Monetary Policy, Housing Market Dynamics, and the Propagation of Shocks

Inaugural-Dissertation

zur Erlangung des akademischen Grades eines Doktors der Wirtschaftswissenschaften an der Wirtschaftswissenschaftlichen Fakultät der Bayerischen Julius-Maximilians-Universität Würzburg

vorgelegt von

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Würzburg, Mai 2016

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Zusammenfassung

In dieser Dissertation werden die Wechselwirkungen zwischen Geldpolitik und Immobilienmärkten empirisch untersucht. Hierbei beleuchtet die Arbeit potentielle Interaktionen aus *drei unterschiedlichen Perspektiven*: Erstens wird die *systematische Reaktion von Geldpolitik* auf veränderte Immobilienfinanzierungskonditionen untersucht. Zweitens wird der Einfluss des *institutionellen Rahmens einer Währungsunion* auf die Entstehung von Immobilienblasen in Teilen der Währungsunion analysiert und drittens werden die Effekte *exogener Zinsimpulse* auf die Makroökonomie und vor allen Dingen auf Häusermärkte quantifiziert, wobei für die Analyse solcher Effekte explizit Interdependenzen mit Finanzmarktkonditionen Berücksichtigung finden. Methodisch kommen zum Zwecke dieser Analysen vor allem zeitreihenökonometrische Ansätze wie Vektor-Autoregressionen (VAR) oder lokale Projektions-Modelle zur Anwendung.

Nach einer kurzen Einleitung in Kapitel 1 wirft Kapitel 2 der Arbeit die Frage nach den makroökonomischen Effekten von Veränderungen der Beleihungsgrenzen (LTV ratio) für Wohnungsbaukredite in den USA auf. In einem strukturellen VAR-Modell steigen sowohl das Bruttoinlandsprodukt (BIP) als auch die Unternehmensinvestitionen nach einer Lockerung der Beleihungsgrenze signifikant an, wohingegen Wohnbauinvestitionen fallen. Der Einfluss des LTV-Schocks auf die Wohnbauinvestitionen hängt offenbar von der systematischen Reaktion der Geldpolitik ab. Im vorliegenden historischen VAR-Experiment reagiert die Federal Reserve (FED) näm-

lich mit kontraktionärer Zinspolitik auf den Schock und erhöht hierbei (indirekt) sowohl die Hypothekenzinsen als auch die Zinserwartungen der Privathaushalte. In einem kontrafaktischen Politikexperiment wird schließlich der marginale Effekt dieser endogenen Zinskontraktion heraus gearbeitet. Der selbe expansive LTV-Schock führt – in einem hypothetischen Umfeld konstant bleibender Notenbankzinsen – zu einem stärkeren Anstieg der Nicht-Wohnungsbau BIP-Komponenten. Die Wohnbauinvestitionen selbst steigen in diesem Szenario signifikant an. Ähnliches gilt zudem für die Reaktion der Immobilienverschuldung in Folge eines LTV-Schocks. Die systematische Zinsreaktion der FED auf Lockerungen der Beleihungsgrenzen für Wohnbaukredite wird demnach als die entscheidende Einflussgröße für die Reaktion von Immobilieninvestitionen und -verschuldung herausgearbeitet.

Das dritte Kapitel der Dissertation untersucht die negative Korrelation zwischen der Entwicklung der Häusermärkte und der Leistungsbilanz in Spanien. Unter Verwendung eines robusten Vorzeichenrestriktionen-Ansatzes, welcher von einem DSGE-Modell für eine Währungsunion abgeleitet wird, werden die Effekte von in Spanien entstehenden Schocks (pull Faktoren) sowie die Effekte von Schocks im Rest der Währungsunion (push Faktoren) auf den spanischen Immobilienmarkt und die Leistungsbilanz untersucht. Hierbei findet ein VAR-Ansatz Anwendung, welcher die gleichzeitige Verwendung von Monats- und Quartalsdaten erlaubt. Unter den vier identifizierten strukturellen Schocks sind "Savings Glut"-Schocks im Rest der Eurozone, Risikoprämien-Schocks auf spanische Anleihen und spanische Häuserspekulations-Schocks in der Lage, die negative Korrelation in den Daten zu erklären. Lockerungen der spanischen Finanzierungskonditionen sind hingegen nicht im Stande einen Anstieg der spanischen Immobilieninvestitionen oder der Häuserpreise und zur gleichen Zeit eine Passivierung der Leistungsbilanz zu verursachen. Mit Blick auf die empirische Bedeutsamkeit der untersuchten Schocks besitzen "Savings Glut"-Schocks im Rest der Eurozone den größten Erklärungsgehalt für die spanische Häuserpreisentwicklung, wohingegen Risikoprämien-Schocks Fluktuationen der Immobilieninvestitionen in Spanien am besten erklären.

Das letzte Kapitel der Dissertation widmet sich der Fragestellung, ob geldpolitische Impulse in den USA stärkere makroökonomische Effekte und vor allen Dingen auch größere Anpassungen auf Immobilienmärkten auslösen, wenn sie in einem Umfeld von Spannungen im Finanzsystem stattfinden. Dieser Thematik nähert sich die Arbeit unter Verwendung eines lokalen Projektions-Modells, in welchem sowohl die dynamischen Effekte der Geldpolitik als auch die geldpolitischen Impulse selbst von einem Maß für Spannungen im Finanzmarkt abhängen können. Die Messgröße für Spannungen der Finanzmarktkonditionen ist hierbei die sogenannte Überschussprämie auf Unternehmensanleihen (excess bond premium). Die Ergebnisse des Kapitels legen den Schluss nahe, dass geldpolitische Impulse makroökonomische Aggregate, Finanzmarktdaten und Häusermärkte stärker beeinflussen, wenn Verspannungen im Finanzsystem hoch sind. Zudem scheint kontraktionäre Geldpolitik Spannungen im Finanzsystem vor allem dann zu verstärken, wenn diese sich bereits auf hohem Niveau befinden.

Acknowledgment

I wrote this dissertation in the course of my employment as a research and teaching assistant at the Chair for Monetary Policy and International Economics at the University of Würzburg. The work at this chair constituted a stimulating environment to conduct my research and to learn more about the exciting fields of monetary policy and macroeconomics, in general.

Above all, I am deeply grateful to Peter Bofinger, my primary thesis supervisor, for patiently encouraging my work and for inspiring debates during the last four years. He taught me a lot and he definitely shaped my perception of economics. First and foremost, I learned from him to continuously evaluate my research along its applicability to deal with recent macroeconomic problems and to assess my work along its usefulness for policymakers.

Furthermore, I want to thank current and former members at the Chair for Monetary Policy and International Economics for creating a productive research environment and a pleasant workplace – also from a personal perspective. A big thank you goes to Sebastian Debes, Daniel Garcia, Johannes Gareis, Daniel Maas, Eric Mayer, Alix Pianka, Mathias Ries, Petra Ruoss, and my office colleague Philipp Scheuermeyer. Especially, I am much obliged to Eric Mayer for awaking my fascination for dynamic macroeconomics already during my Master's studies and for productive collaborations throughout my appointment at the University of Würzburg.

In addition, I feel blessed that I also had the opportunity to learn from and collaborate with researchers from other Universities. In this regard, I want to particularly thank my secondary thesis supervisor, Johann Scharler, from the University of Innsbruck as well as Rüdiger Bachmann from the University of Notre Dame for their readiness to conduct research with me. These co-operations undoubtedly enriched my doctoral studies in a very instructive way.

Ultimately and most importantly, I want to express my deepest gratitude to my family and my friends for their invaluable support of so many different and important kinds. Without their backing, it would not have been possible for me to realize this dissertation project.

Würzburg, May 2016 Sebastian Rüth

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Chapter 1

Introduction

"All that said, given the fundamental factors in place that should support the demand for housing, we believe the effect of the troubles in the subprime sector on the broader housing market will likely be limited, and we do not expect significant spillovers from the subprime market to the rest of the economy or to the financial system."

(see Bernanke, 2007)

Ben Bernanke at Federal Reserve Bank of Chicago on May 17, 2007
43rd Annual Conference on Bank Structure and Competition

At the onset of the twenty-first century, the US – like several other industrialized economies – saw an unprecedented surge of property prices, residential investment activity, and household sector mortgage debt. While policymakers and, most notably, central bankers (see, e.g., the quote at the beginning) appeared to disregard the macroeconomic risks of overheating housing markets, the downturn in the US residential sector of the years 2006 and 2007 allegedly set the stage for the Great Recession.

One popular narrative of the underlying sources of this housing cycle not only characterizes US monetary policy as having underrated potential residential sector spillovers to the broader economy, but as itself having been the main cause for the run-up in housing market activity. For instance, Taylor (2007) argues that following the dot-com bubble, the Federal Reserve systematically deviated from the Taylor rule by setting its policy instrument too low – thereby fueling housing markets. This view is highly controversial, though, and several other competing hypotheses have been put forward in the literature (see, e.g., Sá and Wieladek, 2015).[1] Albeit the literature so far does not converge to a coherent perception of what kind of shocks predominantly drive fluctuations in housing markets, a cautious consensus view is yet that *housing market developments* spread significantly to the broader economy and *do not only passively reflect overall macroeconomic conditions* (see, among others, Iacoviello and Neri, 2010). Therefore, to the extent that housing market dynamics cause business cycles and ultimately also inflation fluctuations, understanding the impact of monetary policy on housing sector activity – and vice versa – is indispensable for the conduct of monetary policy.

This dissertation studies the interrelations between housing markets and monetary policy from *three different perspectives*. First, it identifies housing finance-specific shocks and analyzes their impact on the broader economy and, most importantly, the *systematic monetary policy reaction* to such mortgage sector disturbances. Second, it investigates the implications of the *institutional arrangement of a currency union* for the potential buildup of a housing bubble in a member country of the monetary union by, inter alia, fostering border-crossing capital flows and ultimately residential investment activity. This dissertation, third, quantifies the effects of *autonomous monetary policy shifts* on the macroeconomy and, in particular, on housing markets by conditioning on financial sector conditions.

From a methodological perspective, the dissertation draws on a recent literature that applies time-series econometrics to study the monetary policy transmission mechanism (see, e.g., Uhlig, 2005; Gertler and Karadi, 2015; Tenreyro and Thwaites,

[1] Among others, Mian and Sufi (2009) document a negative correlation between mortgage credit and income growth in the US from 2002 to 2005 and identify an outward shift of mortgage lenders' (effective) loan supply curve as a driver of the housing cycle. While, e.g., Bernanke (2005), Bernanke (2010), and Sá and Wieladek (2015) emphasize the role of external capital inflows for housing market fluctuations, contributions like Case and Shiller (2003), Shiller (2007), or Towbin and Weber (2016) propagate the notion of expectations-driven housing cycles, which are – at least to some extend – disconnected from fundamental factors.

forthcoming) or to examine housing market fluctuations (see, e.g., Duca et al., 2011; Walentin, 2014; Sá and Wieladek, 2015; Jordà et al., 2015). The dissertation, in particular, makes use of vector autoregressions (VARs), which were introduced by Sims (1980) into macroeconometric research. For instance, Hamilton (1994) or Lütkepohl (2010) provide an accessible textbook overview over VARs, while Watson (1986) surveys the early VAR literature, and Stock and Watson (2001) critically review the VAR methodology. The multivariate approach of modeling time-series in the form of VARs is shown to perform well in terms of summarizing macroeconomic and financial data and also in terms of out-of-sample forecasting properties. To conduct structural inference with these models, i.e., to recover the structural representation of a reduced form VAR, a researcher needs to impose little restrictions on the data.[2] Concretely, this dissertation makes use of three different identification procedures, namely, a Cholesky identification scheme (e.g., Bernanke et al., 1997; Christiano et al., 2005), a sign-restrictions based approach (e.g., Faust, 1998; Uhlig, 2005), as well as a narrative identification procedure as in Romer and Romer (2004). Moreover, to analyze impulse response functions in a non-linear underlying economic environment, the dissertation applies the local projections method introduced by Jordà (2005). Most notably, the local projections framework straightforward accommodates non-linear, multivariate time-series approaches and is less vulnerable to misspecifying the true model structure relative to a VAR. As a consequence, local projections are increasingly employed in recent macroeconomic research (see, among others, Owyang et al., 2013; Bernardini and Peersman, 2015; Born et al., 2015; Jordà et al., 2015; Tenreyro and Thwaites, forthcoming).

The subsequent chapter 2 analyzes macroeconomic implications of exogenous shifts in loan-to-value (LTV) ratios of residential mortgages in the US. Applying a linear VAR model to quarterly time-series, it finds that measures of aggregate economic activity, e.g., GDP and business investment, rise significantly after an unexpected expansion in LTV ratios. Surprisingly though, the VAR evidence also

[2] Among others, this feature differentiates VARs from the related New Keynesian dynamic stochastic general equilibrium (DSGE) framework, which usually involves rich specification choices and a great number of theoretical restrictions (e.g., Clarida et al., 1999; Smets and Wouters, 2003; Christiano et al., 2005; Smets and Wouters, 2007).

reveals a decline in residential investment after the LTV shock. These macroeconomic dynamics are paralleled by an increase in the Federal Funds rate, i.e., the US monetary authority reacts to looser collateral requirements with its policy instrument, which is in line with Walentin (2014), who reports a similar finding for shocks to mortgage rates or mortgage spreads. The endogenous monetary policy tightening, moreover, raises mortgage rates and interest rate expectations of consumers, making the policy tightening a prime candidate to explain the contraction of residential investment following an expansionary LTV shock. In addition, based on an impulse response decomposition in the spirit of Kilian and Lewis (2011), the monetary policy tightening is best characterized as a direct response to the LTV shock, leading to the conclusion that the FED's reaction function apparently comprises mortgage lending conditions. To flesh out the marginal impact of the FED reaction, chapter 2 conducts a policy counterfactual as in, e.g., Bernanke et al. (1997) or Sims and Zha (2006). In this experiment, the FED's policy instrument stays hypothetically fixed at the pre-shock level, when the LTV shock hits the economy. In such a scenario, the surge of non-residential GDP components is more pronounced and, most notably, the LTV easing corresponds with an increase in residential investment activity. The qualitative and quantitative dynamics of residential investment after a LTV disturbance thus appear to be contingent on the endogenous monetary policy reaction.[3] In summary, the empirical evidence suggests that under conventional monetary policy an expansion of LTV ratios is not likely to cause residential investment or household leverage booms. This also allows to infer a policy lesson: to fulfill their intended purpose, supervisory limits on LTV ratios as a macroprudential policy tool to cool down residential sector activity should either be coordinated with monetary policy or at least be set up to consider the potential interplay with monetary policy.

Chapter 3 analyzes interrelations between housing markets and monetary policy by focusing on the drivers of a housing cycle in a member country of a monetary union. Put differently, the chapter studies the role of the institutional arrangement of a monetary union as a potential amplifier – or even the source – of intra-union

[3]The LTV shock evidence for residential investment extends to the impact on household debt, which only rises in the fixed Federal Funds rate environment, whereas firm debt rises independent from the monetary policy stance.

capital flows, which ultimately may drive housing markets. For the case of Spain, the chapter empirically studies the observed negative correlation between the current account and the housing sector activity, which was observed since the early 1990s. The chapter follows an empirical sign restrictions approach by deriving robust theoretical restrictions from a New Keynesian model for a monetary union to identify structural shocks in Spanish and rest of Euro Area time series along the lines of Uhlig (2005) or Peersman and Straub (2009).[4] Drawing on an open-economy VAR approach, the chapter identifies push and pull factors as a source of capital flows inside the European Monetary Union (e.g., Calvo et al., 1993; Fratzscher, 2012). As in Eraker et al. (2014), estimation and inference follow a Bayesian approach for mixed frequency, i.e., monthly and quarterly, data. This chapter reveals that both identified push factors – a rest of Eurozone specific "savings glut" shock and a shock to Spanish bond risk premia – predict the negative correlation between Spanish current account and housing market data (Sá and Wieladek, 2015). The same holds true for a housing bubble (pull disturbance) shock, which also generates this negative co-movement. Interestingly, and counterfactual to the cycle in Spanish housing markets, Spanish financial easing shocks fall short in causing surges in, both, (real) property prices and residential investment, which is consistent with, e.g., Justiniano et al. (2015). Among the four identified shocks and in terms of explained shares of the forecast error variance decomposition, the savings glut disturbance accounts for most of the fluctuations in real property prices and the shock to the Spanish risk premium explains the largest share of variation in real residential investment.

Ultimately, chapter 4 raises the question whether monetary policy is more powerful, i.e., whether the same monetary policy impulse affects the broader economy and, in particular, housing markets more strongly, when strains in the financial system are high. Applying local projections to US time series as in Jordà (2005) and Tenreyro and Thwaites (forthcoming), the chapter approaches this question empirically by allowing monetary policy shocks and its propagation to the broader economy to smoothly (see Granger and Teräsvirta, 1994) vary according to a measure of financial market tensions – the so-called excess bond premium (EBP). Gilchrist and Zakrajsek

[4]The New Keynesian model builds on, e.g., Rabanal (2009) and Iacoviello and Neri (2010).

(2012) provide convincing evidence that the EBP – measuring deviations of corporate versus Treasury bond spreads from expected default risks – constitutes a good proxy for risk attitudes and strains in financial markets as a whole. The chapter reveals that monetary policy surprises, which are identified in a non-linear counterpart to Romer and Romer (2004), impact macroeconomic, housing, and financial variables stronger and more persistently when financial frictions are high. Furthermore, increasing impulse responses of the EBP after a monetary policy contraction indicate a monetary-policy-induced amplification of financial market tensions, in particular, when tensions are already high, which is consistent with the credit channel theory of monetary policy (see Bernanke and Gertler, 1995; Bernanke et al., 1999; Gertler and Karadi, 2015). The results of this chapter, furthermore, have implications for the most recent financial crisis in the US, when the EBP reached unprecedented levels. This chapter suggests that interest rate policy might have been particularly effective during the crisis. Though, given the zero lower bound, the FED's inability to lower interest rates might – ceteris paribus – have aggravated the financial turmoil and, most importantly, the downturn in housing markets more than hitherto supposed.

Chapter 2

Systematic Monetary Policy and the Macroeconomic Effects of Shifts in Loan-to-Value Ratios

"Several other countries have used tools such as time-varying risk weights and time-varying loan-to-value (...) caps on mortgages. Indeed, international experience points to the usefulness of these tools, whereas the efficacy of new tools in the United States, such as the countercyclical capital buffer, remains untested."

(see Fischer, 2007)

Stanley Fischer, Macroprudential Monetary Policy Conference
at the Federal Reserve Bank of Boston (October 2, 2015)

2.1 Introduction[1]

What are the macroeconomic consequences of (exogenous) changes in residential mortgage market loan-to-value (LTV) ratios? The most recent cycle in US housing

[1]This chapter is based on joint work with Rüdiger Bachmann. A recent version can be downloaded from the authors webpages (see Bachmann and Rüth, 2016).

markets saw a relaxation and subsequent tightening of borrowing conditions, leading many observers to attribute the growth in residential investment and mortgage debt prior to the Great Recession to the loosening of collateral requirements. In addition, many policymakers seem to perceive movements in LTV ratios as significant drivers of housing markets, with supervisory limits on LTV ratios representing one of the most important macroprudential policy tools in advanced economies.[2] However, little is known about the macroeconomic consequences of (autonomous) variations in LTV ratios.

The aim of this chapter is to empirically quantify the effect of exogenous shifts in LTV ratios on aggregate economic activity, in particular investment, as well as the development of household and firm debt, and, moreover, to shed light on the interaction between movements in LTV ratios and systematic monetary policy reactions.[3] Our empirical strategy consists of estimating structural vector autoregressions (VARs), which allow the identification of exogenous shocks to LTV ratios by employing only a few theoretical restrictions. To measure LTV ratios, we rely on survey data from the Federal Housing Finance Agency (FHFA), which polls a sample of US mortgage lenders to report terms and conditions on lending standards for conventional mortgages within the Monthly Interest Rate Survey (MIRS), representing the most extensive source of information for this segment. We isolate exogenous shifts in LTV ratios from endogenous reactions to other macroeconomic fluctuations by imposing a recursive Cholesky identification scheme. Consistent with Gilchrist and Zakrajsek (2012) and Walentin (2014), we recover the structural VAR representation by assuming that LTV shocks affect "slow-moving" macroeconomic aggregates with a time lag of one quarter, while "fast-moving" financial variables respond to shifts in lending standards on impact.

After an expansionary 25 basis point LTV shock, the LTV ratio rises quite per-

[2]See also the quote at the beginning, or IMF (2011) and Claessens (2014) for a summary of macroprudential tools and their implementation across countries.

[3]This chapter is thus related to a recent literature that studies the effects of shocks to bank lending standards and financial market conditions on the macroeconomy (e.g., Gilchrist and Zakrajsek, 2012; Bassett et al., 2014; Walentin, 2014; Peersman and Wagner, 2015). In contrast to the existing literature, we focus specifically on the housing market and on lending conditions in terms of quantities (LTV ratios) rather than spreads. Also, this chapter includes a systematic analysis of the impact of monetary policy reacting to credit conditions.

sistently, and we find positive spillovers to non-residential aggregate quantities, with business investment rising significantly by 0.3 percent after a year, and real GDP increasing by approximately 0.1 percent. Because of these spillovers to non-residential aggregates we view residential mortgage LTV ratios as an indicator of banks' lending propensity in a broader sense rather than just being indicative of lending standards in residential mortgage markets.

