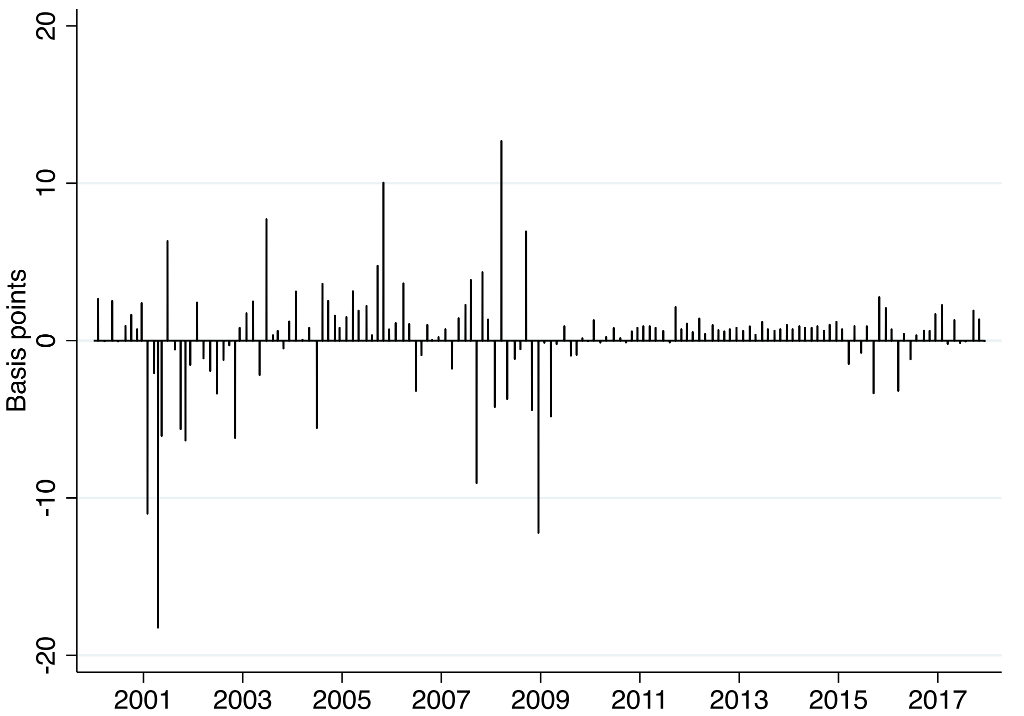
3.3.2 Obtaining biennial household-specific monetary policy shocks

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 would be 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

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

[10]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 3.7 in the Appendix. Between 2001 and 2014 (the Nakamura and Steinsson (2018)'s monetary shock series is available only until end 2014) The correlation between the two shock series is 0.88.

Figure 3.1: The monetary policy news shocks

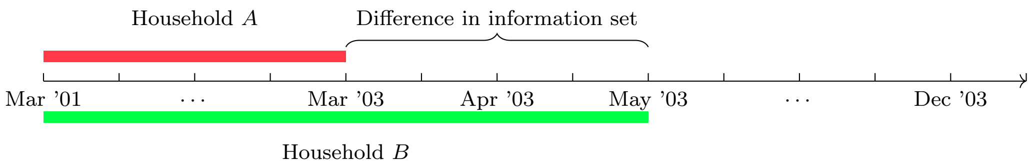


Note: The graph shows the monetary policy news shocks for the period 2001 − 2017 estimated at the daily frequency. The shocks are constructed as the first principal component of the daily change in five interest rate futures around Fed's FOMC meetings. The included futures are the federal funds futures (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 monetary policy shocks is scaled to have a 100 basis point impact on the 1-year US Treasury yield.

regarding the financial variables refer to different periods. Indeed, as can be seen in Figure 3.6 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 have been interviewed during the years 2001 and 2003. Suppose that in the year 2001 they are both interviewed in March, while in 2003, household A is interviewed in March while 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 and then evaluate 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



Note: The figure helps visualizing 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 any 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 set by summing up the monetary policy shocks for each household and taking into account their interview dates,

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

where $MP_{i,t}$ is the biennial monetary policy news shock series for household i in wave t; $I_{i,t-1}$ and $I_{i,t}$ 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 depicts 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 read 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 variable.

Figure 3.3: Boxplot of the biennial household-specific monetary policy news shock by wave

Note: The graphs depicts 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.

Finally, 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).