The picture is different, however, for residential investment: after a small initial increase, residential investment significantly turns negative at minus 0.4 percent in the second year after the shock. We identify the FED's monetary policy instrument as a potential candidate to explain the decline in residential investment after the LTV shock. Indeed, the Federal Funds rate responds to looser lending standards in the residential mortgage market with a hump-shaped tightening of 10 basis points, counteracting the eased *quantity* restriction on mortgage loans. In addition, the endogenous policy contraction significantly passes through to mortgage rates – raising the *price* of mortgage loans – and, furthermore, is anticipated by households as measured by expectations on 12 months ahead borrowing interest rates from the Michigan Survey of Consumers. Our results thus suggest that an exogenous loosening of LTV ratios cannot explain a boom in residential investment at least under conventional monetary policy.

We analyze the systematic monetary policy response along two additional dimensions. First, to answer the question what the FED actually responds to after an LTV easing, we perform an impulse response decomposition as proposed by Kilian and Lewis (2011). At any horizon, the deviation of the Federal Funds rate from its conditional mean can be considered as the sum of the reaction to its own lags and the reaction to realizations of other variables in the VAR. Thus the impulse response decomposition reveals which variables trigger the policy tightening. As the LTV shock causes no inflationary pressure – price inflation even slightly falls in the medium run – we find no evidence for a preemptive price stabilization motive of monetary policy. By contrast, based on the impulse response decomposition, the policy response is better characterized as a *direct* response to the altered lending conditions, rather than an *indirect* response operating through the shock propagation via other variables in the system. For short horizons, the LTV ratio itself accounts for the systematic interest rate contraction almost entirely, and for longer horizons, lags of the Federal Funds rate explain the majority of the policy response. We conjecture that bank lending standards, as represented by LTV ratios, are thus part of the FED's reaction function.

Second, to isolate the impact of systematic monetary policy in the transmission of an LTV shock to the broader economy, we rely on policy counterfactuals as proposed in Bernanke et al. (1997) and Sims and Zha (2006), and recently applied in, e.g., Kilian and Lewis (2011) and Bachmann and Sims (2012). This methodology consists of creating a hypothetical economy, for which we "shut down" the FED's interest rate reaction to an LTV shock. By generating hypothetical sequences of *exogenous* monetary policy surprises that completely offset the *endogenous* Federal Funds rate response, the policy instrument remains constant over the impulse response horizon. Differences between the original and the counterfactual economy, then, indicate the quantitative importance of the systematic monetary policy tightening. We find that the positive non-residential investment response is magnified by the passive monetary policy stance. More importantly, however, with a counterfactually fixed interest rate, residential investment exhibits a quite persistent *increase*, peaking at around 0.4 percent after a year, and deviates statistically significant from the original economy from quarter three onwards. The systematic monetary policy response, hence, determines residential investment activity not only quantitatively, but also qualitatively. This strong interest rate sensitivity of residential investment is in line with, e.g., Erceg and Levin (2006), Monacelli (2009), and Calza et al. (2013).

These counterfactual fixed interest rate results also shed some light on the Great Recession. While our VAR is a linear tool, our results are in line with the perception that a tightening of LTV ratios may have exacerbated the downturn in housing markets at the onset of the Great Recession (see also Guerrieri and Iacoviello, 2015). The reason is the asymmetry represented by the zero lower bound on nominal interest rates. Historically, the FED would have lowered interest rates in the face of the LTV tightening, however, with interest rates bounded at zero, this cushioning mechanism had to be absent. According to our fixed interest rate results, such a situation should then be associated with a drop in residential investment, which was indeed observed during the financial crisis.

We also analyze the LTV shock propagation to measures of firm and household debt. From a theoretical perspective, collateral constraints on household borrowing represent the backbone of models that integrate durable housing goods into the dynamic stochastic general equilibrium (DSGE) framework.[4] Following the mechanism proposed in Kiyotaki and Moore (1997), household borrowing in these models is endogenously tied to a fraction of the (expected) housing value, where the down payment rate is pinned down by an exogenously fixed parameter, the LTV ratio. We find that following a loosening of bank lending conditions for residential mortgages, which we interpret as a loosening of bank lending conditions more broadly, *firms* increase their debt levels, measured either by total bank loans or mortgage loans. This propensity to leverage is, perhaps surprisingly, not affected by the monetary policy reaction. We infer that for firms, the quantity restriction on loans, i.e., the LTV ratio, dominates the price effects on short-term and mortgage interest rates, which are influenced by monetary policy.

In contrast and resembling the evidence for residential investment, the evolution of *household* debt is contingent on the Federal Funds rate reaction. LTV shocks have a small (negative) impact on household debt even as the Federal Funds rate and mortgage rate rise. This is in line with Justiniano et al. (2015), who find in a DSGE model for the US that exogenous shifts in LTV ratios do not appear to have a strong impact on leverage.[5] Household debt, however, increases under a counterfactually fixed interest rate policy, making the shock transmission through the monetary policy instrument, i.e., the systematic interest rate reaction, the crucial channel of how LTV shocks affect household debt.

We interpret shifts in residential LTV ratios as a *supply* indicator of banks' (mortgage) lending. This conclusion is supported by an exercise in the spirit of Bassett et al. (2014), which consists of removing influences of financial sector and macroeconomic conditions from the *raw* LTV series that might drive lending standards,

[4]See, among others, Iacoviello (2005), Monacelli (2009), and Iacoviello and Neri (2010).

[5]Similarly, Midrigan and Philippon (2016) argue that monetary policy appears to counteract shocks to household debt outside of zero lower bound episodes.

but at the same time might also affect the *demand* for residential mortgage loans. We find that, first, the impact of such factors on the LTV ratio is rather small and for the majority of control variables insignificant, and, second, the macroeconomic consequences of shocks to the *purged* LTV ratio are very similar to specifications in which we employ the raw LTV series.

Finally, the residential mortgage LTV ratio, as measured by the FHFA survey, captures all home owners, i.e., first-time as well as repeated home buyers. While Mian and Sufi (2011) show that *existing home owners* contributed substantially to the most recent leverage cycle, as about 65 percent of households already owned a property prior to the cycle, Duca et al. (2011, 2013) stress the importance of *first-time home owners* as being particularly subject to collateral constraints. These authors provided us with a cyclically adjusted LTV series only for the group of first-time home buyers based on the American Housing Survey (AHS). Although the results are – due to noise in first-time home buyer data – not quite as clean, the qualitative evidence for first-time home buyers is very close to the results for all home owners.

The remainder of the chapter is structured as follows. Section 2.2 describes the data, explains the empirical strategy, and presents the core empirical findings. Section 2.3 proposes possible extensions and reviews the results along some robustness dimensions. Section 2.4 concludes the chapter.

2.2 LTV shocks and monetary policy

This section presents the methodological framework and our main empirical findings. Section 2.2.1 describes the data. Section 2.2.2 discusses the VAR identification strategy and presents the main macroeconomic effects of LTV shocks. Section 2.2.3 characterizes the systematic monetary policy response to LTV shocks in detail and isolates the policy reaction in a counterfactual analysis. Section 2.2.4 analyzes the role of household and firm debt.

2.2.1 Data

We study the effects of putatively autonomous movements in residential mortgage LTV ratios on aggregate economic, in particular, investment activity, and monetary policy. Accordingly, our parsimonious benchmark model comprises four variables at the quarterly frequency: non-residential investment (i_t^{nr}), residential investment (i_t^r), the LTV ratio (ltv_t), and the nominal Federal Funds rate (r_t). We obtain the two investment series from Bureau of Economic Analysis (BEA) in seasonally adjusted real terms and take the natural logarithm. The monetary policy instrument is the quarterly average of the effective Federal Funds rate. The sample covers the period 1973Q1 to 2008Q4, where the availability of LTV data dictates the start of the sample. We confine the sample to 2008Q4 when the FED's policy instrument reached the zero lower bound. Since then, the FED engaged in several unconventional policies so that historical policy reaction functions are likely to no longer be valid during the financial crisis episode (see, e.g., Kilian and Lewis, 2011; Peersman and Wagner, 2015).

Our benchmark LTV measure is the quarterly average of the seasonally adjusted monthly LTV ratios on conventional mortgage loans from the Monthly Interest Rate Survey (MIRS) conducted by the Federal Housing Finance Agency (FHFA), which provides the most extensive data on terms and conditions of US mortgages. For instance, towards the end of our sample, the survey covers roughly 82,000 loan contracts.[6] The survey polls a sample of mortgage lenders (savings associations, commercial banks, and mortgage companies) to report interest rates and conditions on all fully amortized single family loans closed within the last five business days of each month.[7] As part of the survey, mortgage lenders are asked to report the agreed LTV ratios at purchase of the properties. Importantly, these LTV ratios include all types of home owners, i.e., owner occupiers as well as first-time home buyers. According to Mian and Sufi (2011), existing home owners contributed substantially to

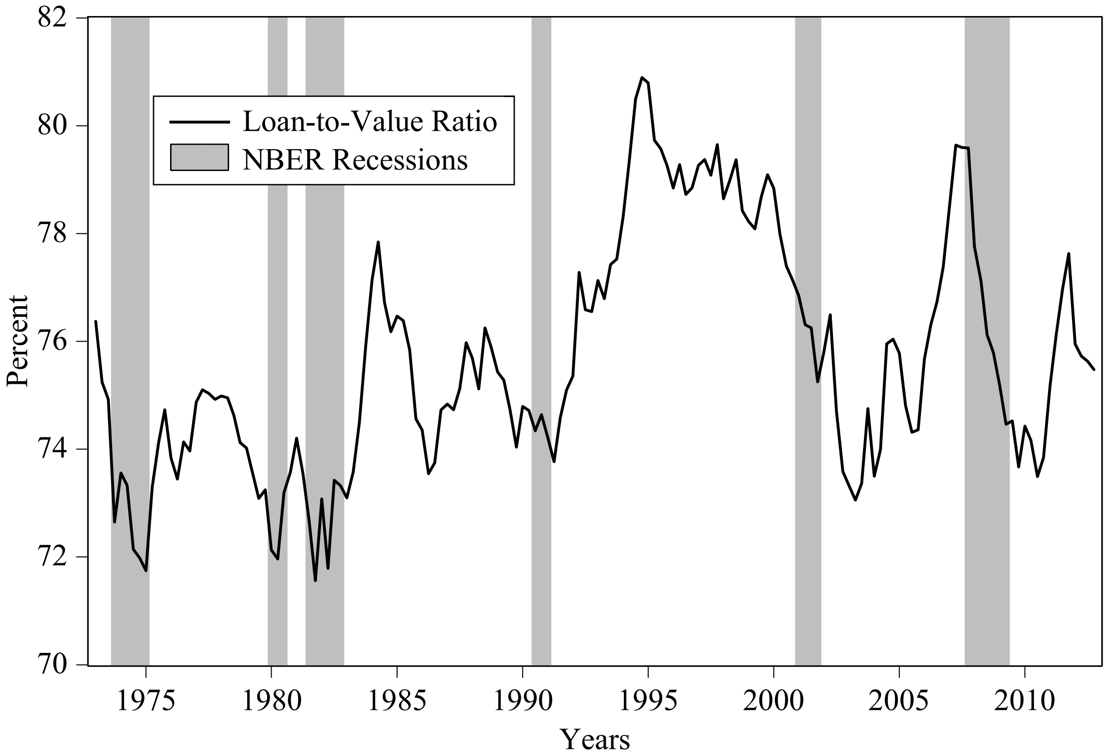
[6]See, e.g., page 6 of `www.reginfo.gov/public/do/DownloadDocument?objectID=19707402`.

[7]The survey does not comprise the following loan types: mortgages insured by the Federal Housing Administration or guaranteed by the Department of Veterans Affairs, multifamily mortgages, mortgages for mobile homes or farms, and mortgages created by refinancing existing mortgages.

the buildup in household leverage during the 2002 to 2006 house price acceleration.[8]

Figure 2.1 plots the FHFA LTV series, i.e., the average ratio of granted mortgage loans for single family houses and the underlying property prices multiplied by 100, across time. The shaded areas represent NBER-dated recession episodes in the US. The LTV ratio is procyclical and exhibits pronounced swings. Borrowing limits eased during the housing boom of the years 2002 to 2006, even though LTV ratios did not reach the levels seen at the end of 1994. This is partly driven by existing home owners, which used their increased housing wealth to roll over into new properties with lower average LTV ratios. At the onset of the Great Recession LTV ratios tightened sharply.

Figure 2.1: Loan-to-value ratio of residential mortgage loans



Notes: The figure displays the seasonally adjusted average loan-to-value ratio on conventional single family mortgage loans, which we obtain from the Federal Housing Finance Agency. Data are at the quarterly frequency, and we express them in percent, i.e., as a ratio of the granted mortgage loan and the underlying house price multiplied by 100. The shaded areas represent NBER-dated recession episodes in the US.

[8]We analyze first-time home buyer data in Section 2.3.2.

In what follows, we augment the baseline model along three dimensions:[9] first, to analyze the interest rate pass-through of the Federal Funds rate, we add the nominal contract mortgage rate (r_t^m) on existing single family home purchases provided by the FHFA and a measure of consumers' interest rate expectations (r_t^e), which we obtain from the Michigan Survey of Consumers, to the VAR. Second, to characterize the monetary policy response in detail, we allow for a more conventional monetary policy reaction function by including real GDP, y_t, and consumer price inflation, π_t. Third, to study the propagation of LTV shocks through measures of firm and household debt, we either use total bank-provided loans to non-financial businesses, b_t^f, and total household debt, b_t^h, or, alternatively and more specifically, firm and household mortgages, b_t^{fm} and b_t^{hm}. All debt series are stock variables measuring the outstanding amount of loans at the end of each quarter, where we apply the GDP deflator to transform them in real terms (see Justiniano et al., 2015).[10]

2.2.2 Identification of LTV shocks

Structural VAR

To analyze the macroeconomic consequences of exogenous shifts in LTV ratios, we rely on the vector autoregressive framework. A structural representation of the variables of interest can be formulated as

$$\boldsymbol{A_0}\boldsymbol{x}_t = \sum_{l=1}^{p} \boldsymbol{A}_l \boldsymbol{x}_{t-l} + \boldsymbol{\varepsilon}_t, \tag{2.1}$$

where we drop the intercept without loss of generality for notational convenience. $\boldsymbol{A}_l$ is a $n \times n$ matrix including autoregressive coefficients at lag, $l = 1, ..., p$, and $\boldsymbol{A}_0$

[9]Each extension is defined and described relative to the baseline four-variable VAR.

[10]We use the BEA investment and GDP series from NIPA Table 1.1.3., lines 1, 9, and 13. From the FRED database, we obtain the Federal Funds rate (FEDFUNDS), the GDP deflator (GDPDEF), and the quarterly change of the Consumer Price Index (CPIAUCSL). The debt measures are from the Flow of Funds database with identifiers: Z1/Z1/FL144123005.Q for non-financial business loans, Z1/Z1/FL143165005.Q for non-financial business mortgages, Z1/Z1/LA153165105.Q for home mortgages, and Z1/Z1/LA153166000.Q for consumer credit of households and nonprofit organizations. The LTV series and our measure of mortgage rates are from the MIRS, Table 17 (all homes), which can be downloaded from FHFA: `http://www.fhfa.gov/DataTools/Downloads/Pages/Monthly-Interest-Rate-Data.aspx`. We apply the Census X-12 filter to seasonally adjust the LTV series and those debt series that are not seasonally adjusted in the Flow of Funds database.

captures contemporaneous impact coefficients. p is the lag length, and $\boldsymbol{\varepsilon}_t$ represents mutually uncorrelated structural shocks. The $n \times 1$ vector $\boldsymbol{x}_t$ comprises the following n variables in this order, $\boldsymbol{x}_t = [i_t^{nr} \; i_t^{r} \; \Delta ltv_t \; r_t]'$.

We need to restrict elements in $\boldsymbol{A}_0$, to disentangle exogenous LTV movements from endogenous reactions to other variables in $\boldsymbol{x}_t$, i.e., to uniquely recover the structural VAR. Structural LTV shocks could arise from internal reassessments of the quality of borrowers, new business models, or shifts in the supervisory and regulatory environment under which banks operate (see Bassett et al., 2014). We follow Gilchrist and Zakrajsek (2012) and Walentin (2014) by assuming that shocks in "slow-moving" macroeconomic variables (i_t^{nr}, i_t^{r}) impact financial variables (Δltv_t, r_t) contemporaneously, whereas shocks in "fast-moving" financial variables affect the real economy with a time lag (see also Christiano et al., 1996; Peersman and Wagner, 2015). We implement the identification strategy by applying a Cholesky factorization to the variance-covariance matrix of the reduced form regression residuals, $\boldsymbol{u}_t$. Then we use the Cholesky factor for $\boldsymbol{A}_{\boldsymbol{0}}$, which delivers the linear mapping $\boldsymbol{u}_t = \boldsymbol{A}_{\boldsymbol{0}}{}^{-1}\boldsymbol{\varepsilon}_t$ and recovers the structural representation. Within the recursive identification scheme, we allow LTV shocks, $\varepsilon_{L,t}$, to simultaneously cause shifts in the monetary policy instrument, where the subscript L is the position of Δltv_t in $\boldsymbol{x}_t$. However, results are not sensitive to, e.g., ordering the Federal Funds rate, r_t, before Δltv_t, which prevents an immediate impact of LTV shocks on monetary policy (see Section 2.3.4).

As in Gilchrist and Zakrajsek (2012) and Bassett et al. (2014), we estimate the VAR with two lags – a lag length suggested by both the Schwarz and Hannan-Quinn information criterion. Results are, however, robust to higher lag orders (see Section 2.3.4). i_t^{nr} and i_t^{r} enter the VAR as natural logarithms (multiplied by 100), and we measure r_t in percent. For the LTV ratio, the null hypothesis of a unit root cannot be rejected based on the augmented Dickey-Fuller test. Therefore, we include ltv_t in first differences explicitly allowing for a unit root and thus permanent movements in LTV ratios following the shock, whereas the event of including ltv_t in levels and estimating a root equal to 1.0 would have a probability of zero (e.g., Born et al., 2015).[11] To illustrate the dynamics of LTV ratios in the VAR, we present cumulative impulse responses for this variable, which we can interpret as LTV ratio changes in percentage points.

LTV shocks: empirical evidence

Figure 2.2 traces out the impulse responses of the variables in $\boldsymbol{x}_t$ following an exogenous 25 basis point increase in the LTV ratio. We present a 25 basis point LTV shock instead of a one standard deviation shock for better comparability across specifications. Notice that this shock size is frequently used in monetary policy VARs. A one standard deviation shock to the LTV ratio would amount to 74 basis points, while a monetary policy shock has a standard deviation of 89 basis points in the benchmark model. Thus the impulse of our LTV shock is of similar strength as a conventional monetary policy shock. The solid lines display the point estimates of impulse response functions and the shaded areas are one standard error confidence intervals, which we obtain from 5,000 replications of the recursive-design wild bootstrap procedure of Goncalves and Kilian (2004).

The LTV ratio exhibits a small and sluggish increase before levelling off at the 25 basis point level, i.e., the exogenous shock has a very persistent effect on the LTV ratio. The shock significantly affects non-residential investment, which features a hump-shaped increase with a peak around 0.3 percent after one year, and then reverts back to the pre-shock level. Because of these spillovers to non-residential investment (and GDP, as we will show) we view residential mortgage LTV ratios as an indicator of banks' lending propensity in a broader sense rather than just being indicative of lending standards in residential mortgage markets.

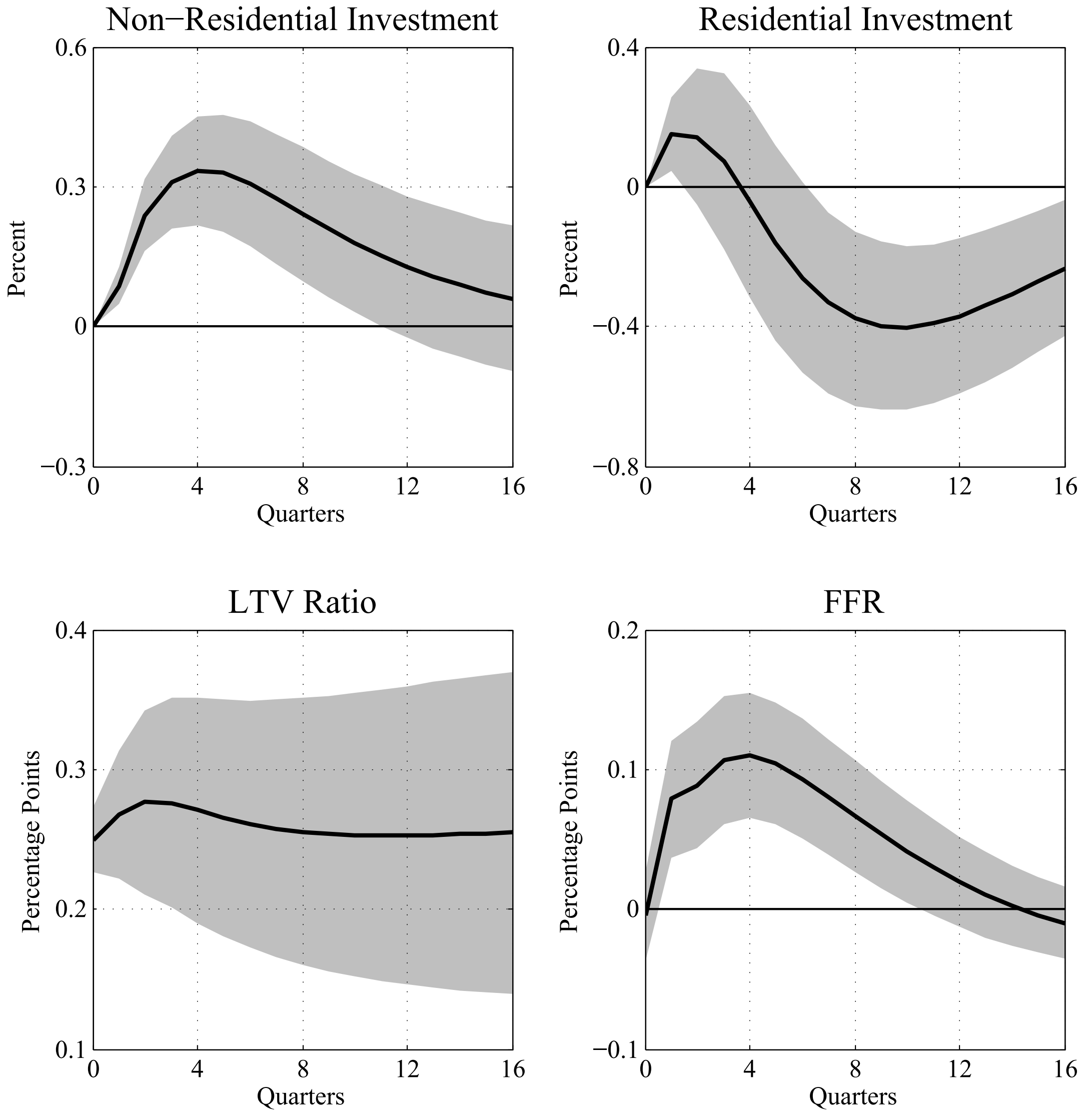
In contrast, the impulse response of residential investment rises by 0.15 percent in the first quarter, but then falls significantly by 0.4 percent until it reaches its trough after two and a half years before slowly reverting back to its pre-shock level.

[11] The null of a unit root, likewise, cannot be rejected for residential investment. We carefully deal with this issue in Section 2.3.4 and find similar results for specifications, in which all variables enter the VAR in levels or in which all variables − except for interest rates − enter the VAR as first differences. In addition, we use population-standardized investment in levels, which somewhat alleviates concerns about a unit root in residential investment and find almost identical results.

This result is perhaps surprising as it is inconsistent with the view that loose LTV ratios lead to construction booms, and, perhaps, housing bubbles. But why does a shock that eases borrowing constraints in the residential mortgage market coincide with a slowdown of residential investment? The impulse response in the lower right panel of Figure 2.2 represents a candidate answer. Monetary policy reacts to the eased lending standards by significantly raising the Federal Funds rate by more than 10 basis points. This finding is consistent with Bassett et al. (2014), who report a significant reaction of monetary policy after a shock to their broadly defined indicator of banks' loan supply, and with Walentin (2014), who finds a monetary policy easing after contractionary mortgage spread shocks.

The persistent contractionary shift in monetary policy counteracts the initial easing in mortgage markets and seems to be dominating the expansionary effects of the LTV increase, at least, with respect to residential investment. The literature (e.g., Erceg and Levin, 2006; Monacelli, 2009; Calza et al., 2013) supports this hypothesis by documenting a strong interest rate sensitivity of consumer durables and residential investment, for which the impact of monetary policy shocks is several times larger compared to non-housing related GDP components.

Figure 2.2: Loan-to-value ratio shock in the benchmark 4-variable VAR

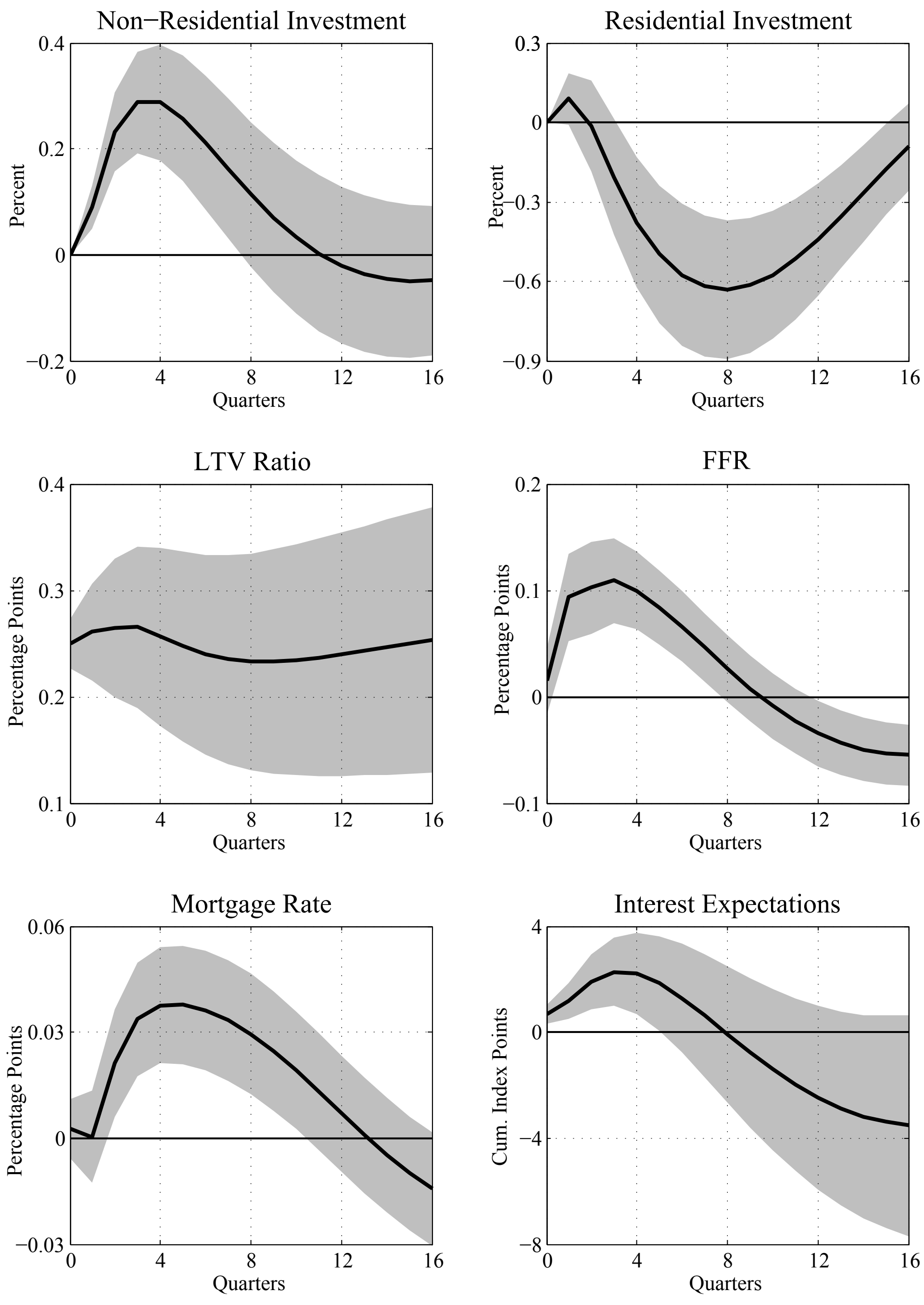


Notes: The x-axis represents time in quarters. The solid lines represent point estimates of impulse response functions for the VAR, using $\boldsymbol{x}_t = [i_t^{nr} \; i_t^r \; \Delta ltv_t \; r_t]'$. Shaded areas display one standard deviations confidence intervals, which we obtain from 5,000 replications of the recursive-design wild bootstrap procedure of Goncalves and Kilian (2004).

We examine the monetary policy reaction hypothesis along two further dimensions. First, we analyze whether the endogenous interest rate hike is part of the information set of households, i.e., whether housing investors anticipate the adverse interest environment, and, second, we study whether the shift in monetary policy passes through to interest rates that are more relevant for housing markets, i.e., mortgage rates. To do so, we add the nominal contract mortgage rate (r_t^m) on existing single family home purchases provided by the FHFA to the VAR. Furthermore, we include a measure of consumers' interest rate expectations (r_t^e), which we obtain from the Michigan Survey of Consumers. On a monthly basis, the survey asks consumers the following question: *"No one can say for sure, but what do you think will happen to interest rates for borrowing money during the next 12 months – will they go up, stay the same, or go down?"*. We use a balance score, i.e., the share of consumers expecting rates to go up minus the share of consumers expecting rates to go down, plus 100. Thus the scale is qualitative and positive values indicate a less favorable expected interest rate environment. We re-estimate the VAR with these additional variables ordered as follows $\boldsymbol{x}_t = [i_t^{nr} \; i_t^r \; \Delta ltv_t \; r_t \; r_t^m \; r_t^e]'$. The recursive ordering allows the Federal Funds rate to pass through to mortgage rates contemporaneously (Bernanke et al., 1997). It also allows expectations to adjust to macroeconomic and financial conditions on impact. Figure 2.3 presents the LTV shock propagation to the newly introduced variables in the lowest panels. We report the cumulative impulse response of interest expectations to recover the qualitative expectations on the *level* of interest rates.

The endogenous monetary policy tightening transmits significantly to mortgage rates. Thus an increase in mortgage borrowing costs (prices) counteracts the loosening of LTV ratios on mortgage loans (quantities). The policy reaction, in addition, is reflected by consumers' qualitative expectations on borrowing conditions, which move instantaneously and remain significantly positive for more than a year. The evidence of both variables supports the perception of systematic monetary policy being a candidate for explaining the decrease in residential investment after an expansionary LTV shock.

Figure 2.3: Mortgage rate and interest expectations following an LTV shock



Notes: The x-axis represents time in quarters. The solid lines represent point estimates of impulse response functions for the VAR, using $\boldsymbol{x}_t = [i_t^{nr} \; i_t^r \; \Delta ltv_t \; r_t \; r_t^m \; r_t^e]'$. Shaded areas display one standard deviations confidence intervals, which we obtain from 5,000 replications of the recursive-design wild bootstrap procedure of Goncalves and Kilian (2004).

2.2.3 Systematic monetary policy

This section studies the systematic monetary policy reaction in detail, first, by isolating the drivers of the Federal Funds rate response, and, second, by performing a counterfactual analysis meant to illustrate the quantitative importance of the monetary policy response.

What drives the monetary policy reaction?

Before we isolate the effect of the systematic monetary policy reaction, it is instructive to analyze which variables in the VAR actually trigger the policy reaction, i.e., what is the central bank responding to after an LTV shock. We do so by decomposing the impulse response of the FED's policy instrument into contributions from the variables in $\boldsymbol{x}_t$, as in Kilian and Lewis (2011). The rationale behind this exercise is as follows: LTV disturbances cause the Federal Funds rate to deviate from its steady state. This response can be considered as the sum of a policy reaction, first, to lags of the policy instrument itself, and, second, to deviations of other measures in $\boldsymbol{x}_t$ from their steady state values. The relative contributions of variables in $\boldsymbol{x}_t$ to the Federal Funds rate response, then, identify the forces underlying the monetary policy contraction.

It is convenient to express the structural VAR as follows

$$\boldsymbol{x}_t = \boldsymbol{C}\boldsymbol{x}_t + \sum_{l=1}^{p} \boldsymbol{A}_l \boldsymbol{x}_{t-l} + \boldsymbol{\varepsilon}_t, \tag{2.2}$$

where the $n \times n$ matrix $\boldsymbol{C}$ is strictly lower triangular. Furthermore, we can compactly summarize the structural parameters as $\boldsymbol{B} = [\boldsymbol{C} \;\; \boldsymbol{A_1} \ldots \boldsymbol{A_p}]$.

To isolate the contribution of variable j to the Federal Funds Rate response at horizon h after a time $t = 0$ shock to the LTV ratio ($\boldsymbol{\Xi}_{F,j,h}$), we define

$$\boldsymbol{\Xi}_{F,j,h} = \sum_{m=0}^{\min(p,h)} \boldsymbol{B}_{F,mn+j} \boldsymbol{\Phi}_{j,L,h-m}, \tag{2.3}$$

with subscripts F and L denoting the position of the Federal Funds rate and LTV ratio in the system, and $h = 0, 1, 2, ..., 16$ as well as $j = 1, 2, ..., n$. $\boldsymbol{\Phi}_{j,L,h-m}$ is the $\{j, L\}$ entry of the parameter matrix of impulse reponses, $\boldsymbol{\Phi}_{h-m}$.

Given the FED's objective of macroeconomic stabilization and taking its "dual mandate" into account, we augment the benchmark model for the impulse response decomposition exercise to allow for a more conventional monetary policy reaction function (e.g., Bernanke et al., 1997; Kilian and Lewis, 2011), i.e., we add real GDP, y_t, and consumer price inflation, π_t, to the VAR, and study the contributions of these variables to the policy response. The augmented model thus includes the following six variables in the following order $\boldsymbol{x}_t = [y_t \; i_t^{nr} \; i_t^{r} \; \pi_t \; \Delta ltv_t \; r_t]'$. With this identification restriction, monetary policy surprises impact other variables with a time lag of one quarter, and monetary policy reacts to realizations of macroeconomic aggregates contemporaneously, i.e., contemporaneous as well as previous realizations of all variables in $\boldsymbol{x}_t$ are reflected by the FED's time t information set. We thus follow an established literature of recursively identified monetary policy VARs, e.g., Bernanke et al. (1997), Christiano et al. (2005), and Erceg and Levin (2006).[12]

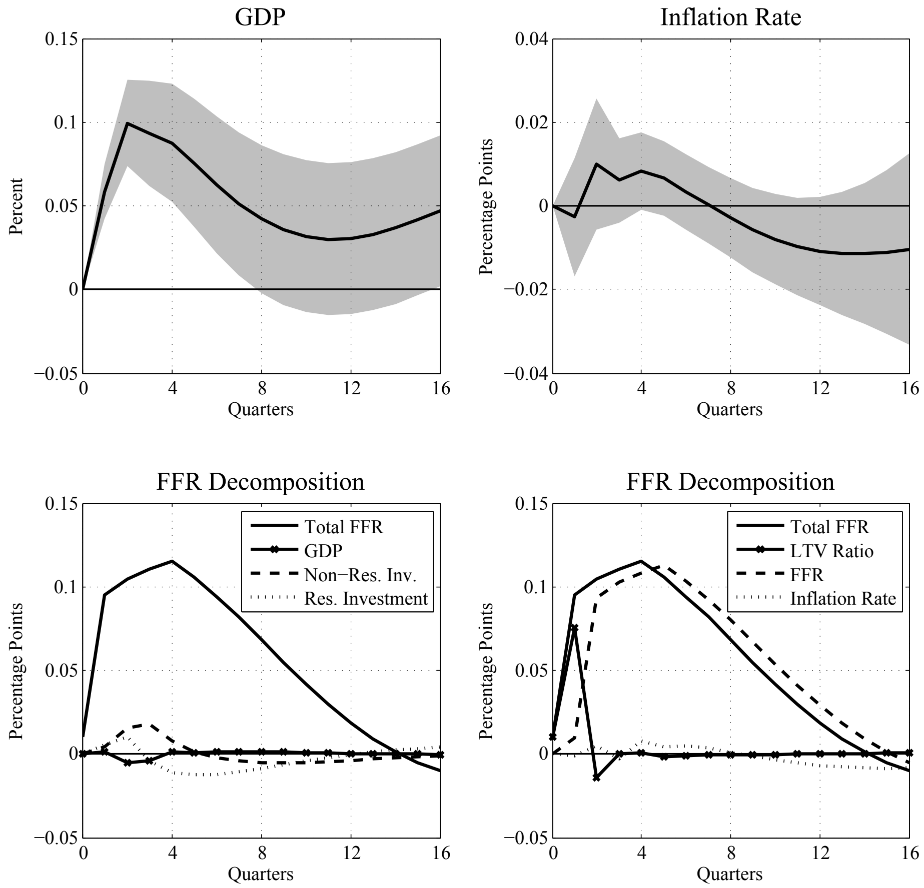
The upper panels of Figure 2.4 plot the dynamics of y_t and π_t after the 25 basis point disturbance to the LTV ratio.[13] Real GDP displays a hump-shaped rise of 0.1 percent, which peaks in quarter two and becomes insignificant from quarter eight onwards. The inflation rate is initially sticky and then increases for six quarters before moving into negative territory from quarter seven onwards. However, the inflation response is economically small and statistically insignificant over the whole forecast horizon. This suggests that the monetary policy contraction (solid lines in the lower panels of Figure 2.4) cannot be explained by a "leaning against the wind" of inflationary pressure. Indeed, the lower panels of Figure 2.4 reject such a narrative as they do with a "curbing the output boom" story: the endogenous monetary policy contraction is a *direct* response to the LTV shock, rather than an *indirect* response operating through other variables, in particular, output or inflation in the VAR. In the first quarter after the shock, the LTV ratio accounts for the bulk

[12] An identified contractionary monetary policy shock in our six-variable VAR features negative hump-shaped responses of y_t, i_t^{nr} and i_t^{r}. The response of i_t^{r} is roughly three times as strong as that of i_t^{nr}, that is, residential investment is significantly more interest-sensitive than non-residential investment. ltv_t significantly and sluggishly falls after a contractionary monetary policy disturbance, suggesting a "risk-taking channel" (on the part of lenders) of monetary policy, that is, after a monetary contraction, banks grant less or less risky mortgage loans. π_t reacts positively for four quarters (price puzzle) before turning negative. Results are available upon request.

[13] Figure 2.5 shows the rest of the impulse response functions.

of the Federal Funds rate response (line with nodes) and for subsequent horizons, the lags of Federal Funds rate itself explain the Federal Funds rate response almost entirely (dashed line). The contributions of output, inflation, and both investment measures are negligible. Apparently, lending standards in the banking industry as reflected by residential mortgage LTV ratios are part of the FED's reaction function, and a move against more expansionary lending practices drives the policy instrument following the LTV shock.

Figure 2.4: Decomposition of the Federal Funds rate response following an LTV shock



Notes: The x-axis represents time in quarters. In the upper panels, the solid lines represent point estimates of impulse response functions for y_t and π_t from the VAR with $\boldsymbol{x}_t = [y_t \ i_t^{nr} \ i_t^{r} \ \pi_t \ \Delta ltv_t \ r_t]'$, and shaded areas display one standard deviations confidence intervals, which we obtain from 5,000 replications of the recursive-design wild bootstrap procedure of Goncalves and Kilian (2004). The solid lines in the lower panels are the point estimate of the Federal Funds rate impulse response function after an LTV shock. The dashed, dotted, and lines with nodes represent the contribution of the respective variable to the reaction of the FED's policy instrument.

Shutting down the monetary policy response

To flesh out the impact of the Federal Funds rate reaction, we follow the methodology of creating policy counterfactuals as proposed in Bernanke et al. (1997) and Sims and Zha (2006), and recently applied in, e.g., Kilian and Lewis (2011) and Bachmann and Sims (2012). The procedure consists of analyzing the dynamics of variables in the VAR model for a hypothetical economy, for which we completely remove the impact of the LTV shock from the FED's reaction function, i.e., where the FED does not react to the LTV shock and its effects at all. To create such an environment, we generate hypothetical sequences of monetary policy shocks that suffice to completely "shut down" the Federal Funds rate response after the LTV shock. The counterfactual economy features the same structural characteristics as the benchmark economy, however, in the face of the LTV shock, the former economy is subject to a sequence of counteracting *exogenous* monetary policy disturbances that "zero out" the *endogenous* policy response.