3.3.3 Households' heterogeneous exposure to monetary policy

The small size and transitory nature of monetary shocks makes it challenging to obtain tight standard errors for estimated responses of household variables to the monetary shocks obtained in the previous section, especially when considering the effects of those shocks on data on low frequency data. However, the estimate precision increases greatly for estimates of the *differential* responses across households. Households differ along different dimensions that can affect their response to monetary policy (see, among the others, Wong, 2019; Forti Grazzini, 2020; Cumming and Hubert, 2020). This heterogeneity can be exploited to understand their exposure to monetary policy and, in turn, their different response to it. Thus, comparing the behavior of households that are supposed to be more and less affected should improve identifying the effect of monetary policy on the variables of interest. Moreover, the differential responses allow for a better understanding of the transmission mechanisms of monetary policy. We follow Forti Grazzini (2020) and consider their financial wealth as source of household heterogeneity. The intuition behind this choice is that the more financial wealth a household holds, the more it is affected by monetary shocks due to their *valuation effects*: monetary policy impacts yields and prices of assets and, thus, it affects the value of households' stock of financial wealth. The more a household is impacted by monetary policy, the stronger it should react to it. Thus, 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). Afterwards, we focus exclusively on stock market participants (i.e., having positive investment in stocks for at least two consecutive waves) and analyze the effect of monetary policy on their trading activity. Throughout this section we evaluate the impact of a contractionary monetary policy scaled such 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 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 excercise 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 the monetary policy impact changes for different household groups (β_2), for different values of the exposure variable (β_3) and for 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. It shows marginal effects at the means, i.e. evaluated at the sample mean of the explanatory variables. We only show the marginal effects of the parameters of interest. Column (1) and (2) present the baseline results, where we compare single female- and male-headed households. Thus, the dummy $Head_i$ equals 1 if the household head is single and female, 0 if the head is male (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 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). Then 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 decrease by 11%. On the contrary, the entry probability of MHH is not significant.

Note that the male-headed household group contains both single and married households. To make sure that the baseline results are not driven by any 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 significant impact on households' stock market exit decision and it only affects the entry choice of single female-headed ones.

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 housholds 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 market, suggesting that female- and male-headed households do not behave differently once they participate in financial markets. This implies that there is some unobserved characteristics that correlates with the interaction between gender/marital status and monetary policy and that makes nonparticipant single female- and male-headed house-

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

holds significantly different from each other in their entry decision but not in their exit choice.

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

Note: 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.

3.4.1.1 Monetary policy and stock market participation at different levels of liquid assets

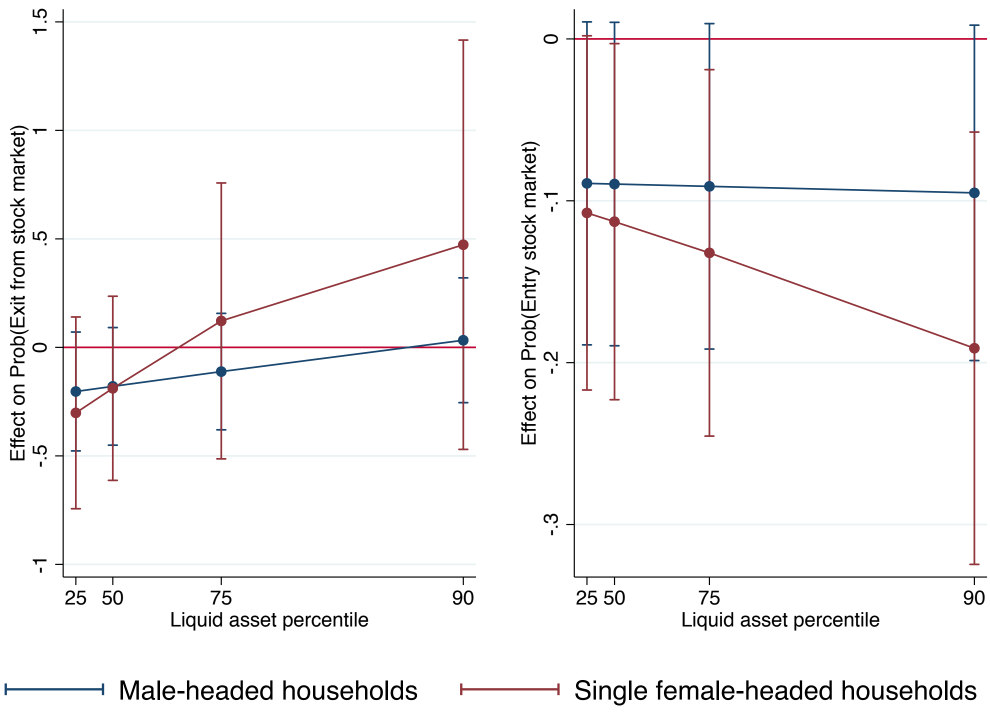
As explained in Section 3.3.3, we improve the the identification of monetary policy using lagged liquid assets, $W_{i,t-1}$, as the exposure variable to monetary policy. The intuition behind this choice is that the more financial wealth a household holds, the more it is affected by monetary policy through the valuation effects of monetary shocks. In the previous section we decribe the (marginal effect) of monetary policy on exit and entry decisions for both female- and male-headed households considering the average holding of lagged liquid assets.[12] In this section we refine the analysis looking at how the reponse changes for different levels of liquid assets.