We can recursively calculate the monetary policy shocks required to force the policy response to zero over the whole forecast horizon as follows:

$$\boldsymbol{\varepsilon}_{F,h} = -\sum_{j=1}^{n} \boldsymbol{B}_{F,j}\boldsymbol{y}_{j,F} - \sum_{m=1}^{\min(p,h)} \sum_{j=1}^{n} \boldsymbol{B}_{F,mn+j}\boldsymbol{z}_{j,h-m}. \tag{2.4}$$

$\boldsymbol{y}_{j,0}$ is the time $t = 0$ impact of the LTV disturbance on variable j in the benchmark VAR, whereas the same impact in the counterfactual economy reads:

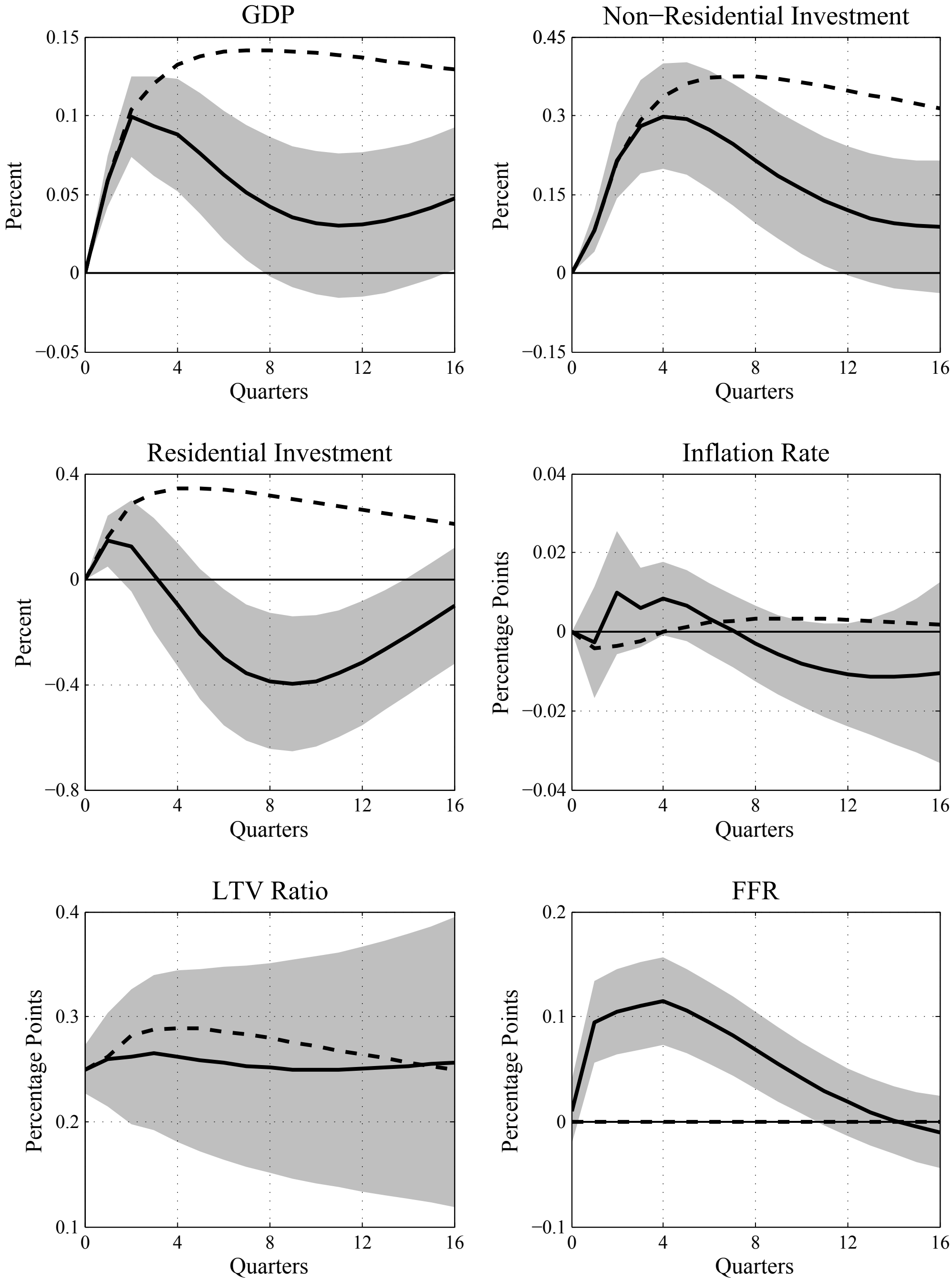
$$\boldsymbol{z}_{j,0} = \boldsymbol{y}_{j,0} + \frac{\boldsymbol{\Phi}_{j,F,0}\boldsymbol{\varepsilon}_{F,0}}{\sigma_F}. \tag{2.5}$$

The standard deviation of the monetary policy disturbance is σ_F. For horizons beyond the impact period, $h > 0$, we calculate

$$\boldsymbol{y}_{j,h} = \sum_{m=1}^{\min(p,h)} \sum_{i=1}^{n} \boldsymbol{B}_{j,mn+i}\boldsymbol{z}_{j,h-m} + \sum_{i<j}^{n} \boldsymbol{B}_{j,i}\boldsymbol{y}_{i,h} \text{ and } \boldsymbol{z}_{j,h} = \boldsymbol{y}_{j,h} + \frac{\boldsymbol{\Phi}_{j,F,0}\boldsymbol{\varepsilon}_{F,h}}{\sigma_F}. \tag{2.6}$$

The solid lines in Figure 2.5 show the impulse responses of the variables in $\boldsymbol{x}_t$ after an LTV shock together with one standard error confidence intervals (shaded area) for the original economy. The dashed lines represent impulse response functions for a counterfactual economy, in which monetary policy does not respond to the dynamics triggered by the LTV shock at any horizon.

Figure 2.5: LTV shock and a policy counterfactual



Notes: The x-axis represents time in quarters. The solid lines represent point estimates of impulse response functions for the VAR, using $\boldsymbol{x}_t = [y_t \; i_t^{nr} \; i_t^{r} \; \pi_t \; \Delta ltv_t \; r_t]'$. Shaded areas display one standard deviations confidence intervals obtained from 5,000 replications of the recursive-design wild bootstrap procedure of Goncalves and Kilian (2004). The dashed lines represent counterfactual impulse responses for the case of a passive monetary policy authority that does not react to the shock at all, as in Bernanke et al. (1997) and Sims and Zha (2006).

By construction, the impulse response of r_t is zero over the whole time horizon in the counterfactual experiment, and, as a consequence of passive monetary policy, ltv_t features a slightly stronger and more persistent increase compared to the benchmark economy. In a statistical sense, however, the counterfactual response of ltv_t is not different from the benchmark response. GDP, y_t, and non-residential investment, i_t^{nr}, both increase more strongly and more persistently in the absence of the policy tightening, while the response of the inflation rate, π_t, remains rather flat. The dynamics of residential investment, i_t^r, are most affected by the altered monetary policy regime. In the counterfactual economy, i_t^r continues the initial surge of the benchmark case by further increasing in a hump-shaped manner to almost 0.4 percent after one year. The response then remains strictly positive over the whole forecast horizon, whereas in the benchmark economy i_t^r turns significantly negative after one year. The effect of the LTV shock on residential investment thus crucially depends on the endogenous reaction of monetary policy, both in a quantitative and qualitative sense. This finding is consistent with the strong interest sensitivity of the housing sector documented in, e.g., Erceg and Levin (2006), Monacelli (2009), and Calza et al. (2013); see also Footnote 12 above.

Our results also suggest that an exogenous loosening of LTV ratios is unlikely to explain a boom in residential investment at least under conventional monetary policy. From a macroprudential policy perspective, the evidence represents a caveat for the use of supervisory limits on LTV ratios as a tool to curb overheating housing markets, again, at least in times of conventional monetary policy.[14] More generally, the efficacy of such policy measures seems to be contingent on the reaction function of monetary policy; macroprudential policy measures, therefore, should be designed to take into account interactions or coordinate with monetary policy.

[14]In the face of the Great Recession a number of countries introduced, tightened, or at least considered the introduction of supervisory limits for LTV ratios as a macroprudential policy tool. Among them are for instance, Canada, the Netherlands, Norway, Sweden, or the United Kingdom (IMF, 2011).

2.2.4 Household and firm debt

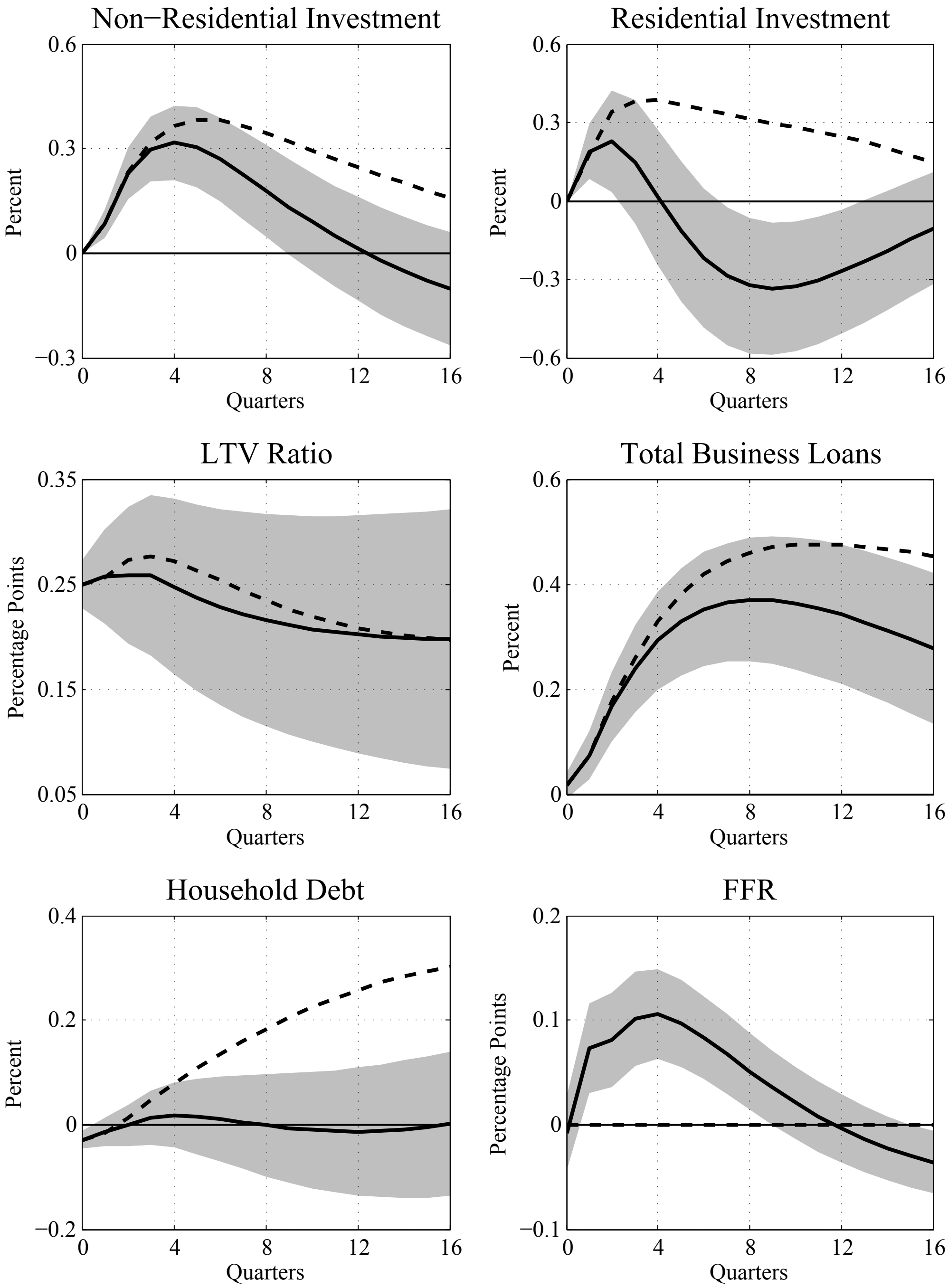
We next analyze the implications of looser collateral requirements on mortgage loans for the leverage of households and firms. We do this against the backdrop of the most recent US housing cycle of the years 2000 to 2007, where borrowing of households and firms increased substantially. In real terms, household debt rose by more than 70 percent and bank provided loans to non-financial businesses by 50 percent during this period. The unprecedented surge of private debt led to a number of theoretical contributions studying the interaction between leverage and the broader economy (see, e.g., Eggertsson and Krugman, 2012; Justiniano et al., 2015; Midrigan and Philippon, 2016). By using data on LTV ratios, the approach taken here adds a time series perspective on the role of changes in collateral requirements as a potential driver of leverage cycles to this literature, explicitly accounting for the role of monetary policy.

We follow Monacelli (2009) and use the natural logarithm (multiplied by 100) of real household debt, b_t^h, which consists of home mortgage loans and consumer credit provided by banks. For firms, we focus on bank provided loans to non-financial businesses, b_t^f. We re-estimate the VAR including the following variables in this order $\boldsymbol{x}_t = [i_t^{nr}\ i_t^r\ \Delta ltv_t\ b_t^f\ b_t^h\ r_t]'$. According to the maintained Cholesky identification strategy, LTV shocks move the newly introduced debt measures contemporaneously, and we allow monetary policy to respond to all financial variables on impact (see Gilchrist and Zakrajsek, 2012).

Figure 2.6 plots the impulse responses of the debt-augmented model, both for the original (solid line with confidence bands) and the counterfactual (dashed line) economy. The impulse responses of i_t^{nr}, i_t^r, ltv_t, and r_t are hardly affected by the introduction of business loans and household debt, compared to the benchmark four-variable VAR. Bank loans to the non-financial business sector display a pronounced increase, which is significant over the whole forecast horizon, i.e., firms quickly and persistently take advantage of the loosened availability of loans. In contrast, the impulse response function of household debt is even slightly negative in the impact quarter and otherwise barely reacts to the shock. Turning to the results for the counterfactual economy (dashed line) in which the FED remains passive after the LTV shock, no pronounced differences emerge for the evolution of business loans. Thus the *price* of loans – indirectly influenced by the systematic monetary policy tightening – appears to be second order for firms' propensity to borrow from banks, whereas the relaxation of the *quantity* restriction on loans, i.e., the LTV ratio, when interpreted as a broader indicator of loan supply, emerges as the dominating factor. Shutting down the monetary policy tightening, however, crucially alters the debt position of the household sector. For the counterfactual economy, household debt is slowly building up after an initial dip and is statistically different from the benchmark economy from the second year onwards. Restricting the analysis only to *mortgage* loans of households, b_t^{hm}, and firms, b_t^{fm}, as illustrated in Figure 2.7, reveals similar results, i.e., following a loosening of borrowing constraints, firms increase mortgage loans independently of the monetary policy response, whereas households reduce their mortgages in the historical experiment and strongly increase their mortgage leverage in the counterfactual economy. The strong interest rate sensitivity of household debt thus helps rationalize the differences in the counterfactual impulse responses for residential investment, which display the same qualitative behavior as household debt and mortgage dynamics.

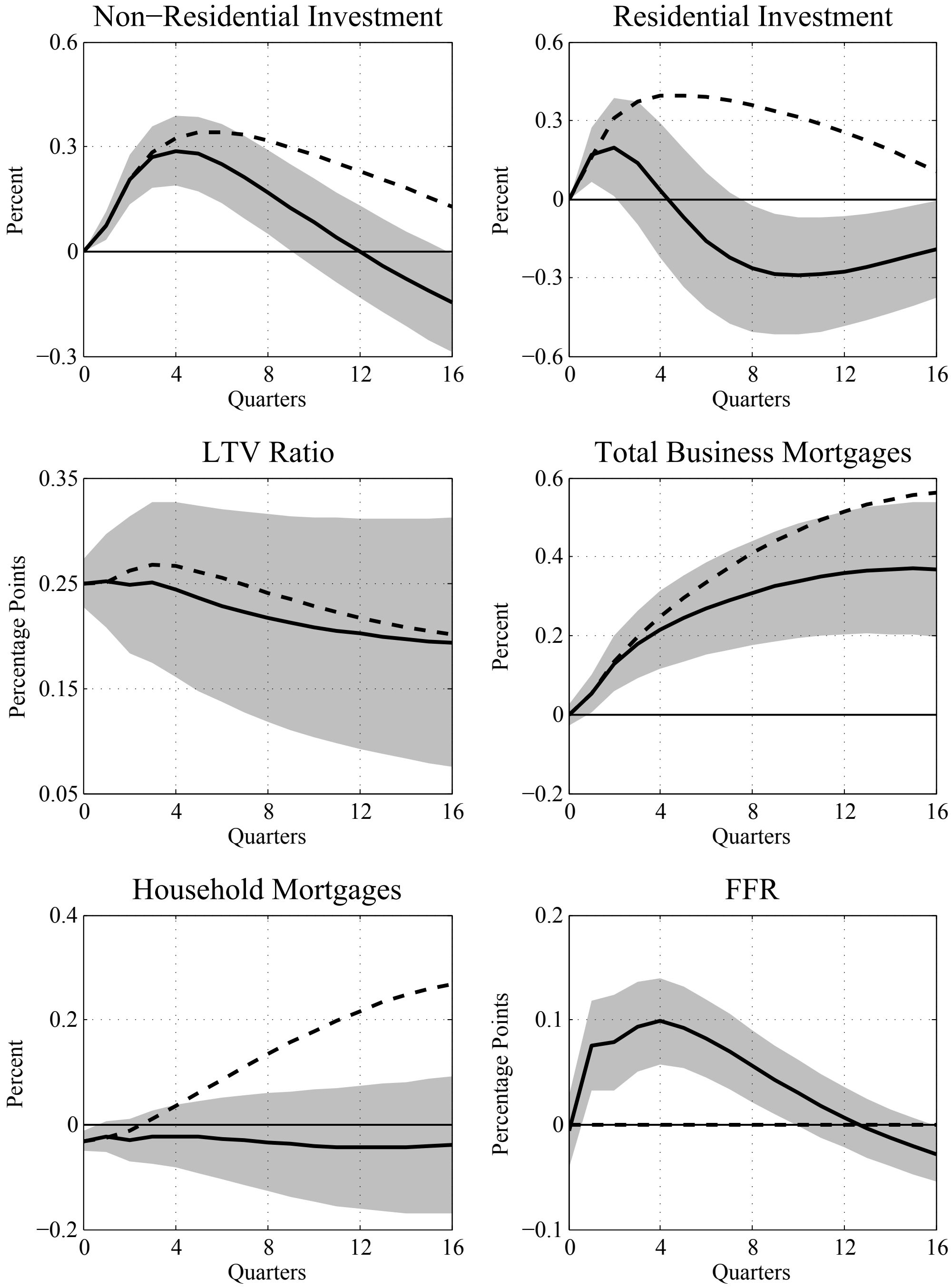
Lastly, while our VAR is a linear tool and we exclude the Great Recession quarters from our analysis, and while we find no significant role for changes in LTV ratios as a driver of residential investment and household debt – including the housing cycle from 2001 to 2007 –, our results support the view that a tightening of LTV ratios may have exacerbated the downturn in housing markets at the onset of the Great Recession. The reason is that the zero lower bound on nominal interest rates represents an asymmetry. Historically, the FED would have lowered interest rates in the face of the LTV tightening; however, with interest rates bounded at zero, this cushioning mechanism was absent. According to our fixed interest rate results, such a situation should then be associated with a drop in residential investment, which was indeed observed during the financial crisis.

Figure 2.6: LTV shock and debt of households and firms



Notes: The x-axis represents time in quarters. The solid lines represent point estimates of impulse response functions for the VAR, using $\boldsymbol{x}_t = [i_t^{nr}\ i_t^{r}\ \Delta ltv_t\ b_t^f\ b_t^h\ r_t]'$. Shaded areas display one standard deviations confidence intervals obtained from 5,000 replications of the recursive-design wild bootstrap procedure of Goncalves and Kilian (2004). The dashed lines represent counterfactual impulse responses for the case of a passive monetary policy authority that does not react to the shock at all, as in Bernanke et al. (1997) and Sims and Zha (2006).

Figure 2.7: LTV shock and mortgage debt of households and firms



Notes: The x-axis represents time in quarters. The solid lines represent point estimates of impulse response functions for the VAR, using $\boldsymbol{x}_t = [i_t^{nr} \; i_t^{r} \; \Delta ltv_t \; b_t^{fm} \; b_t^{hm} \; r_t]'$. Shaded areas display one standard deviations confidence intervals obtained from 5,000 replications of the recursive-design wild bootstrap procedure of Goncalves and Kilian (2004). The dashed lines represent counterfactual impulse responses for the case of a passive monetary policy authority that does not react to the shock at all, as in Bernanke et al. (1997) and Sims and Zha (2006).

2.3 Extensions and robustness

In this section, we scrutinize our main findings along the following dimensions. In Section 2.3.1, we examine further the validity of the benchmark LTV measure as a supply indicator of bank lending. Section 2.3.2 analyzes the effects of an LTV shock restricted to the group of first-time home buyers, and Section 2.3.3 provides results when we use an alternative procedure to create policy counterfactuals. Section 2.3.4 presents robustness checks on the VAR specification and identification.

2.3.1 LTV ratios: a loan "supply" indicator?

Thus far, we have interpreted Cholesky-identified shifts in LTV ratios as an indicator of changes in the effective mortgage loan *supply* of banks. However, given the potential endogeneity of the LTV ratio, such shifts could also, at least to some extent, reflect variations in loan *demand* of borrowers. Using data on changes of banks' lending standards from the FED's Senior Loan Officer Opinion Survey (SLOOS), Bassett et al. (2014) propose a procedure to purge loan supply measures from influences that, on the one hand, drive loan supply, but on the other hand, might independently affect loan demand. We apply their methodology to the raw LTV series and remove the effects of variables capturing (a) the current state of the economy, (b) the economic outlook, and (c) general financial sector conditions. We then re-run our six-variable VAR from Section 2.2.3 with the adjusted LTV measure.

Specifically, to control for changes in lending standards that are reflective of the *current state of the economy*, we follow Bassett et al. (2014) and account for the four quarter percentage change of real GDP, $y_t - y_{t-4}$, the four quarter percentage change of the unemployment rate, $u_t - u_{t-4}$, and the quarterly change in the real Federal Funds rate, Δrr_t. According to, e.g., the SLOOS evidence in Bassett et al. (2014), another important driver of changes in banks' lending practices is the *outlook about the future evolution of the economy*. Thus we purge the LTV series from the year ahead expectations on the growth rate of real GDP, $\mathbb{E}_{t-1}\{y_{t+4} - y_t\}$, and the expected change in the unemployment rate, $\mathbb{E}_{t-1}\{u_{t+4} - u_t\}$. Both expectation measures are available from the Survey of Professional Forecasters. Furthermore, we include the change in the term spread, $\Delta tspr_t$, which we measure as the slope of the yield curve for US Treasuries with constant maturity, i.e., the spread between three month and ten year Treasury yields. This spread captures expectations on the future evolution of policy rates. Finally, we control for the following indicators reflecting *financial sector conditions*. First, we include the change in the credit spread index, $\Delta cspr_t$, developed by Gilchrist and Zakrajsek (2012), which represents a corporate bond spread calculated on the basis of secondary market (individual) bond prices. The index serves as an indicator of tensions in financial markets as well as perceived default risks and is shown to have a good predictive capability for measures of real economic activity. Second, we use changes in the excess bond premium, Δebp_t, also proposed in Gilchrist and Zakrajsek (2012), to address potential movements in financial sector risk aversion and, third, we include the percentage change in private depository institutions' net worth, Δnw_t, to account for the influence of banks' capital position on lending policies.[15]

We run a regression of Δltv_t on the described set of control variables to purge the LTV ratio from these factors. We perform the estimation by ordinary least squares and report Newey-West standard errors in parentheses. The resulting equation is given by:

$$\begin{aligned}\Delta ltv_t = &\underset{(0.28)}{0.10} - \underset{(0.27)}{0.46}\mathbb{E}_{t-1}\{u_{t+4} - u_t\} - \underset{(0.30)}{0.42}(u_t - u_{t-4}) - \underset{(9.49)}{1.51}\mathbb{E}_{t-1}\{y_{t+4} - y_t\} \\ &- \underset{(0.11)}{0.08}(y_t - y_{t-4}) + \underset{(0.09)}{0.04}\Delta rr_t - \underset{(0.17)}{0.14}\Delta tspr_t + \underset{(0.0003)}{0.0006}\Delta nw_t - \underset{(0.49)}{0.15}\Delta ebp_t \\ &- \underset{(0.36)}{0.23}\Delta cspr_t + \varepsilon_t^{ltv},\end{aligned} \tag{2.7}$$

where the residual of the regression, ε_t^{ltv}, denotes the "cleaned up" LTV series. Except for changes in the expected unemployment rate and banks' net worth, none of the controls is significant at the 10 percent level, and with an adjusted R^2 of 0.098, the overall explanatory power of the regressors is weak. This supports the notion that the raw LTV series is a fairly clean measure of movements in banks' *loan*

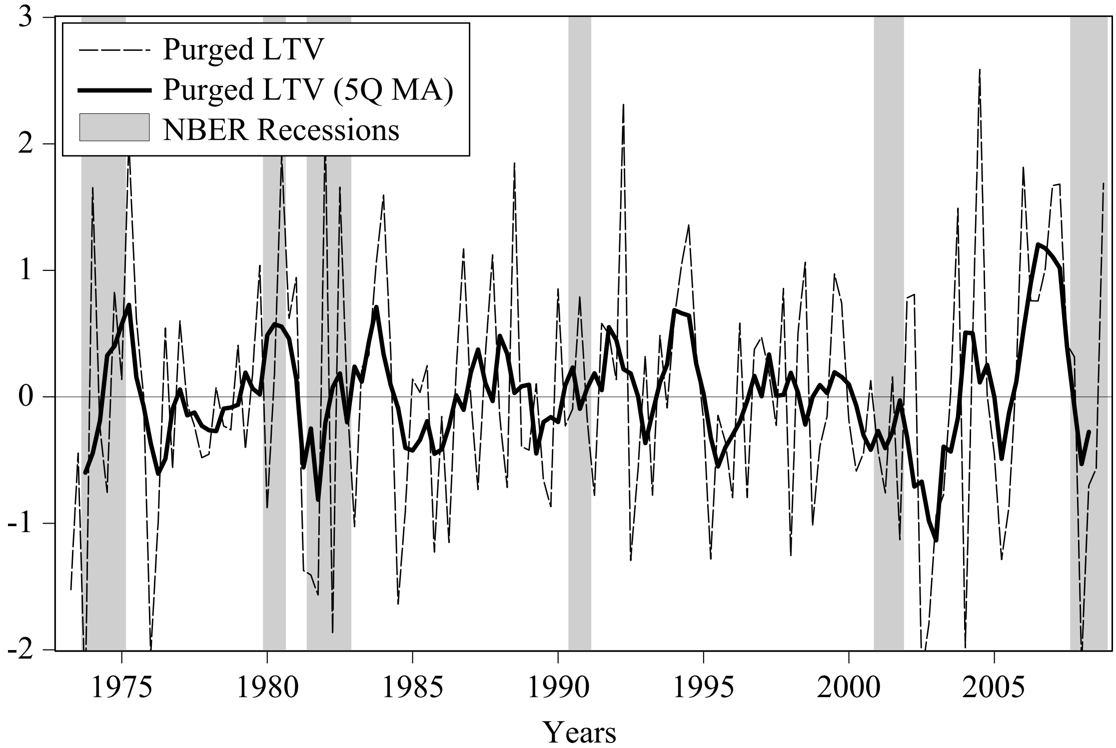
[15]We use the following data sources: from the FRED database, we obtain u_t (identifier: UNRATE) and from the Flow of Funds database, we obtain nw_t (identifier: Z1/Z1/FL702090095.Q). For $cspr_t$, ebp_t, rr_t, and $tspr_t$, we draw on the data set of Gilchrist and Zakrajsek (2012), which is provided on `https://www.aeaweb.org/articles.php?doi=10.1257/aer.102.4.1692`, while the historical Survey of Professional Forecasters data can be downloaded from `https://www.philadelphiafed.org/research-and-data/real-time-center/`.

supply unaffected by *demand factors*. The finding of a significant negative impact of the expected change in the unemployment rate on Δltv_t is consistent with Duca et al. (2011), who report a negative correlation and a significant causal impact of changes in the unemployment rate on first-time home owner LTV ratios in a related exercise.

Figures 2.8 and 2.9 plot the adjusted LTV ratio, ε_t^{ltv}: down payment constraints were particularly pronounced following the early 2000 recession, and, interestingly, lending standards eased most during the latest housing cycle. Next, we use the adjusted LTV ratio series to re-run the structural VAR from Section 2.2.3 and study the transmission of the LTV shock for the new LTV measure. Figure 2.10 traces out the impulse responses to an exogenous shock in the purged LTV series. As in Bassett et al. (2014), the core results are little affected by the removal of factors that might drive loan supply and loan demand simultaneously, i.e., the impulse responses are similar compared to the raw LTV specification. As before, the LTV shock exhibits fairly persistent effects on the LTV ratio itself and raises non-residential investment in a hump-shaped manner. Residential investment displays a small initial surge before falling significantly, and monetary policy responds to the shock with an interest rate hike of more than 10 basis points. Shutting down the policy response in the counterfactual economy, again predicts a stronger reaction of non-residential investment and changes the quantitative and qualitative dynamics of residential investment. The latter rises sluggishly in the counterfactual economy in a way similar to the non-residential counterpart. In summary, the main findings of this chapter are not affected by the purging exercise of Bassett et al. (2014). We therefore conclude that shifts in the raw LTV ratio series are a good measure of changes in banks' mortgage loan supply. Furthermore, by featuring significant spillovers to non-residential sectors, the residential mortgage LTV ratio is obviously indicative of shifts in the effective loan supply of banks in a broad sense, rather than in a narrow sense with a focus solely on real estate.[16]

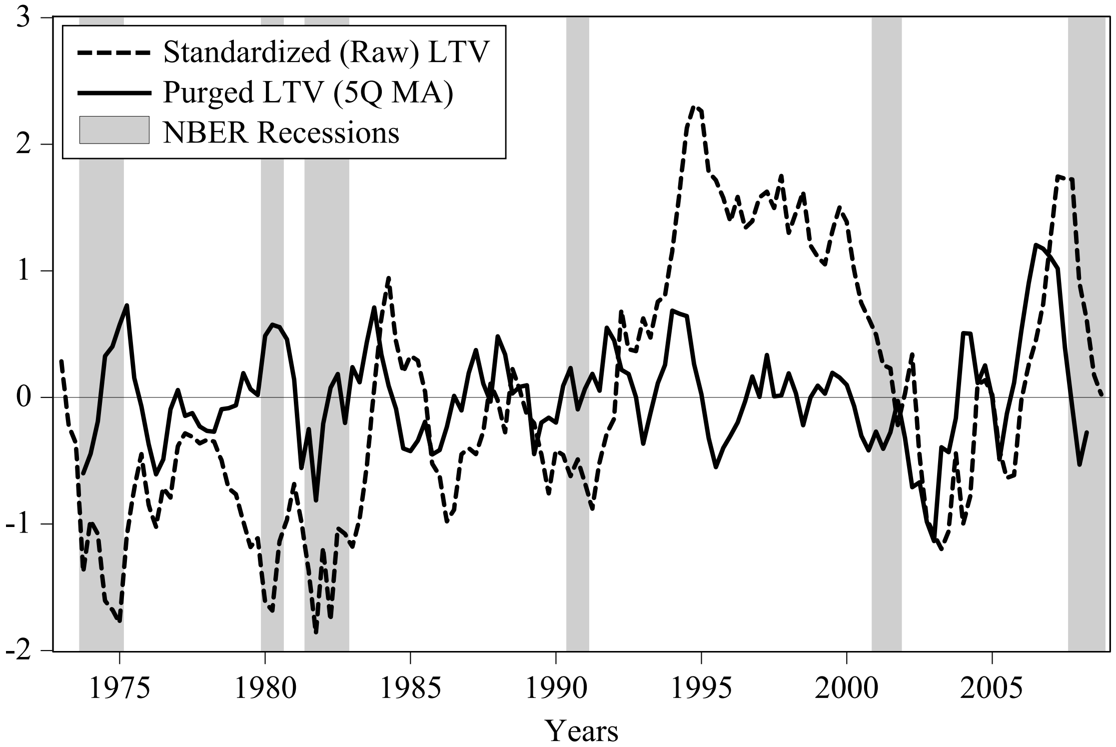
[16]This view is supported by Figure 2.16 (appendix), where we show that the results are essentially unaltered when we substitute an auto loan LTV ratio series for the residential mortgage LTV series. A priori, it is not clear that a LTV ratio for auto loans should have any impact on GDP, business/residential investment or the Federal Funds rate. That it does, supports our view that both LTV ratios, for mortgages and auto loans, proxy for bank lending conditions more broadly.

Figure 2.8: Loan-to-value ratio purged from putative demand factors I



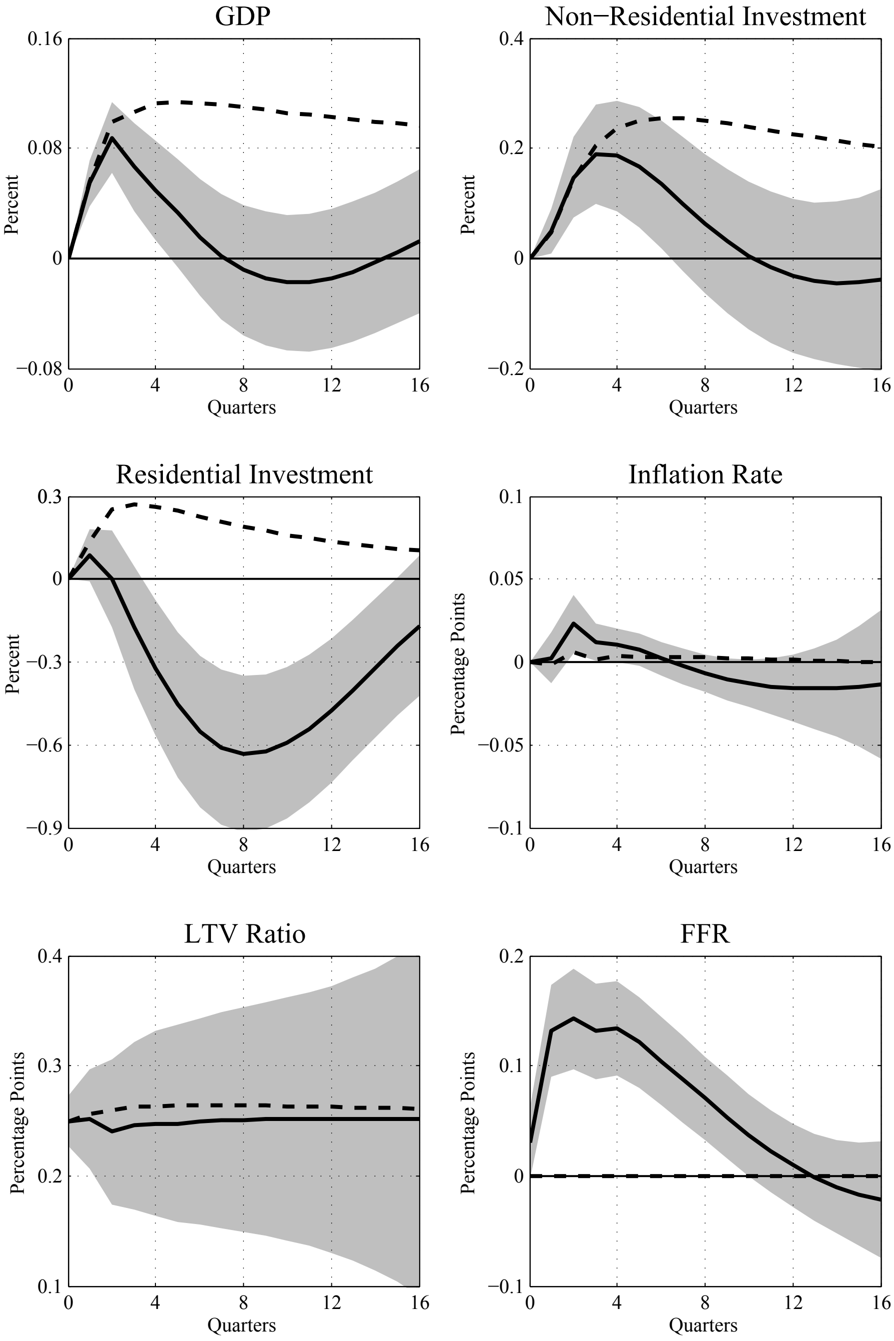
Notes: The figure displays the residuals of Equation (2.7), standardized by the standard error of estimation, together with a central five-quarter moving average of these standardized residuals. The shaded areas represent NBER-dated recession episodes in the US.

Figure 2.9: Loan-to-value ratio purged from putative demand factors II



Notes: The figure displays the five-quarter moving average series of the standardized residuals together with demeaned and standardized-by-its-standard-deviation raw LTV ratio series from Figure 2.1. The shaded areas represent NBER-dated recession episodes in the US.

Figure 2.10: Shock to LTV ratio purged from putative demand factors



Notes: The x-axis represents time in quarters. The solid lines represent point estimates of impulse response functions for the VAR, using $\boldsymbol{x}_t = [y_t\ i_t^{nr}\ i_t^{r}\ \pi_t\ \Delta ltv_t\ r_t]'$, where Δltv_t is now given by the purged LTV ratio series, that is, the residuals of Equation (2.7). Shaded areas display one standard deviations confidence intervals obtained from 5,000 replications of the recursive-design wild bootstrap procedure of Goncalves and Kilian (2004). The dashed lines represent counterfactual impulse responses for the case of a passive monetary policy authority that does not react to the shock at all, as in Bernanke et al. (1997) and Sims and Zha (2006).

2.3.2 LTV shocks and first-time home owners

In this section, we study the macroeconomic effects of an LTV shock for an alternative, more limited measure of LTV ratios. Duca et al. (2011, 2013), using data from the American Housing Survey (AHS), emphasize the role of *first-time home owners* for mortgage markets, as a major share of this marginal group of home buyers should be subject to credit and collateral constraints. Analogously to Bassett et al. (2014), Duca et al. (2011, 2013) adjust their raw first-time buyer LTV ratio series from AHS for certain cyclical factors, such as, e.g., the unemployment rate, seasonal factors, and some exceptional events.

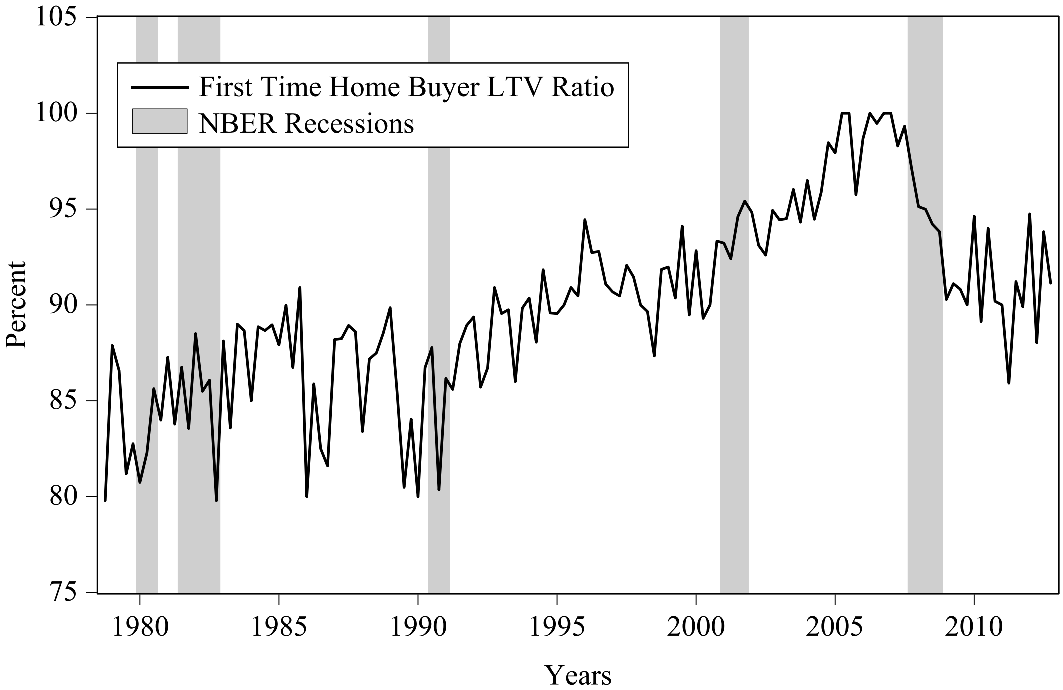
Figure 2.11 plots the adjusted median LTV ratio for first-time home buyers provided to us by Duca et al. (2011, 2013).[17] The sample starts in 1978Q4 because the AHS data is available from then onwards. The first-time home buyer LTV series is noisier than the overall LTV ratios because the number of first-time buyers in any AHS quarter is small. The series exhibits a range of variation of about 20 percentage points, which is about twice as large as for the benchmark FHFA LTV ratio. Furthermore, the average value over time of first-time home buyer LTV ratios in the sample amounts to 90 percent, whereas the counterpart for all home buyers is only slightly above 75 percent. The series fluctuates around a mean of about 85 percent in the 1980s. Then first-time home buyer LTV ratios steadily increase before declining again at the onset of the Great Recession.

In Figure 2.12, we show the effects of changes in first-time home buyer LTV ratios in our benchmark structural VAR. All the effects are less pronounced both in terms of magnitude and statistical significance, presumably reflecting the smaller number of first-time home buyers in the data. Yet, the qualitative behavior of the impulse response functions is consistent with the benchmark specification, despite the aforementioned fact that, unconditionally, the LTV ratio for first-time home owners exhibits a noticeably different time series behavior than the one for all home owners. Again, the loosening of LTV ratios triggers fairly persistent movements in

[17] We kindly thank Duca et al. (2011, 2013) for providing us with their data. They add back in the Hodrick-Prescott trend which they had removed from their data time series before the purging procedure, which is why this data can be interpreted as a fraction and is directly comparable to the LTV ratio series for all home owners displayed in Figure 2.1.