Figure 3.4 plots the effect of a contractionary shock normalized so to increase the 1-year Treasury bond yields by 100 basis points on the probability of exit the stock market (left panel) and entry (right panel) for single female-headed households (red line) and male-headed household (blue line) at different percentiles of the lagged liquid assets distribution.[13] The reported coefficients are the marginal effects (with the 90% confidence bands) calculated after estimating Equation (3.4) all evaluated at the sample mean of the explanatory variables but the exposure variable. Figure 3.4 confirms previous section findings. Specifically, monetary policy is found to have no 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), irrespective of the level of liquid assets. 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 gets affected, the more it responds to monetary policy. In fact, the impact of monetary shocks moves from being insignificant for female households in the 25th liquid assets percentile to bigger (in absolute value) significant coefficients, up to a -19% at a p-value of 0.02 for households in the 90th percentile.

[12] The average holding of lagged liquid assets is equal to \$260,064 in the sample used to estimate the impact of monetary policy on exit decisions and it is equal to \$26,672 in the sample used to estimate the impact of monetary policy on entry choice.

[13] 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.

Figure 3.4: Monetary policy and stock market participation decision at different levels of liquid asset distribution - marginal effects



Note: 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.

3.4.1.2 Monetary policy and stock market participation when allowing for more heterogeneity across household groups

The baseline results described in the previous two sections and contained in columns 1 and 2 of Table 3.2 are marginal effects evaluated at sample averages of the explanatory variables, obtained 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 averages. All regressors are interacted with the dummy $Head_i$,

$$\begin{aligned} & y^*_{i,t} = \delta_t + \delta_r + Head_i \times (\alpha X_{i,t-1} + \beta_1 MP_{i,t} + \beta_2 W_{i,t-1} + \beta_3 MP_{i,t} \times W_{i,t-1}) + u_{i,t}, \\ & y_{i,t} = 1\left[y^*_{i,t} > 0\right]. \end{aligned} \tag{3.5}$$

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

Results are reported in Table 3.3. 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.

3.4.2 Monetary policy and active saving

In this section, we focus exclusively 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. (2009a) employ a fixed effect model,

$$\begin{aligned} \mathrm{AS}_{i,t} = & \delta_i + \delta_t + \alpha X_{i,t-1} + \beta_1 MP_{i,t} + \beta_2 MP_{i,t} \times Head_i + \beta_3 MP^*_{i,t} + \\ & \beta_4 (MP^*_{i,t} \times Head_i) + \beta_5 (W_{i,t-1} \times Head_i) + \beta_6 W_{i,t-1} + \beta_7 Head_i + \varepsilon_{i,t} \end{aligned} \tag{3.6}$$

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.

Table 3.3: Monetary policy and stock market participation decision - marginal effects at the group-specific sample average

	Exit	Entry
	(1)	(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

Note: 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 (3.3) with data from 2001 to 2017 and are evaluated at the group-specific 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.

All remaining terms capture the three-way interaction effect. We are interested in the coefficients β_1-β_4, as they capture the mean effect of monetary policy (coefficient β_1) and how the monetary policy effect changes for different household groups (coefficient β_2), along different values of the exposure variable (coefficient β_3) and for different household groups along different values of the exposure variable (coefficient β_4). Since we consider households that participate in the stock market both in $t-1$ and t, in this exercise 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 3.4. 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 columns 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 households, and after a contractionary monetary policy shock households sell stocks (-0.003, MP^* coefficient). We calculate the economic numbers of this effect: a shock that increases 1-year Treasury yields by 100 basis points, induces investors to sell, on average, \$762 of their stock investment.[14] Nevertheless, the gender/marital status of the household head seems to play no role as the coefficients attached to any of the terms that include the dummy $Head_i$ are not significant. This result indicates that once female headed-households participate in the stock market, their active saving decisions are not systematically different from those of housholds with a male head. This homogeneous response to monetary policy may seem controversial to the literature that documents behavioral differences between genders in the financial markets. Nonetheless, it is important to point out that our results do not reject the fact that different household groups invest heterogeneously, but rather provide evidence that, when participating in the 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. We perform two robustness checks. First, we use the value of the financial portfolios of the previous wave as an 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 3.10 in the Appendix, and they are very similar to the baseline results.

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

Column (2) in Table 3.4 show that previous results on stock investment rebalancing following a constractionary monetary shock are not specific to the single female- and male-headed households contrast, but can be extended to the other two comparison (Column 2). The coefficients attached to (MP^*) are identical to the baseline result in Column (1) (-0.003, significant at the 1% level) and the triple interactions are non significant, implying that there is no statistical difference between the compared household groups.

To sum up, the analysis shows that, consistent with finding in the previous section, single women do not behave differently from single and married male-headed households once they participate in financial markets. This suggests that these household groups are not systematically different in their stock active saving decisions when controlling for financial and demographic characteristics.

3.4.2.1 Monetary policy and riskless active saving

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), a contractionary monetary policy shock should induce investors to rebalance their portfolio by selling risky assets (like stocks) to purchase safer options (like Treasury bonds). In this section, as an additional result, we investigate whether there exists any significant difference among different groups of households in the way they rebalance their riskless investment.

We re-estimate Equation (3.6) using as dependent variable the net purchase amount of riskless assets of household i between $t-1$ and t.[15] Results are shown in Table 3.5. Also in this case, we report the results of two comparisons, SFHHs vsMHHs (Column 1) and SFHHs vs SMHHs (Column 2).

The results on the comparison of single female-headed with all male-headed households are mixed. Column (1) shows that, consistent with the rebalancing channel of monetary policy, after a contractionary monetary shock, on average households buy more riskless assets (+0.088, MP^* coefficient). 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 (-0.700), suggesting that the overall effect for female-headed households is negative. One possible explanation for this finding is the choice of using the 1-year Treasury to calculate the riskless asset active saving component. Thus, we repeat the analysis using alternative US Treasury maturities. Results are reported in Table 3.9. For ease of comparison, Column (1) shows the same results as the corresponding column in Table 3.5. Column (2) and

[15]Please refer to Section 3.2.1.2 for the construction of the *stock active saving* variable.

Table 3.4: Monetary policy and stock active saving

Dummy $Head = 1$ if Dummy $Head = 0$ if	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

Note: This table presents the results of the fixed effect model in Equation (3.6) estimated using the sub-sample of households participating in the stock market over the years 2001-2017. 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.

(3) include results when the dependend variable is calculated using the 2-year and the 10-year Treasury, respectively. While the coefficients attached to MP^* is very similar in all three columns (0.088-0.089), in columns 2 and 3 there is no significant difference in the way single female- and male-headed households the riskless investment. This implies that the gender/marital status impact on the response to monetary policy is sensitive to the Treasury maturity used for the calculation of the dependent variable. To investigate further more, we repeat the analysis using the value of the financial portfolios of the previous wave as an exposure variable and 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 3.11. The robustness check exercises support the finding that there is no gender/marital status-specific response to monetary policy in the riskless investment rebalancing choice.

The remaining column in Table 3.5 shows that there is no significant difference also when contrasting single female- with only married male-headed households.[16]

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

One might suspect that the absence of a household group specific response to monetary policy is driven by the investment decisions of wealthier households. In fact, differences in risk aversion, inertia, and financial literacy across gender could 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.

In order to examine this hypothesis, we repeat our baseline analysis splitting our sample in two. Results are reported in Table 3.12. In panel A we compare single female- and male-headed households in the top 50% of ther respective group liquid assets distribution; in panel B we analyze households in the bottom 50%. The dependent variable is stock active saving. Let us first concentrate on Column (1). Panel A shows that the coefficients are similar to our baseline estimates presented in Table 3.4 (Column 1), although slighly 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.

Taken together, these findings confirm that first, there is no gender/marital status-specific response to montary policy; second, they show that the baseline results in

[16]We repeat the analysis in Column (2) using the 2-year and the 10-year Treasury for the calculation of the riskless asset active saving and we find that the results remain robust. This extra set of findings are not reported here but are available upon request.

Table 3.5: Monetary policy and riskless asset active saving

Dummy $Head = 1$ if	Single female-headed HH	Single female-headed HH
	VS	VS
Dummy $Head = 0$ if	Male-headed HH	Married male-headed HH
	(1)	(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

Note: This table presents the results of the fixed effect model in Equation (3.6) estimated using the sub-sample of households participating in the stock market over the years 2001-2017. The dependent variable is riskless asset 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.

Column (1) of tables 3.4 and 3.5 are driven by the wealthiest households, supporting once again the choice of using financial wealth as exposure variable.

3.4.2.3 Conventional vs. unconventional monetary policy

On December 16, 2008, the federal funds rate - the conventional monetary policy instrument of the Fed - reached the effective zero lower bound. Subsequently, the Fed introduced unconventional monetary policy measures in the attempt to further lower the long-term interest rates of the economy, thus boosting the stagnated economy. The two main unconventional tools were "forward guidance", communication by the FOMC about the expected future path of the federal funds rate over the next several quarters, and "large scale asset purchases" (LSAP), purchases of hundreds of billions of dollars of longer-term U.S. Treasury bonds and mortgage-backed securities.

The Nakamura and Steinsson (2018)' identification strategy pools the exogenous response of financial market participants coming from both conventional and unconventional monetary policies. On the one hand, this is extremely useful because it allows us to contruct a single variable able to capture monetary policy throughout the full 2001-2017 sample. On the other hand, this method does not allow us to distinguish between different types of monetary measures.