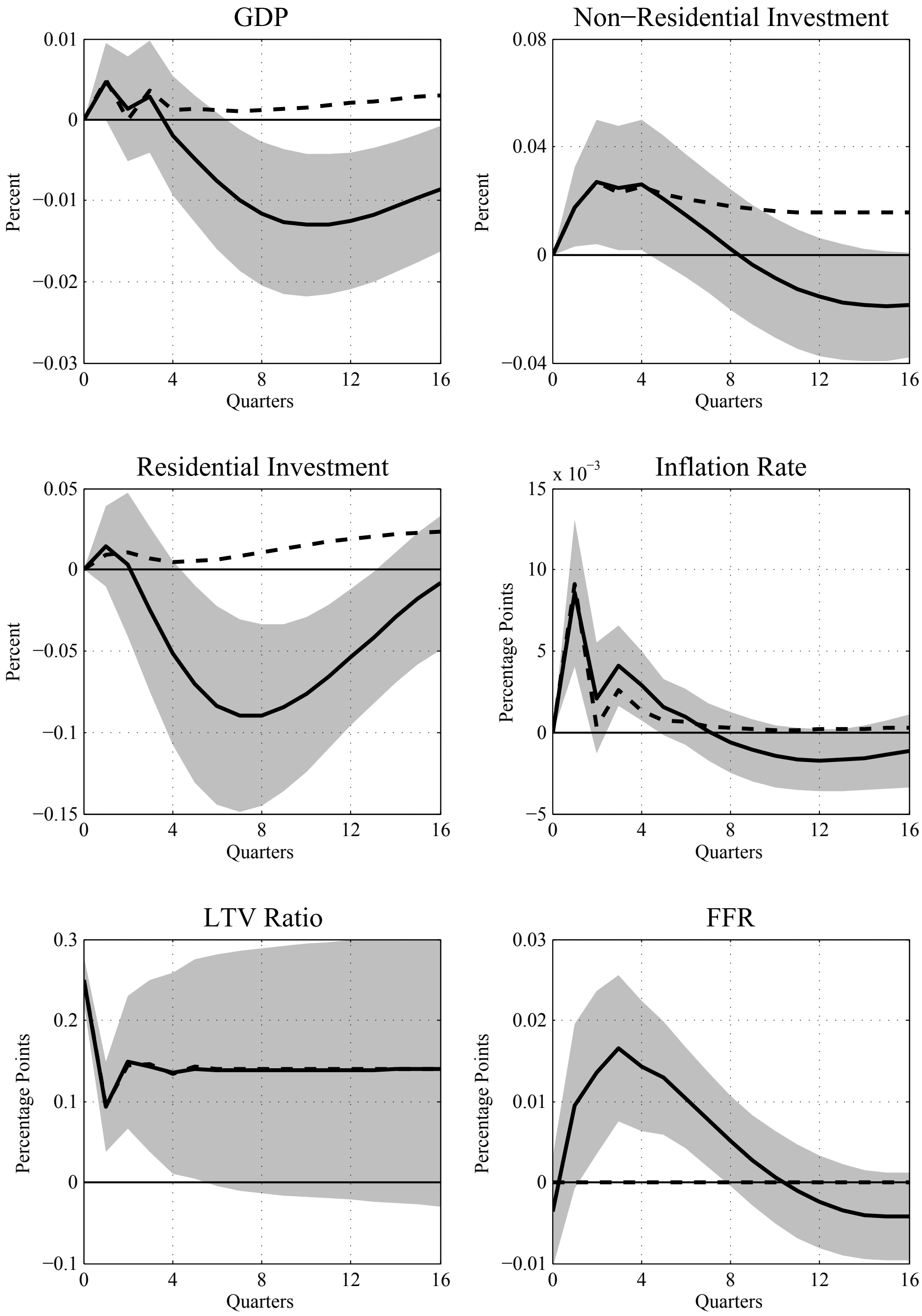
LTV ratios. Non-residential investment increases, whereas we find a contractionary impact on residential investment. The Federal Funds rate increases. Given the less significant tightening of monetary policy, the differences between the original and the counterfactual economy are also less pronounced. However, the passive monetary policy experiment still predicts more expansionary effects for non-residential investment and a surge of residential investment following the LTV shock.

Figure 2.11: Loan-to-value ratio of first-time home buyer mortgage loans



Notes: The figure displays the loan-to-value ratio for first-time home buyer mortgage loans, based on the American Housing Survey (AHS) and purged for certain cyclical factors, such as, e.g., the unemployment rate, seasonal factors, and some exceptional events (see for details Duca et al., 2011, 2013). This series has been provided to us by Duca et al. (2011, 2013). Data are at the quarterly frequency and we express them in percent, i.e., as a ratio of the granted mortgage loan and the underlying house price multiplied by 100. The shaded areas represent NBER-dated recession episodes in the US.

Figure 2.12: Shock to the first-time buyer LTV ratio



Notes: The x-axis represents time in quarters. The solid lines represent point estimates of impulse response functions for the VAR, using $\boldsymbol{x}_t = [y_t\ i_t^{nr}\ i_t^{r}\ \pi_t\ \Delta ltv_t\ r_t]'$, where Δltv_t is now given by the first-time buyer LTV ratio series, which has been provided to us by Duca et al. (2011, 2013). Shaded areas display one standard deviations confidence intervals obtained from 5,000 replications of the recursive-design wild bootstrap procedure of Goncalves and Kilian (2004). The dashed lines represent counterfactual impulse responses for the case of a passive monetary policy authority that does not react to the shock at all, as in Bernanke et al. (1997) and Sims and Zha (2006).

2.3.3 Alternative policy counterfactual

Kilian and Lewis (2011), in an application to oil price shocks, propose an alternative, less restrictive monetary policy counterfactual to isolate the effects of the policy response, compared to the framework in Bernanke et al. (1997) and Sims and Zha (2006) that we have used in Section 2.2.3. Recall that there the counteracting monetary policy surprises completely offset the endogenous interest rate response. That counterfactual assumed that the FED does not react to the impact of the LTV loosening *at all*, i.e., the Funds rate remained constant at any horizon. Following the approach pioneered by Kilian and Lewis (2011), we now study a counterfactual economy in which we only shut down the *direct* impact of the LTV shock with counteracting monetary policy shocks, but allow the FED to respond to the *indirect* effects of the LTV shock operating through its propagation to other variables in the system.

Using the definitions of $\boldsymbol{z}_{j,h}$ and $\boldsymbol{y}_{j,h}$ from Equation (2.6), we can recursively calculate the sequence of monetary policy shocks required to remove the direct influence of the LTV shock from the FED's reaction function as follows:

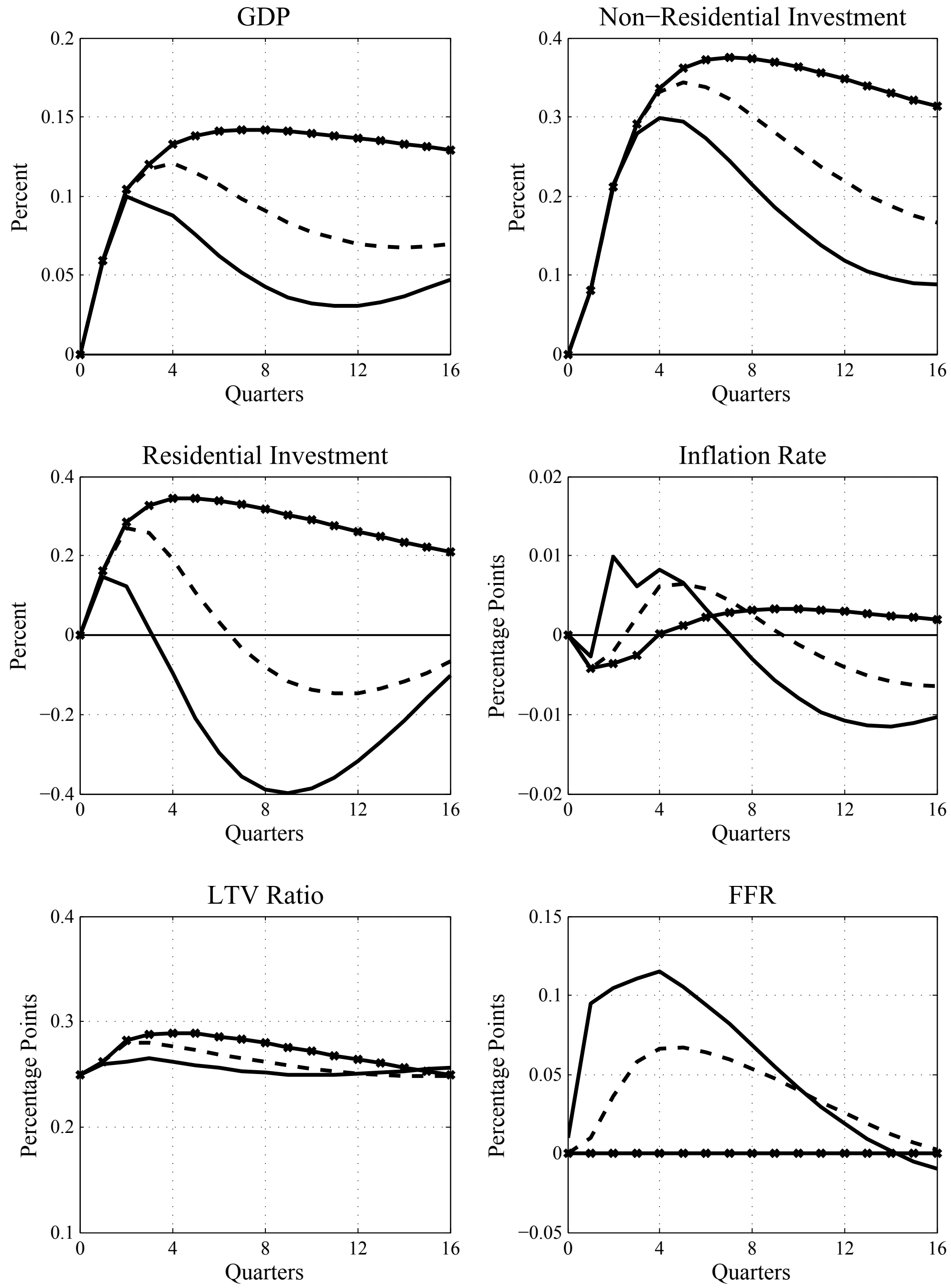
$$\boldsymbol{\varepsilon}_{F,h} = -\boldsymbol{B}_{F,L}\boldsymbol{y}_{L,h} - \sum_{m=1}^{\min(p,h)} \boldsymbol{B}_{F,mn+L}\boldsymbol{z}_{L,h-m}, \tag{2.8}$$

where the subscript L represents the position of the LTV ratio in the structural VAR.

Figure 2.13 traces out the impulse responses in the conventional monetary policy VAR used in Section 2.2.3 for the original (solid lines), the Kilian and Lewis (2011) counterfactual (dashed lines), and the counterfactual economy in which the Federal Funds rate remains constant (lines with nodes). The experiment of removing only the direct effect of the LTV shock from the FED's reaction function (dashed lines) still predicts a surge in the policy instrument. Yet, the response is more sluggish and less pronounced compared to the original economy. The direct reaction to the LTV shock accounts for roughly one third of the observed policy tightening after an expansionary LTV shock. Due to the still contractionary – but less pronounced – interest rate environment in this counterfactual experiment, the deviations from the original economy are consequently also less pronounced for the remaining variables compared to the Bernanke et al. (1997) and Sims and Zha (2006) economy.

In fact, almost all responses in the Kilian and Lewis (2011) counterfactual lie in between the original economy and the zero interest rate reaction counterfactual. Regarding the impact on residential investment, we find an increasing impulse response in the Kilian and Lewis (2011) counterfactual for one and a half years, which subsequently abates more like in the original economy. The initial surge in residential investment peaks at almost 0.3 percent, however, which is not too far away from the close to 0.4 percent peak in the Bernanke et al. (1997) and Sims and Zha (2006) economy.

Figure 2.13: LTV shock and different policy counterfactuals



Notes: The x-axis represents time in quarters. The solid lines represent point estimates of impulse response functions for the unrestricted VAR, using $\boldsymbol{x}_t = [y_t \ i_t^{nr} \ i_t^{r} \ \pi_t \ \Delta ltv_t \ r_t]'$. The dashed line displays the adjustment patterns following an LTV shock for the Kilian and Lewis (2011) policy counterfactual. The lines with nodes represent counterfactual impulse responses for the case of a passive monetary policy authority that does not react to the shock at all, as in Bernanke et al. (1997) and Sims and Zha (2006).

2.3.4 Robustness

Finally, we assess whether our main findings remain valid in a battery of additional robustness checks concerning (i) the VAR specification, (ii) the data sample, and (iii) the identification assumptions.

First, we re-run the benchmark VAR by allowing for higher lag orders of $p = 3, 4$, and 6 quarters. We also re-estimate the VAR with LTV entering in levels, and with all variables – except for the Federal Funds rate – entering in differences. Moreover, we estimate the benchmark model after dividing i_t^{nr} and i_t^{r} by the civilian non-institutional population (FRED identifier: CNP16OV). This transformation alleviates somewhat potential concerns about unit roots: the p-value of an augmented Dickey-Fuller test, with the null hypothesis of a unit root, decreases through the normalization from 0.31 to 0.15 for residential investment and from 0.06 to 0.05 for non-residential investment.[18]

Second, we check the robustness of our results with respect to the sample choice. Motivated by relative low US inflation rates and modest output fluctuations since the 1980s, Clarida et al. (2000), among others, document a significant shift in the conduct of monetary policy for post 1979 data. Beginning with the appointment of Paul Volcker as the FED's chairman, their estimated monetary policy reaction function changes considerably toward a more proactive attitude of controlling the inflation rate. Following Clarida et al. (2000), we therefore re-estimate the VAR by excluding the pre-Volcker era and starting the sample in 1979Q3. In addition, banks' lending standards eased considerably in the buildup phase to the most recent US housing cycle, which suggests large bank lending shocks during this episode (see Figures 2.1, 2.8, and 2.9). To study whether our results are driven by this perhaps extraordinary period, we exclude it from the sample and re-estimate the VAR in yet another specification with data ending in 1999Q4.

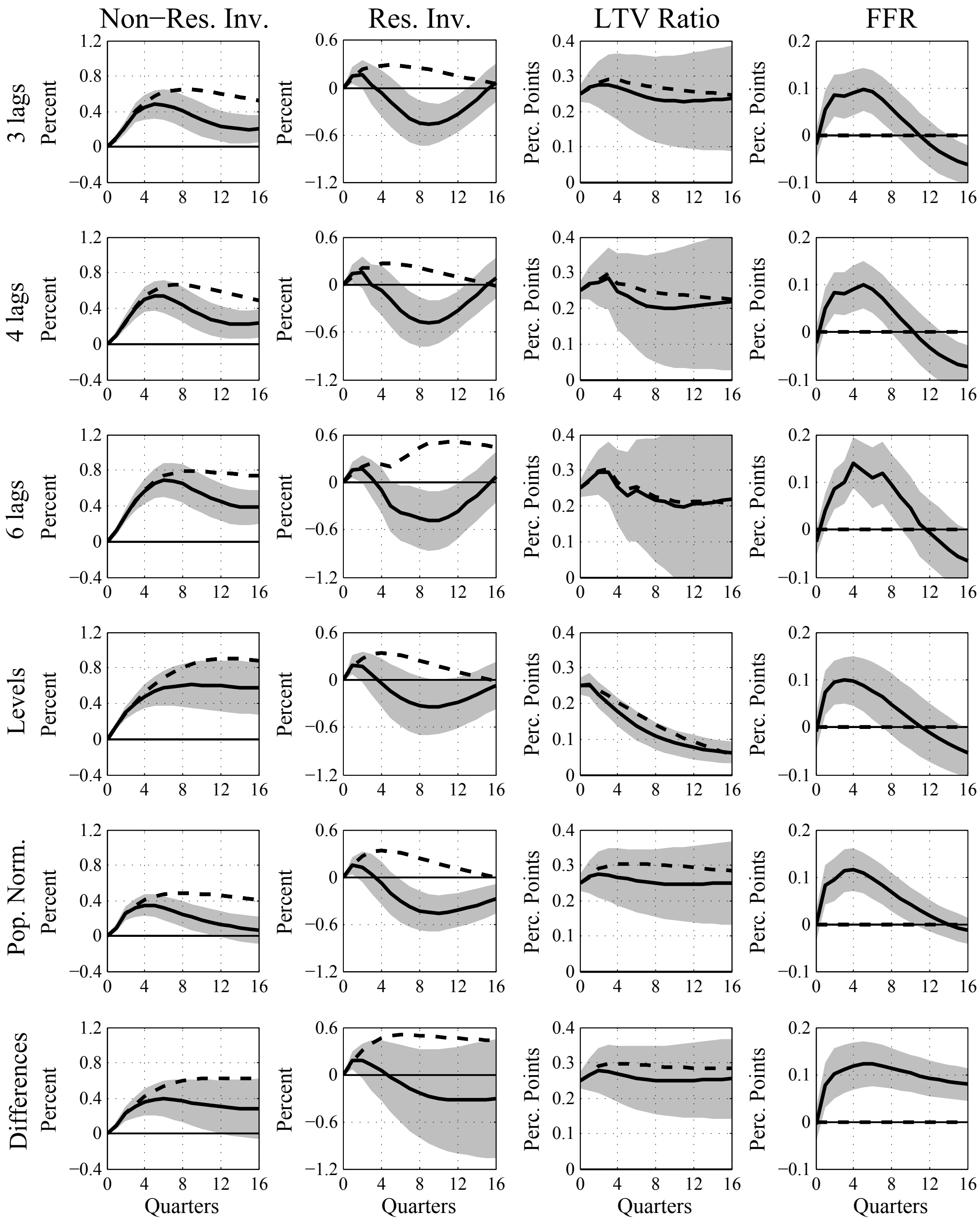
Third, we analyze the sensitivity of our results to the ordering of the variables

[18]In addition, we estimate the richer VAR specifications from Sections 2.2.3 and 2.2.4 with population-normalized data. After the population transformation, the null hypothesis of a unit root can be rejected for all debt measures based on conventional significance levels. For real GDP, the corresponding p-value declines from 0.40 to 0.22 after the normalization. All of these results are almost identical compared to the non-standardized versions and available from the authors upon request.

in the recursive identification scheme. Thus far, we have assumed that LTV shocks affect monetary policy on impact, yet, exogenous shifts in monetary policy propagate to lending standards with a time lag of one quarter. We implemented this notion by ordering Δltv_t before r_t within the block of financial variables in the benchmark model. Now, we assume that LTV shocks propagate to all other variables with a delay of one quarter, but monetary policy surprises are allowed to influence lending standards in the impact quarter.

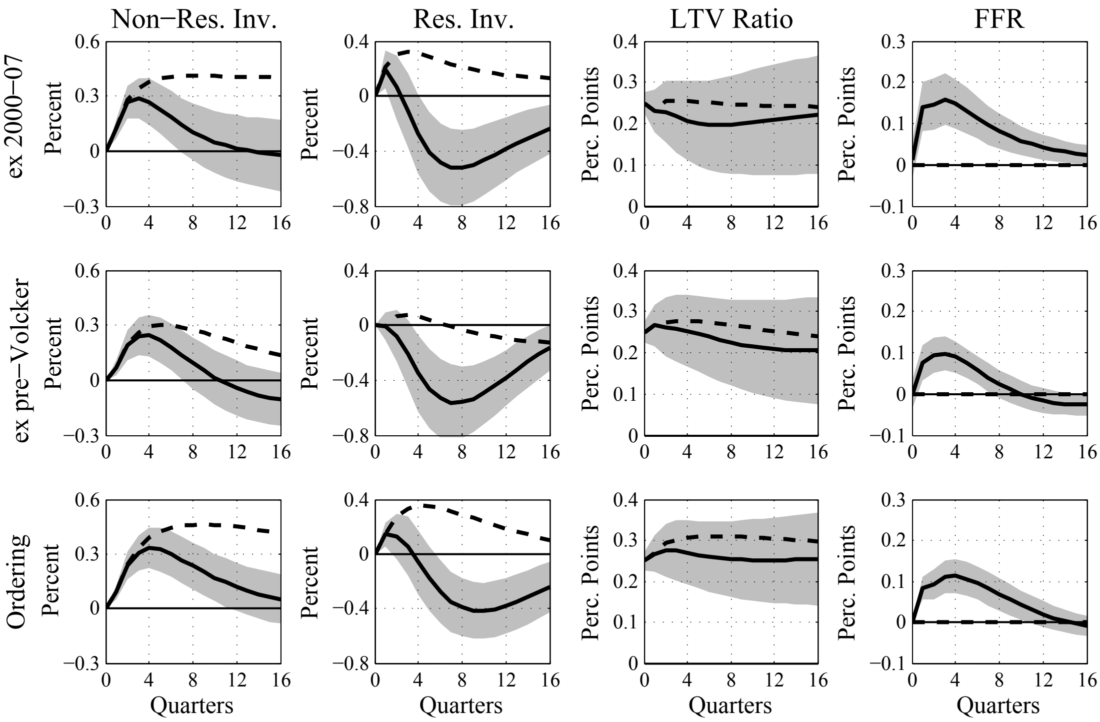
In Figures 2.14 and 2.15 we summarize the results of these robustness exercises, where we display the variables of interest in columns and the different specifications in rows. While the magnitudes differ somewhat across specifications, the qualitative patterns of the impulse responses both in the original and the counterfactual fixed interest rate economy are unaffected by these sensitivity analyses.

Figure 2.14: Robustness: VAR specification



Notes: The x-axis represents time in quarters. The solid lines represent point estimates of impulse response functions for the VARs, which use $\boldsymbol{x}_t = [i_t^{nr}\ i_t^{r}\ \Delta ltv_t\ r_t]'$ as their point of departure. Shaded areas display one standard deviations confidence intervals obtained from 5,000 replications of the recursive-design wild bootstrap procedure of Goncalves and Kilian (2004). The dashed lines represent counterfactual impulse responses for the case of a passive monetary policy authority that does not react to the shock at all, as in Bernanke et al. (1997) and Sims and Zha (2006).

Figure 2.15: Robustness: sample and Cholesky ordering



Notes: The x-axis represents time in quarters. The solid lines represent point estimates of impulse response functions for the VARs, which use $\boldsymbol{x}_t = [i_t^{nr} \; i_t^{r} \; \Delta ltv_t \; r_t]'$ as their point of departure. Shaded areas display one standard deviations confidence intervals obtained from 5,000 replications of the recursive-design wild bootstrap procedure of Goncalves and Kilian (2004). The dashed lines represent counterfactual impulse responses for the case of a passive monetary policy authority that does not react to the shock at all as in Bernanke et al. (1997) and Sims and Zha (2006).