In this section we investigate whether conventional and unconventional monetary policies have comparable effects on participating household groups portfolio decisions. We make use of the high frequency identification method of Swanson (2017), which disentangles between three distinct dimensions of monetary policy shocks. As in Nakamura and Steinsson (2018), Swanson (2017) uses a broad range of interest rate futures changes around FOMC announcements. The main difference is that, instead of extracting one principal component, this method extracts the first three principal components and rotates them to give the factors a structural interpretation. The first factor captures the surprise change in the federal funds rate target ("target shock"), the second is the surprise change in forward guidance ("forward guidance shock"), and the third is interpreted as surprise change in LSAPs ("LSAP shock"). For our purposes, we use only the target shock, which captures the conventional part of monetary policy. Figure 3.8 plots the target shock with our monetary policy news shock of Nakamura and Steinsson (2018). We scale it such that it has a 100 basis points effect on the 1-year Treasury yield to make the estimates comparable to our baseline model.[17] We

[17] To make the target shock comparable to the high frequancy monetary policy news shock described in Section 3.3.1, we construct it as the first (out of three) principle component of the daily changes around FOMC days of the same five interest rate futures, the federal funds futures (the current-month contract rate and thecontract rates for each of the next three months) and Eurodollar futures (at two to four quarters in the future). Then, following the procedure outlined in 3.3.2, we construct

re-estimate Equation (3.6) in its baseline specification, i.e. comparing single female- and male-headed households. Table 3.6 shows the results.

The effect of the target shocks on the active saving of stocks is identical to the effect of the monetary policy news shocks (Table 3.4, columns 1), with the two coefficients being identical. After a contractionary monetary policy shock that increases 1-year Treasure by 100 basis points, households sell stocks by approximately \$712. Moreover, consistent with the baseline results, gender/marital status of the head do not play a role in the way households react to conventional shocks. To conclude, we confirm that households respond to conventional and unconventional monetary policies in a very similar fashion.

3.5 A counterfactual analysis

Our empirical analysis shows that the only household group affected by monetary policy are single female-headed houseolds that do not participate in the stock market. 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 female-headed households is affected by monetary policy shocks during our sample period, 2001 - 2017. Afterwards, we calculate how much financial wealth single female-headed households potentially missed out or gained through monetary policy induced non-participation or entry in the stock market.

To construct the counterfactual of how many female-headed households would have participated in the stock market would not there have been monetary policy, 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 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 3.13, Column 1). Then we use the stock market entry rate per wave of single female-headed households provided by the PSID (Table 3.13, Column 2) to calculate how high the entry rate would be if there were no monetary policy shocks (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 point. From Table 3.2 we know that this decreases single female-headed household to enter the stock market by

biennial househol-specific target shocks by summing the high frequancy target shocks within a two-year window that depends on the household's interview month to the PSID survey in each wave.

Table 3.6: Monetary policy and stock active saving - target shock

Dummy $Head = 1$ if	Single female-headed HH
	VS
Dummy $Head = 0$ if	Male-headed HH
MP^{target}	-3493.8
	(2772.0)
MP^{target} × Head	4398.8
	(3501.9)
$MP^{*target}$	-0.003**
	(0.001)
$MP^{*target}$ × Head	0.015
	(0.020)
Constant	yes
Other inter. terms	yes
Financial var., lag	yes
Demographics, lag	yes
Household FE	yes
Time FE	yes
Observations	2,389
R^2	0.01

Note: This table presents the results of the fixed effect model in Equation (3.6) over the years 2001-2017. Only households participating in the stock market both in $t-1$ and t are included. The dependent variable is the stock active saving. 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 $MP_{i,t}^{target}$ is the household-level biennial cumulated monetary policy shock series constructed in Section 3.4.2.3. The variable $MP_{i,t}^{*target}$ 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.

$0.11 \times 0.115 = 0.013$.[18] Thus, in 2003 single female-headed households are 1.3% less likely to enter the stock market. Then 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 monetary policy shock, $0.104 \times (1 + 0.013) = 0.105$.

At first glance, Table 3.13 tells us that fist, monetary policy is mostly of 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 accomodative); second, if we compare the entry rate with and without monerary policy, we would be inclined to think that their 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? As mentioned in Section 3.1, in 2018 there are about 15 million single female-headed households. If we assume that this figure is constant throughout our sample and that monetary policy has a symmetric effect on households, we can calculate the how many single female-headed households enter or do not enter the stock market due to monetary shocks (red line in Figure 3.5).

As a next step, 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 stock market participants single women's average stock holding, which is \$113215.46 (Table 3.1) and again 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 3.13, 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 3.5 present the final numbers.

Figure 3.5 shows that 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. As a next step, we perform an additional and complementary counterfactual exercise. We use again the average stock holding of single women stock market participants and we approximate capital gains of women by using the yearly average S&P, but now we assume that single women

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

make the monetary policy-driven participation decision on the first year they appear in the PSID and cannot change it until the end of the sample (2017). Suppose a single female-headed household enters the stock market in the year 2003 and stay until 2017. The S&P 500 index in 2003 was 500 and in 2017 it was 2000. Thus, the capital gain during this period is $(2000 - 500/500) = 3$, or 300%. Thus, using again stock market participants single women's average stock holding, the capital gain during this period is 339646.48$. We calculate these capital gains from the stock market for every period until 2017. Afterwards, we multiply the capital gains with the number of single female-headed households that are affected by monetary policy. Figure 3.9 in the Appendix presents the numbers. According to the results, between 2003 and 2017 single women missed almost $13 bl of capital gains.