2.4 Conclusion

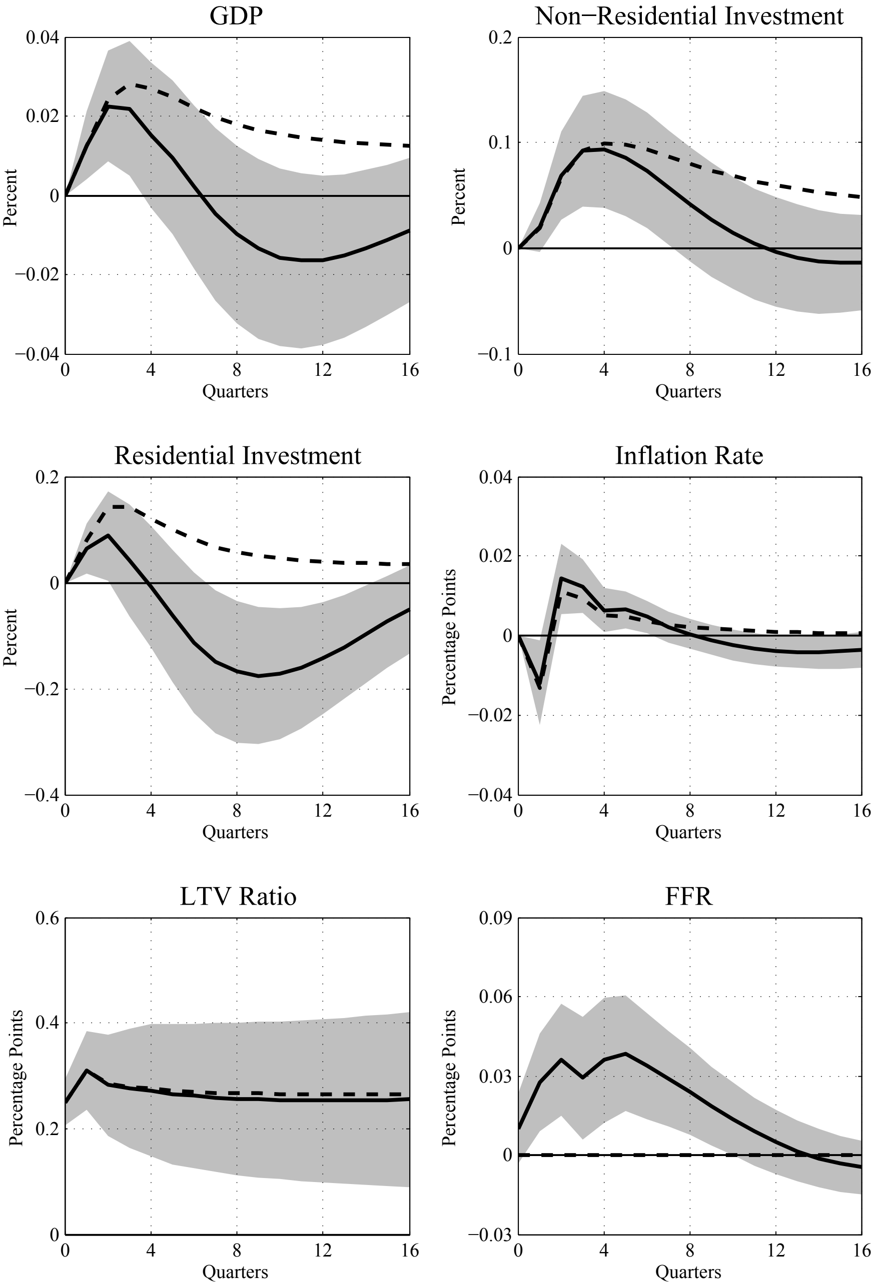
This chapter studies the macroeconomic consequences of shifts in banks' lending standards as measured by residential mortgage LTV ratios. Using LTV data from the Federal Housing Finance Agency, we find that exogenous LTV shocks feature significant spillovers to non-residential sectors, giving LTV ratios the interpretation as a general loan supply indicator. Perhaps surprisingly, however, we also find that shifts in LTV ratios are not likely to be a substantial driver of residential investment and household debt. The reason behind this result is a systematic monetary policy response, which tightens as a reaction to looser LTV ratios. As a result, residential investment and household debt decline.

Our findings suggest an important policy lesson: the use of supervisory limits on LTV ratios as a macroprudential policy tool to curb overheating housing markets should at the very least be calibrated to take into account interactions with monetary policy, ideally it should be coordinated with monetary policy.

Lastly, how can we interpret the events during the Great Recession through the lens of our results? While our VAR is a linear tool and we exclude quite deliberately the Great Recession quarters from our analysis, and while we find no significant role for changes in LTV ratios as a driver of residential investment and household debt – including the housing cycle from 2001 to 2007 – our results are in line with the perception that a tightening of LTV ratios may have exacerbated the downturn in housing markets at the onset of the Great Recession. The reason is the asymmetry represented by the zero lower bound on nominal interest rates. Historically, the FED would have lowered interest rates in the face of the LTV tightening; however, with interest rates bounded at zero, this cushioning mechanism had to be absent. According to our fixed interest rate results, such a situation should then be associated with a drop in residential investment, which was indeed observed during the financial crisis.

2.5 Appendix to chapter 2

Figure 2.16: LTV shock for new car loans at auto finance companies



Notes: The x-axis represents time in quarters. The solid lines represent point estimates of impulse response functions for the VAR, using $\boldsymbol{x}_t = [y_t \ i_t^{nr} \ i_t^{r} \ \pi_t \ \Delta ltv_t \ r_t]'$, where Δltv_t is now given by the LTV ratio series for new car loans at auto finance companies from FRED, with series identifier DTCTLVNLNM. We apply the Census X-12 filter to seasonally adjust this monthly LTV series and then use the quarterly average in the VAR. The sample period is 1973Q1-2008Q4, as in the baseline VAR. Shaded areas display one standard deviations confidence intervals obtained from 5,000 replications of the recursive-design wild bootstrap procedure of Goncalves and Kilian (2004). The dashed lines represent counterfactual impulse responses for the case of a passive monetary policy authority that does not react to the shock at all, as in Bernanke et al. (1997) and Sims and Zha (2006).

Chapter 3

Current Account Dynamics and the Housing Cycle in Spain

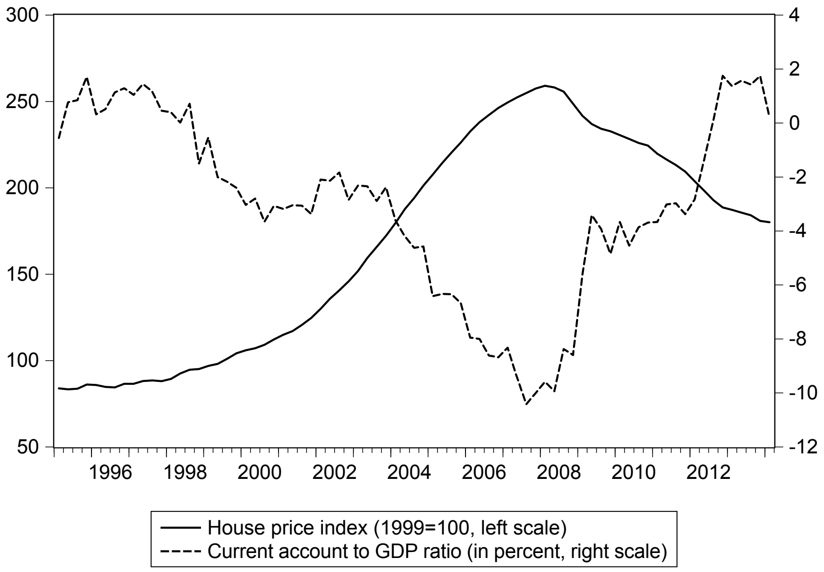
3.1 Introduction[1]

What are common drivers of the well-established, negative correlation between housing markets and the current account in Spain? Spain witnessed a pronounced boom and bust cycle in housing[2], which coincided with a deterioration and subsequent contraction of its current account (see Figure 3.1). From 1995 to 2008 real square meter property prices tripled on average, and during the culmination of the boom one fourth of the Spanish male labor force was employed in the construction sector that temporarily accounted for 20 percent of GDP growth. At the peak of the boom, the current account to GDP ratio recorded minus 10 percent, followed by a sharp correction after the bust.

[1]This chapter is based on joint work with Daniel Maas and Eric Mayer. An earlier version appeared as Maas et al. (2015).

[2]Fernández-Villaverde et al. (2013), Gonzalez and Ortega (2013), and Akin et al. (2014) provide an overview of the Spanish cycle in housing markets. In general, housing is of particular importance in Spain as the rate of home ownership and the share of private wealth allocated to housing both exceed 80 percent, which is considerably beyond European average.

Figure 3.1: Current account and house price dynamics



Notes: The figure presents the current account to GDP ratio and house prices for Spain. We obtain the data from Eurostat and BIS.

This chapter tests four popular hypotheses along their ability to generate the joint behavior of housing markets and the current account that is present in Spanish data. In this regard, we account for Spain-specific as well as for external shocks emerging in the rest of the Eurozone. The comparison of such "pull" (domestic) and "push" (foreign) factors, at least, dates back to Calvo et al. (1993) and is still subject to research on the sources of capital flows (Fratzscher, 2012).

The pull hypothesis emphasizes the importance of domestic factors as potential drivers of the housing boom in Spain. By initiating a domestic boom these factors, ultimately, attract capital inflows from the rest of the Eurozone. Prime candidates for this hypothesis are a relaxation of credit standards that foster credit supply by the banking industry (see, e.g., Helbling et al., 2011; Bassett et al., 2014) as well as housing bubble shocks that fuel markets against the backdrop in belief of ever surging house prices (see, e.g., Shiller, 2005, 2007; In't Veld et al., 2011).

In contrast, the push hypothesis explains housing markets by external factors that proactively allocate capital to Spain. One representative is the risk premium shock (see in't Veld et al., 2014). The creation of the common Euro denominated market eliminated risk premia among the member countries, which led core Eurozone investors to invest in Spain and further lowered risk free rates. Vice versa, the economic turmoil in 2008 reintroduced risk spreads and reverted capital flows. A further push representative is a European version of the "savings glut" shock originally proposed by Bernanke (2005) for the US. The rationale of this shock is that Spain as member of a monetary union was overheated by too low interest rates compared to a Taylor rate. As a consequence, and in line with consumption dynamics, core Europe had systematically higher saving rates than Spain and lower economic momentum during the run-up phase. Consequently, excess savings from the core broke its way through to Spanish housing markets.

We empirically analyze how the competing shocks impact the current account and housing market variables. We study how the shocks propagate through the economy and, furthermore, we judge their quantitative relevance. We do so by applying a robust sign restrictions approach as in Peersman and Straub (2009) to Spanish and rest of Euro Area data. We derive restrictions from a single currency union DSGE model incorporating two countries, i.e., Spain and the rest of the Euro Area. The model builds on Rabanal (2009) and Iacoviello and Neri (2010) and features a variety of nominal and real frictions. Following Kiyotaki and Moore (1997), households consist of two subgroups according to their time preferences, i.e., savers and borrowers (see Monacelli, 2009). As in Iacoviello (2005), borrowers face a collateral constraint such that their borrowing is limited to the present value of their housing multiplied by a loan-to-value (LTV) ratio. In the empirical analysis, we employ an open-economy vector autoregressive (VAR) model, which allows a discrimination of push and pull forces. Due to the short sample size, we follow Eraker et al. (2014) and draw on a Bayesian mixed frequency approach for estimation and inference. The identification of structural shocks is along the lines of Uhlig (2005). Concretely, we identify a savings glut, a risk premium, a financial easing, and a housing bubble shock. Except for the financial easing shock, all identified disturbances are capable of generating the observed, negative correlation of the current account and housing markets. In contrast to the competing macroeconomic disturbances, the financial easing shock predicts no robust, significant drop in the current account and, most notably, a decline in residential investment and house prices. Moreover, the savings glut shock has most explanatory power for real house prices, while the risk premium shock, in particular, has explanatory power for residential investment. Overall, the housing bubble shock accounts for a slightly smaller share of variation in the data, while the financial easing shock explains the key variables to a similar extend as both push disturbances.

Our contribution to the current literature is along the following dimensions. First, for the US there is a number of theoretical and empirical studies analyzing the joint dynamics of the current account and housing markets (see, e.g., Sá and Wieladek, 2015; Justiniano et al., 2014). However, prima facie, it is not evident, which conclusions drawn from US data can be applied to Spain.[3] Most importantly, Spain is member of a currency union and net capital inflows did not come from Asia and oil exporting countries, but largely from the rest of the Euro Area. Thus the study of Spain, in particular, helps to understand the specifics of the nexus between housing markets and the current account inside a monetary union, where shocks propagate differently due to the common conduct of monetary policy. Despite different currency regimes, we reinforce the results of Sá and Wieladek (2015) for the US by also revealing the importance of savings glut shocks for Spain. Second, in't Veld et al. (2014) estimate a rich DSGE model by Bayesian techniques with Spanish data. They find a strong influence of falling risk premia, a loosening of collateral constraints, and asset price shocks on the Spanish output boom and capital inflows. We complement their analysis with a time series approach, which imposes less structure on the data. Furthermore, we focus on the housing boom rather than the Spanish output cycle. We find little support for financial easing shocks in explaining the negative correlation of housing markets and the current account, which is in line with in't Veld et al. (2014). Third, due to limited data availability, contributions like Hristov et al. (2012) or Ciccarelli et al. (2015) rely on panel data approaches to achieve efficiency gains. Likewise, single country VAR approaches often resort to data samples that extend the relevant time period for the same

[3]For instance, Spain has a bank-based financial system operating under the tight Basel regulatory framework, where new constructions were only moderately fueled by sub-prime residential mortgage-backed securities. In contrast, the US is known to be a predominantly market-based financial system, where sub-prime markets were loosely regulated, which was center stage at the crisis (see, e.g., Goddard et al., 2007).

reason. To tackle this issue, we simultaneously employ monthly and quarterly data for Spain in the Bayesian mixed frequency framework as in Eraker et al. (2014).

The chapter is structured as follows. In Section 3.2, we explain the different hypotheses that we empirically test in detail. Section 3.3 discusses the model employed to derive the sign restrictions. Section 3.4 describes the econometric framework and presents the results. Section 3.5 provides some extensions and robustness, while Section 3.6 concludes.

3.2 Four hypotheses

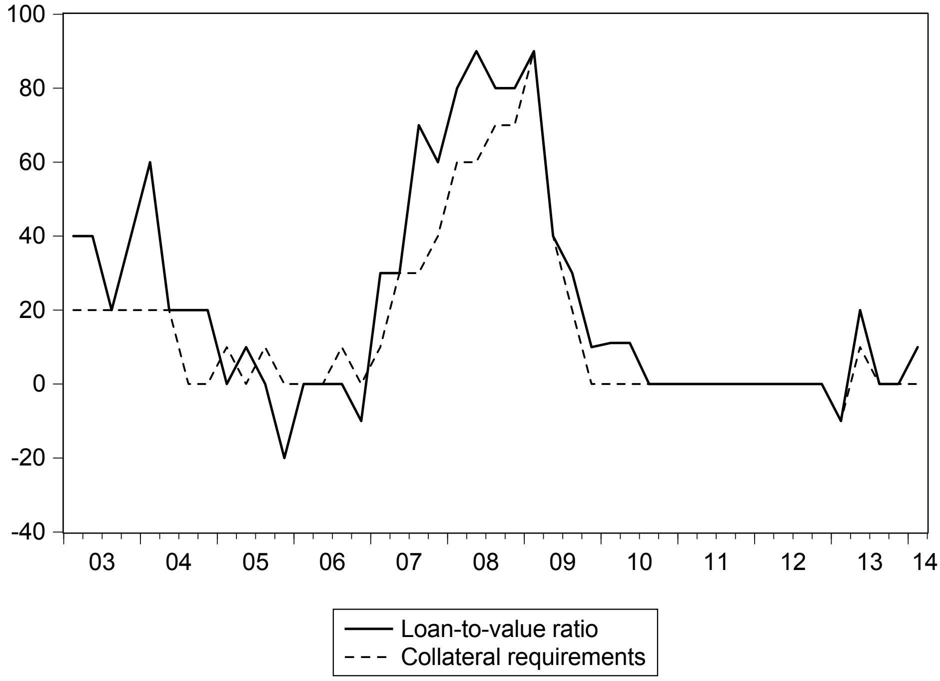
To motivate the analysis, we further discuss four different sources that potentially link the housing and current account[4] cycles in Spain.

We begin the exposition with pull factors of capital flows. In Spain's bank-based financial system the majority of mortgages was supplied by the banking industry. Formally, under the Basel regulatory framework, banks faced stricter equity requirements, once LTV ratios exceeded 80 percent of the collateral value. In practice, banks placed 40 percent of all mortgage loans exactly on the limit of 80 percent. Furthermore, appraisal firms systematically overstated property values (Akin et al., 2014), thereby effectively raising LTV ratios in terms of market values and softening lending standards before the crisis (see Figure 3.2). As the fraction of collateral constrained households is sizeable in Spain (Hristov et al., 2014), the effective loosening of collateral requirements is of first order macroeconomic importance. Beyond, and induced by, inter alia, tough competition in the banking sector, Spanish mortgage rates were 21 percent below European average. The expansion in the effective loan supply of Spanish banks, of course, could also have been driven by changes in the conduct of local banking supervision or the regulatory environment as well as through shifts in industry strategies (e.g., Bassett et al., 2014). As mortgage growth

[4]As argued in Shin (2012) and Acharya and Schnabl (2010), gross financial flows are more crucial for overall financing conditions than net capital flows as reflected by the current account. Yet, Obstfeld (2012) emphasizes the importance of the current account for the scrutiny of policy makers (see Fratzscher et al., 2010). Catão and Milesi-Ferretti (2014) point out the current account as a predictor of external crises. Furthermore, Giavazzi and Spaventa (2011) stress the relevance of the current account, in particular, for the case of a monetary union.

was not backed by domestic wholesale funding, it triggered capital inflows, predominantly, from core Eurozone countries. In summary, we refer to these developments as financial easing shocks.

Figure 3.2: Changes in Spanish banks' lending standards



Notes: The figure shows the change in banks' conditions for housing loans to households over the past three months (frequency of tightened minus eased lending standards). We obtain the data from the ECB's bank lending survey, which is available since 2003.

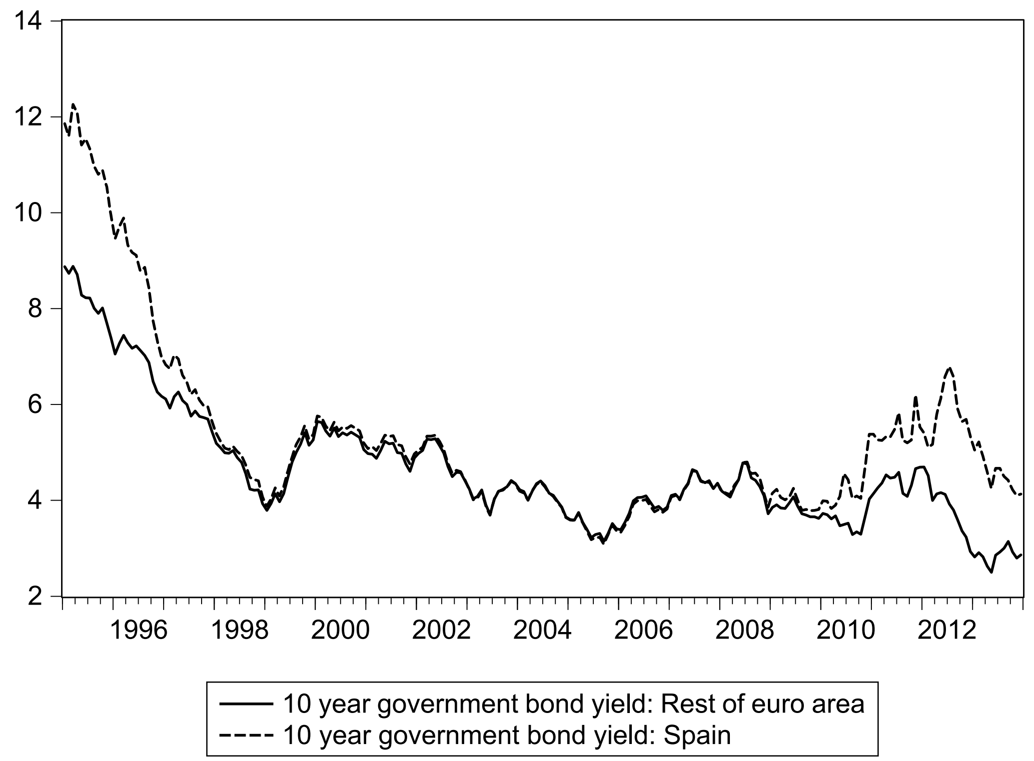
A second prominent pull hypothesis are housing bubble shocks (see, e.g., In't Veld et al., 2011). Following Shiller (2005, 2007), a housing bubble is best described by a social pandemic, which is fueled by the belief of ever increasing house prices thereby raising the willingness to pay higher prices. According to Laibson and Mollerstrom (2010), Adam et al. (2012), and in't Veld et al. (2014), housing bubble shocks, moreover, cause current account deficits and thus capital inflows.[5] Empirically, asset prices are a main driver of the US current account, which is in line with the housing bubble hypothesis (Fratzscher et al., 2010). The rationale of the housing bubble shock is that housing demand is stimulated by the belief of rising house prices. As housing serves as collateral, higher house prices also lead to stronger demand for

[5]Beyond, Cheng et al. (2014) and Ling et al. (2014) stress the importance of housing bubble shocks for a housing boom.

non-durable goods. Accordingly, the domestic demand expansion induces imports causing current account deficits. Besides, housing bubble shocks can explain the coincidence of increasing house prices and strong residential investment, whereas financial easing shocks need not necessarily account for this feature (Justiniano et al., 2014). The dynamics of residential investment are an important facet of the Spanish housing boom, as the ratio of residential investment to GDP almost doubled from 1995 to 2006. As increasing house prices loosen collateral constraints, the overall transmission of housing bubble shocks to the broader economy, however, is similar to financial easing shocks.