Figure 3.5: Capital gains/losses of single female-headed households due to monetary policy

Note: 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' biennial capital gains/losses due to the monetary policy-driven stock market participation or non participation decision (blue bars).

In summary, we show that although the effect of monetary policy on the stock market entry rate 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, negatively affecting their already higher probability to be living in poverty in the retirement years.

3.6 Conclusion

This Chapter investigates potential heterogeneous effects of monetary policy on stock market investment choices of single female-headed households compared to both single and married male-headed households using US household survey data from 2001 to 2017. Our 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 an 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 how households respond 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. In particular, our results indicate that such controls are important but do not explain away gender and marital effects. Because these controls are imperfect, however, and because unobserved differences may affect investment behavior, we interpret the remaining gender and marital effect on stock market entry choices as descriptive, rather than causal.

Since our results show that monetary policy only affects the entry decisions of single women, it is crucial to make them resilient to monetary policy adjustments and to educate them on saving and investing. In the US, investment decisions are particularly important for the accumulation of wealth for retirement and this is true especially for single female-headed households, as they are the most fragile family groups and they are more likely to live longer in retirement, having fewer working years and lower earnings.

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 Fed 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,

$$ff1_{t-\Delta t} = \frac{d1}{D1} r_0 + \frac{D1 - d1}{D1} E_{t-\Delta t}(r_1), \tag{3.7}$$

where $d1$ denotes the day of the FOMC meeting, $D1$ is the number of days in the month.[19] Accordingly, the current-month federal funds rate contract right after the FOMC meeting is,

$$ff1_t = \frac{d1}{D1} r_0 + \frac{D1 - d1}{D1} E_t(r_1). \tag{3.8}$$

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

$$mp1_t \equiv E_t(r_1) - E_{t-\Delta t}(r_1) = (ff1_t - ff1_{t-\Delta t}) \frac{D1}{D1 - d1}. \tag{3.9}$$

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

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:

$$ff2_{t-\Delta t} = \frac{d2}{D2} E_{t-\Delta t}(r1) + \frac{D2 - d2}{D2} E_{t-\Delta t}(r2), \tag{3.10}$$

where $d2$ and $D2$ are the day of that FOMC meeting and the number of days in the month containing that FOMC meeting, respectively.[20] 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,

$$mp2_t \equiv E_t(r_2) - E_{t-\Delta t}(r_2) = \left[(ff2_t - ff2_{t-\Delta t} - \frac{d2}{D2} mp1_t) \right] \frac{D2}{D2 - d2}. \tag{3.11}$$

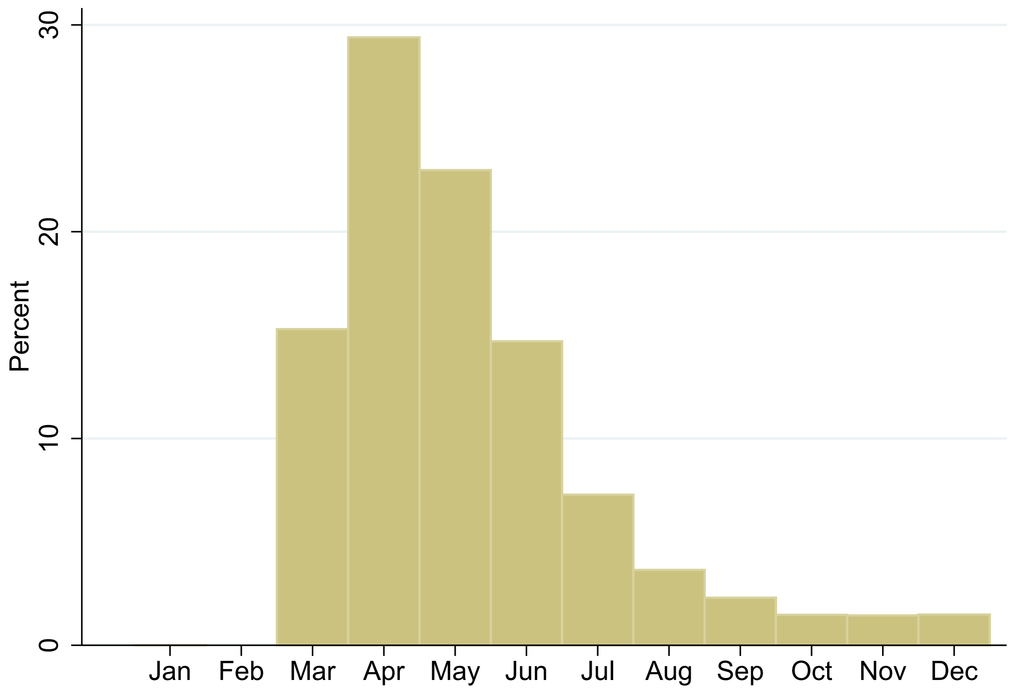
The last set of interest rate futures we use are the change in the price of three Eurodollar futures at the time of the FOMC meetings. Following Nakamura and Steinsson (2018), we use the Eurodollar futures at horizons of two, three, and four quarters in the future.

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.

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

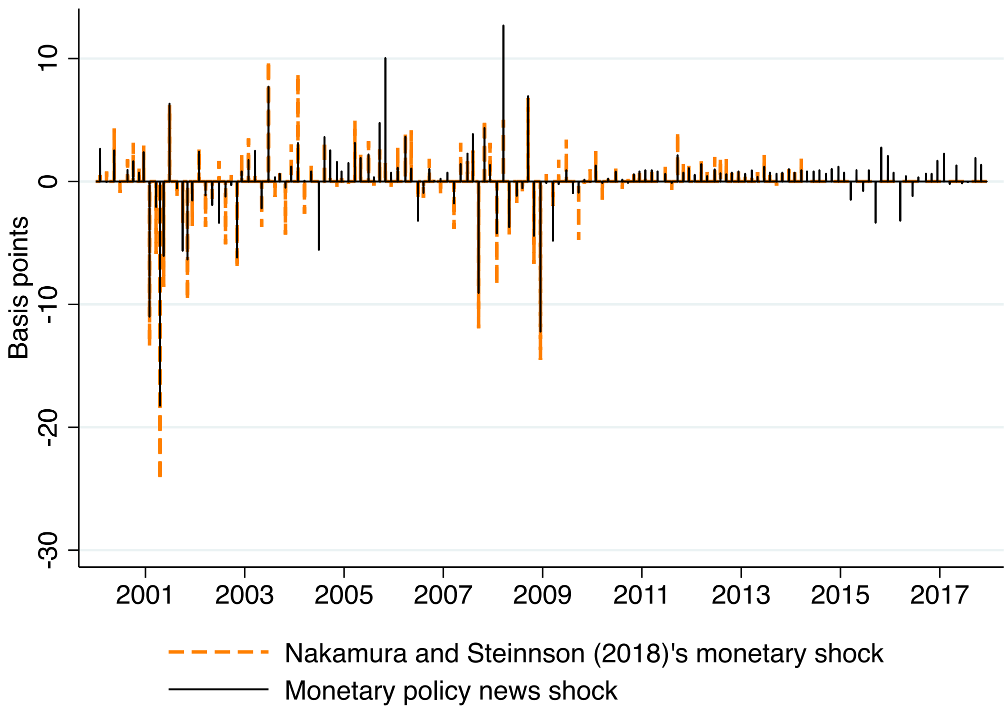
3.B Figures

Figure 3.6: Distribution of PSID interview months, 2003-2017



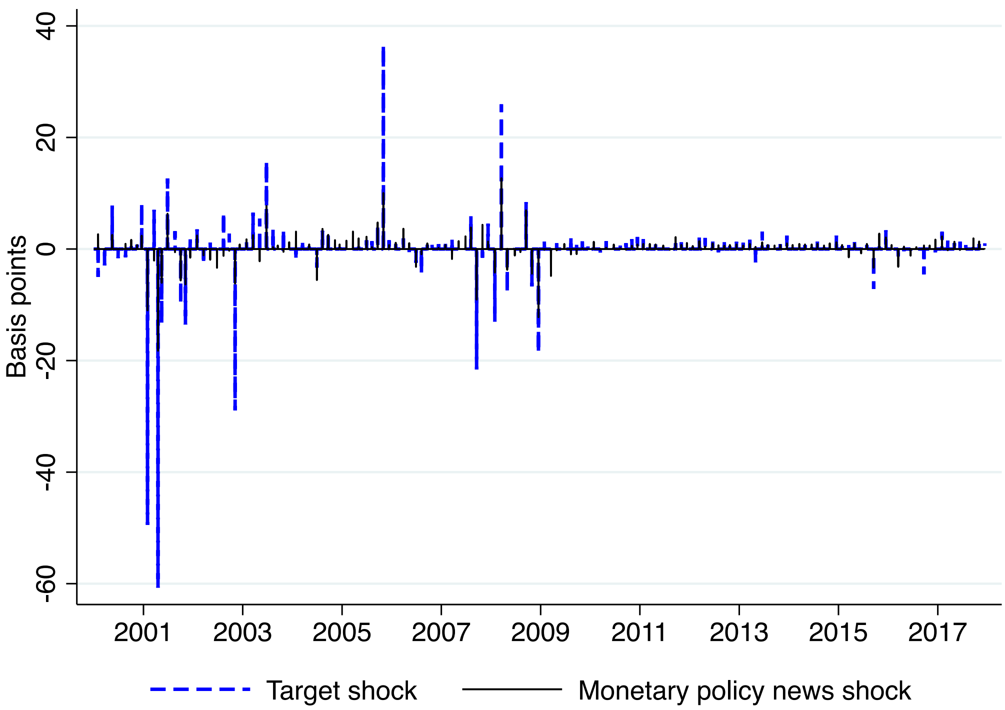
Note: This figure shows the cross sectional distribution of PSID interviews over the year 2003-2017.
Source: PSID

Figure 3.7: The Nakamura and Steinsson (2018)'s original shocks and the monetary policy news shocks



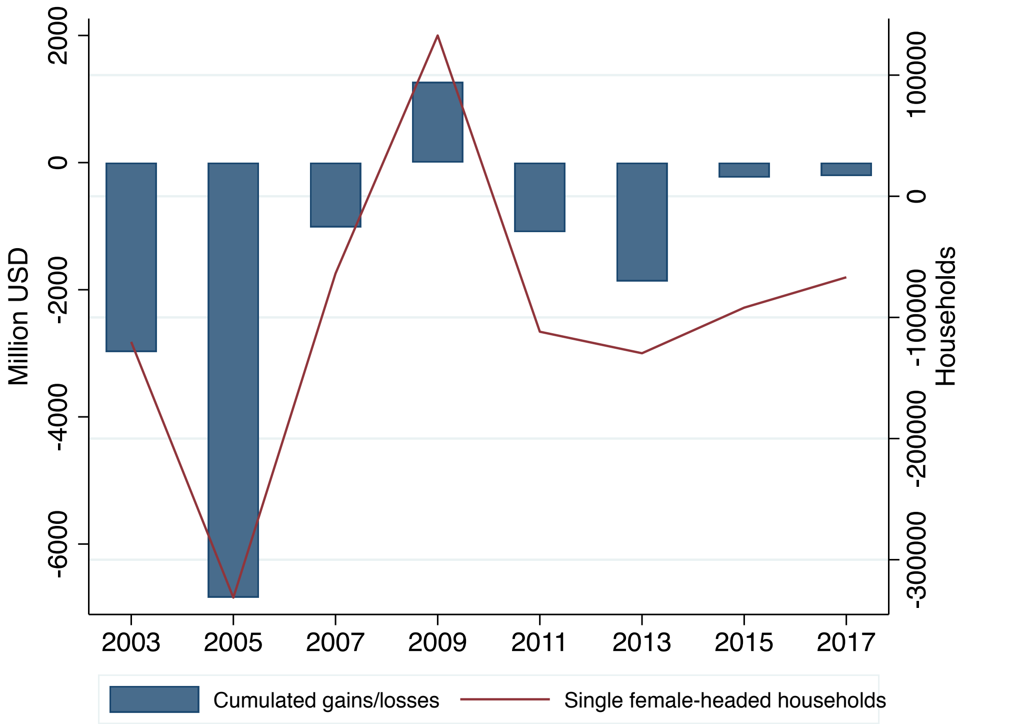
Note: This figure shows two monetary policy shock series. The original Nakamura and Steinsson (2018)'s monetary shock series (dashed orange line) is constructed as the first principal component of the daily change in five interest rate futures around Fed's FOMC meetings and is estimated using intradaily 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 3.8: The target shocks and the monetary policy news shocks



Note: This figure shows two monetary policy shocks estimated at a daily frequency over the $2001 - 2017$ period. The target shocks (dashed blue line) are constructed as the first principal component out of three of the daily change in five interest rate futures around Fed's FOMC meetings. The monetary policy news shocks (solid black line) are constructed as the first principal component of the daily change in five interest rate futures around Fed's FOMC meetings. 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 3.9: Cumulated capital gains/losses of single female-headed households due to monetary policy



Note: 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	

Note: 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

Note: 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: 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

Note: This table the presents results of the fixed effect model in Equation (3.6) 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 3.10: Monetary policy and stock active saving - robustness checks

Dummy $Head = 1$ if	Single female-headed HH	
	VS	
Dummy $Head = 0$ if	Male-headed HH	
	Stock Act. Sav.	Stock Act. Sav.
	(1)	(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

Note: This table presents the results of the fixed effect model in Equation (3.6) 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 3.11: Monetary policy and riskless active saving - only households that participate in the PSID survey for at least two consecutive waves

Dummy $Head = 1$ if Dummy $Head = 0$ if	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 × Head	-5018.0	19281.7
	(52797.8)	39957.9
MP^*	0.081***	0.091**
	(0.010)	(0.012)
MP^* × 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

Note: This table presents the results of the fixed effect model in Equation (3.6) 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.

Table 3.12: Monetary policy and stock active saving - households at the top and bottom of their respective group's liquid asset distribution

Dummy $Head = 1$ if	Single female-headed HH	
	VS	
Dummy $Head = 0$ if	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

Note: This table presents the results of the fixed effect model in Equation (3.6) over the years 2001-2017. Only households participating in the stock market both in $t-1$ and t are included. The dependent variables is stock active saving. In Column (2) 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. The variable $MP_{i,t}$ is the household-level biennial cumulated 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 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.

Table 3.13: 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).

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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.

Caterina Forti Grazzini
Frankfurt am Main, September 2020

Liste verwendeter Hilfsmittel

- Stata 16
- Matlab R2016a
- RStudio 1.1.383 basierend auf R 3.3.0
- Microsoft Excel
- LaTeX
- Siehe auch Literatur- und Quellenangaben