Now, we discuss the competing push hypothesis. The push view, for instance, underlies the so-called risk premium shock (in't Veld et al., 2014). Beginning with the Madrid Summit in 1995, Spanish risk free rates started to converge to the level of German bond rates (see Figure 3.3).

Figure 3.3: 10-year government bond yields



Notes: The figure depicts the development of 10-year government bond yields for Spain and the rest of the Euro Area. We obtain the data from Eurostat.

According to the risk premium narrative, the introduction of the common European currency, as a whole, created an institutional environment that encouraged portfolio investors and banks to expand portfolio investment and lending to the periphery as, e.g., Spanish assets were paying higher yields. First and foremost, the creation of the Euro eliminated currency risks and might even have made investors belief in possible bail outs, decreasing the perception of political risks. Besides, as pointed out in Hale and Obstfeld (2014), the ECB's refinancing policy did not discriminate between Spanish and, e.g., German sovereign bonds, despite their different credit ratings. The same applies to capital requirements that attached zero risk weights to all Euro Area government debt obligations. The introduction of an efficient payment settlement system (TARGET), in addition, eliminated transaction cost. With the financial crisis hitting in 2008, risk spreads re-emerged, the current account reverted, and housing markets collapsed.

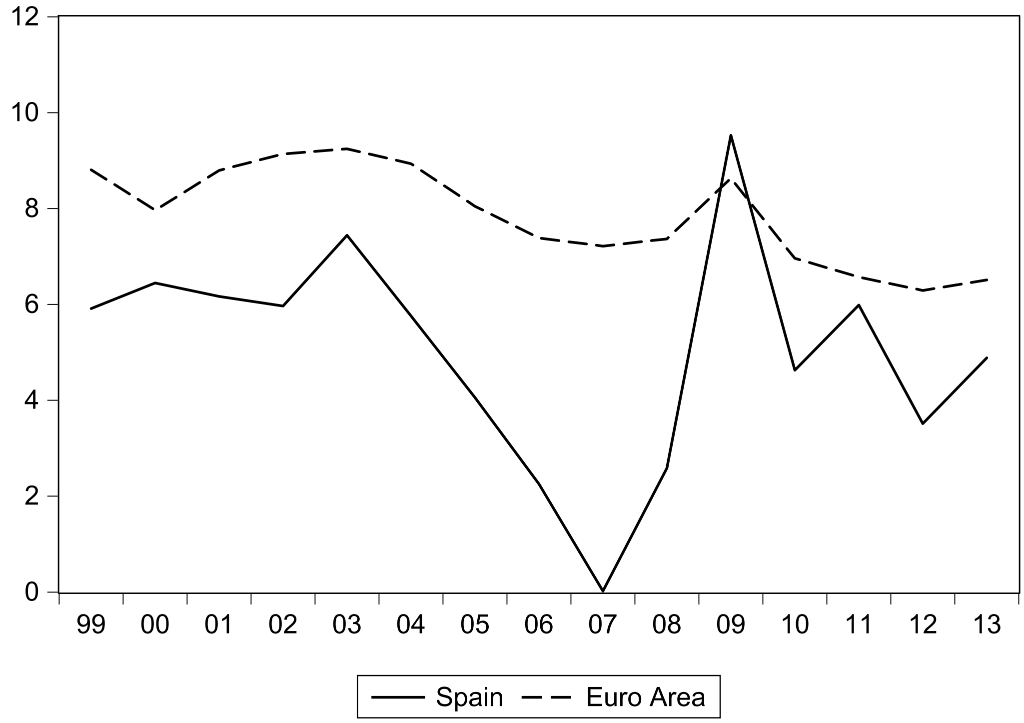
Another push factor conveys a European variant of the "savings glut" (Bernanke, 2005; Mendoza et al., 2009) shock operative for Spain. Clearly, the savings glut hypothesis cannot be literally applied to Spain. The idea of "uphill" flowing money, in particular, from China to the US, due to an underdeveloped Chinese financial system with a limited amount of financial instruments, is US specific. Instead, we argue for the case of Spain as follows. In the course of the housing boom, Spanish GDP and HCPI growth rates were roughly one percentage point higher than in the rest of the Euro Area. Thus monetary conditions, measured against a Taylor rate, were excessively expansionary for Spain and provide another rationale for the current account deficits as low real interest rates, on the one hand, discouraged saving and, on the other hand, fostered investment in housing.[6]

Figure 3.4 depicts net saving rates for Spain and the Euro Area from 1999 to 2013. Since 2003, Spanish net saving rates dropped from 7 to 0 percent, before sharply reverting at the onset of the Great Recession, while the Euro Area counterpart series fluctuated modestly around 8 percent. This setting is reminiscent of a savings glut idea as savings from the core Eurozone were seeking profitable

[6]See also Adam et al. (2012) for the interaction of real interest rate dynamics and beliefs in fueling house price booms.

investment opportunities in the periphery. Slack in core economies depressed Spanish exports, while the booming Spanish economy attracted imports and triggered current account deficits.

Figure 3.4: Net household saving as percentage of net disposable income



Notes: The figure portrays net household saving as a percentage of net disposable income for Spain and the rest of the Euro Area. We obtain the data from the OECD.

3.3 DSGE model sign restrictions

In this section, we develop a New Keynesian DSGE model building on Rabanal (2009), Iacoviello and Neri (2010), and Aspachs-Bracons and Rabanal (2011).[7] We use the predictions of the model to derive robust sign restrictions of impulse response functions, which we employ for identification in the empirical analysis.

3.3.1 Model

The model features two economies in a closed monetary union, i.e., a domestic (Spain of size n) and a foreign country (rest of Eurozone of size $1-n$). In both economies, households are composed of two types, i.e., borrowers and savers, where the latter have the higher discount factor as in Kiyotaki and Moore (1997). Firms consist of monopolistically competitive intermediate goods producers as well as perfectly competitive final goods bundlers, and are partitioned into two sectors. By employing capital and labor services, firms in the first sector produce non-durable consumption and investment goods, which are traded across countries. Firms in the second sector produce housing by employing land in addition to the input factors capital and labor, with savers owning the stocks of capital and land. Households maximize lifetime utility subject to a budget constraint, where utility concavely increases in consumption of non-durables and housing, and convexly decreases in labor. Optimizing borrowers and savers allocate resources among each other, which results in equilibrium debt. As in Iacoviello (2005), debtors borrow against housing. The expected present value of housing multiplied by a LTV ratio, as a consequence, determines borrowers' collateral constraints and thus their leverage (see also Kiyotaki and Moore, 1997). Following Smets and Wouters (2003) and Christiano et al. (2005), the model considers several real and nominal frictions.

We derive sign restrictions from the DSGE model, exclusively, for shocks that are necessary for identification in the empirical analysis and which ensure orthogonality to other macroeconomic disturbances. We restrict the presentation to the optimiza-

[7] Mayer and Gareis (2013) estimate a model similar to ours with Bayesian techniques to study the housing boom and bust cycle in Ireland.

tion problems of home country households and firms as there exists symmetry across the home country and the rest of the single currency area.

Borrowers' program

We denote the continuum of borrowing households (see Monacelli, 2009) with $b \in [0, \omega]$. b represents a borrower, the share of borrowers in the economy is $\omega < 1$, and

$$\mathbb{E}_0 \left\{ \sum_{t=0}^{\infty} \zeta_{\beta,t} \tilde{\beta}^t \left(\alpha \log(\tilde{C}_t(b) - h\tilde{C}_{t-1}) + (1-\alpha) \log(\tilde{D}_t(b)) - \frac{\tilde{L}_t(b)^{1+\eta}}{1+\eta} \right) \right\} \tag{3.1}$$

is the intertemporal utility function. $\tilde{\beta}$ is the discount factor of borrowers (indicated with $\tilde{}$), where borrowers are less patient than savers, i.e., $\tilde{\beta} < \beta$. $\zeta_{\beta,t}$ is an exogenous shock disturbing the discount factor and logarithmically follows $\log(\zeta_{\beta,t}) = \rho_\beta \log(\zeta_{\beta,t-1}) + \epsilon_{\beta,t}$, with $\epsilon_{\beta,t} \sim \mathcal{N}(0, \sigma_\beta)$ and $\rho_\beta > 0$. $\mathbb{E}_t$ represents expectations formation at time t. Consumption of dwellings, $\tilde{D}_t(b)$, i.e., the stock of housing, increases borrowers' utility, whereas an index of labor supply, $\tilde{L}_t(b)$, negatively affects utility. η stands for the inverse Frisch elasticity. Consumption of a composite index comprising domestic and foreign non-durables, $\tilde{C}_t(b)$, is subject to external habits, with h determining the degree of habit formation.

The basket of non-durables is $\tilde{C}_t(b) = (\tau^{\frac{1}{\iota}} \tilde{C}_{H,t}(b)^{\frac{\iota-1}{\iota}} + (1-\tau)^{\frac{1}{\iota}} \tilde{C}_{F,t}(b)^{\frac{\iota-1}{\iota}})^{\frac{\iota}{\iota-1}}$, where subscripts indicate whether the non-durable is produced in the home, H, or foreign country, F. ι is the substitution elasticity between both non-durable goods, and τ defines the fraction of goods produced in the home country. Reallocating labor services from the non-durable consumption goods sector, $\tilde{L}_{C,t}(b)$, to the housing sector, $\tilde{L}_{D,t}(b)$, is subject to frictions as in Iacoviello and Neri (2010) and Aspachs-Bracons and Rabanal (2011). $\iota_L \geq 0$ measures cost associated with labor reallocation, and ϱ is the size of the housing sector, where the index of labor services is $\tilde{L}_t(b) = ((1-\varrho)^{-\iota_L} \tilde{L}_{C,t}(b)^{1+\iota_L} + \varrho^{-\iota_L} \tilde{L}_{D,t}(b)^{1+\iota_L})^{\frac{1}{1+\iota_L}}$.

Borrowers are constrained by the following sequence of budget restrictions

$$P_{C,t}\tilde{C}_t(b) + P_{D,t}\tilde{X}_t(b) + R_{t-1}\tilde{S}_{t-1}(b) = \sum_{j}^{C,D} \frac{W_{j,t}}{\mathcal{M}_{j,t}} \tilde{L}_{j,t}(b) + \tilde{S}_t(b) + \Pi'_t(b). \tag{3.2}$$

$P_{j,t}, W_{j,t}$, and $\mathcal{M}_{j,t}$ denote prices, wages, and nominal wage markups in sector $j = C, D$, with C denoting the non-durable and D indicating the durable consumption goods sector. The markups result from monopolistic competition that drives a wedge between wages paid by producers and those earned by borrowing households. $\tilde{X}_t(b)$ is borrowers' investment in residential property, and $\tilde{S}_t(b)$ represents one period debt that borrowers hold against domestic savers for a gross interest rate of $R_t > 1$. Ultimately, labor unions pay dividends, $\Pi_t^{\prime}(b)$.

Indebted households borrow against the expected present value of their dwellings, which serve as collateral (see Iacoviello, 2005). The nominal collateral constraint holds in every period and reads

$$R_t \tilde{S}_t(b) \leq \zeta_{LTV,t}(1-\chi)(1-\delta)\mathbb{E}_t\left\{P_{D,t+1}\tilde{D}_t(b)\right\}, \tag{3.3}$$

where χ is the rate of down-payment, i.e., $1-\chi$ the LTV ratio, respectively. $\zeta_{LTV,t}$ represents an exogenous AR(1) shock to the loan-to-value ratio with unconditional mean of zero, which eases or tightens lending standards for borrowers. Furthermore, the housing stock depreciates with rate δ and has the accumulation equation $\tilde{D}_t(b) = (1-\delta)\tilde{D}_{t-1}(b) + \tilde{X}_t(b)$. To ensure a well-defined steady state of nominal debt (Schmitt-Grohé and Uribe, 2003), borrowers in the home country pay a risk premium on the union-wide risk free bond rate, which inversely relates to deviations of the net foreign asset position from its non-stochastic steady state as in Aspachs-Bracons and Rabanal (2011)

$$\frac{R_t}{R_t^*} = \exp\left[-\kappa\left(b_t^{\prime} - b^{\prime}\right) + \zeta_{RP,t}\right]. \tag{3.4}$$

$b_t^{\prime}$ is the net foreign asset to nominal GDP ratio and $b^{\prime}$ the respective steady state. $\kappa \geq 0$ measures how sensitive the risk premium, R_t/R_t^*, reacts to fluctuations in $b_t^{\prime}$, where the union-wide (indicated with *) risk free bond rate is R_t^*. $\zeta_{RP,t}$ is an exogenous disturbance that stochastically manipulates the risk premium, with $\zeta_{RP,t} = \rho_{RP}\zeta_{RP,t-1} + \epsilon_{RP,t}$ and $\epsilon_{RP,t} \sim \mathcal{N}(0, \sigma_{RP})$.

Borrowers optimally choose non-durable consumption as well as debt holdings such as to maximize (3.1) subject to (3.2), which gives

$$\tilde{U}_{C,t} = P_{C,t}\tilde{\lambda}_t \quad \text{and} \quad R_t^{-1} = \tilde{\beta}\mathbb{E}_t\left\{\frac{P_{C,t}}{P_{C,t+1}}\frac{\tilde{U}_{C,t+1}}{\tilde{U}_{C,t}}\right\} + \tilde{\psi}_t. \tag{3.5}$$

$\tilde{U}_{C,t}$ denotes the marginal increase in utility associated with consumption of one extra unit of the non-durable good. $\tilde{\lambda}_t$ and $\tilde{\lambda}_t\tilde{\psi}_t$ are multipliers on the budget and collateral constraint, respectively. The optimal choice of the housing stock yields

$$\zeta_{B,t}\frac{\tilde{U}_{D,t}}{\tilde{U}_{C,t}} = \frac{P_{D,t}}{P_{C,t}} - (1-\delta)\left(\tilde{\psi}_t\zeta_{LTV,t}(1-\chi)\mathbb{E}_t\left\{\frac{P_{D,t+1}}{P_{C,t}}\right\} - \tilde{\beta}\mathbb{E}_t\left\{\frac{\tilde{U}_{C,t+1}}{\tilde{U}_{C,t}}\frac{P_{D,t+1}}{P_{C,t+1}}\right\}\right), \tag{3.6}$$

where $\tilde{U}_{D,t}$ denotes the marginal increase in utility from an extra unit of dwellings. $\zeta_{B,t}$ is a stationary AR(1) shock representing a near rational bubble process in housing prices as in In't Veld et al. (2011). In the spirit of Bernanke and Gertler (1999), this disturbance temporarily shocks the housing Euler Equation, which is the relevant asset Equation, and drives a wedge between the expected house price and the counterpart value under fully rational expectations. Hence, for housing investors, such a bubble is similar to a risk premium that is unrelated to fundamentals. By allowing only for small deviations from rational expectations on future fundamentals, we are able to introduce this stationary, non-fundamental disturbance and still can solve for the unique rational expectations equilibrium. Overall, $\zeta_{B,t}$ captures the ideas promoted, inter alia, in Shiller (2005, 2007), who calls for explanations of housing cycles beyond fundamentals and describes housing bubbles as periods of optimism followed by panic reactions, i.e., pessimism regarding future housing market conditions.

Finally, the demand for domestic and foreign produced non-durables read $\tilde{C}_{H,t} = \tau(P_{C,t}/P_{H,t})^{\iota}\tilde{C}_t$ and $\tilde{C}_{F,t} = (1-\tau)(P_{C,t}/P_{F,t})^{\iota}\tilde{C}_t$, with $P_{H,t}$ and $P_{F,t}$ denoting the price of consumption goods produced in country $i = H, F$. Thus domestic consumers' price index is a composite, i.e., $P_{C,t} = (\tau P_{H,t}^{1-\iota} + (1-\tau){P_{F,t}}^{1-\iota})^{\frac{1}{1-\iota}}$.

Savers' program

The continuum of saving households is $s \in [\omega, 1]$, where each saver (s) has the lifetime utility function

$$\mathbb{E}_0\left\{\sum_{t=0}^{\infty}\zeta_{\beta,t}\beta^t\left(\alpha\log(C_t(s) - hC_{t-1}) + (1-\alpha)\log D_t(s) - \frac{L_t(s)^{1+\eta}}{1+\eta}\right)\right\}, \tag{3.7}$$

and maximizes it subject to the following sequence of nominal budget constraints

$$P_{C,t}C_t(s) + P_{D,t}X_t(s) + P_{I,t}\sum_j^{C,D} I_{j,t}(s) + S_t(s) + B_t(s) = \sum_j^{C,D} \frac{W_{j,t}}{\mathcal{M}_{j,t}} L_{j,t}(s)$$
$$+ \sum_j^{C,D} R_{j,t}Z_{j,t}(s)K_{j,t-1}(s) - P_{I,t}\sum_j^{C,D} a\left(Z_{j,t}(s)\right)K_{j,t-1}(s) + R_{l,t}l(s) + R_{t-1}S_{t-1}(s)$$
$$+ R_{t-1}B_{t-1}(s) + \Pi_t''(s) + \Pi_t'''(s). \tag{3.8}$$

Savers have access to international bond markets, $B_t(s)$, which is not the case for domestic, borrowing households. $R_{l,t}l(s)$ are revenues from renting out land, $l(s)$, to producers in the construction sector at rate $R_{l,t}$. $\Pi_t''(s)$ and $\Pi_t'''(s)$ denote dividends obtained from intermediate goods firms and labor unions, respectively.[8] Moreover, savers invest in non-residential capital, $K_{j,t}(s)$, of sector $j = C, D$, where $I_{j,t}(s)$ is a composite of home and foreign non-durable investment goods defined as $I_{j,t}(s) = (\tau^{\frac{1}{\iota}} I_{H,t}^j(s)^{\frac{\iota-1}{\iota}} + (1-\tau)^{\frac{1}{\iota}} I_{F,t}^j(s)^{\frac{\iota-1}{\iota}})^{\frac{\iota}{\iota-1}}$. As the home country's weight, τ, is the same as in the counterpart index for consumption goods, it holds that $P_{I,t} = P_{C,t}$. Building on, e.g., Christiano et al. (2005) and Smets and Wouters (2007), saving households optimally decide on the capital utilization rate, $Z_{j,t}(s)$. Adjusting this intensive margin of capital is subject to cost, $a\left(Z_{j,t}(s)\right)$, where the cost function has the properties as in Pariès and Notarpietro (2008). $R_{j,t}$ is the rental price of capital in nominal terms, which determines savers' income from supplying the effectively used capital stock, $Z_{j,t}(s)K_{j,t-1}(s)$, to producers in sector $j = C, D$. Sector-specific capital accumulates over time as follows

$$K_{j,t}(s) = (1-\delta_j)K_{j,t-1}(s) + \left[1 - S\left(\frac{I_{j,t}(s)}{I_{j,t-1}(s)}\right)\right] I_{j,t}(s), \tag{3.9}$$

and depreciates with rate δ_j. Following Christiano et al. (2005), varying investment is costly, where $S(\cdot)$ is a cost function with $S(1) = S'(1) = 0$ and $S''(1) = \rho > 0$.

The solution to savers' decision problems with respect to their optimal choices of non-durable consumption and bond holdings results in the following FOC's

$$U_{C,t} = P_{C,t}\lambda_t \quad \text{and} \quad R_t^{-1} = \beta \mathbb{E}_t \left\{ \frac{\lambda_{t+1}}{\lambda_t} \right\}. \tag{3.10}$$

[8] Definitions of non-durable consumption goods and labor supply indices as well as consumption demand are analogue to those of borrowing households.

Optimal consumption of the housing good implies

$$\zeta_{B,t}\frac{U_{D,t}}{U_{C,t}} = \frac{P_{D,t}}{P_{C,t}} - \beta(1-\delta)\mathbb{E}_t\left\{\frac{U_{C,t+1}}{U_{C,t}}\frac{P_{D,t+1}}{P_{C,t+1}}\right\}. \tag{3.11}$$

Furthermore, savers optimize the stock of capital and its utilization rate as well as investment into sector-specific capital, which amounts to the subsequent FOC's

$$Q_{j,t} = \beta\mathbb{E}_t\left\{\frac{U_{C,t+1}}{U_{C,t}}\left[(1-\delta_j)Q_{j,t+1} + \left(\frac{R_{j,t+1}}{P_{C,t+1}}Z_{j,t+1} - a(Z_{j,t+1})\right)\right]\right\}, \tag{3.12}$$

$$Q_{j,t}\left[1 - S\left(\frac{I_{j,t}}{I_{j,t-1}}\right) - S'\left(\frac{I_{j,t}}{I_{j,t-1}}\right)\left(\frac{I_{j,t}}{I_{j,t-1}}\right)\right] =$$

$$1 - \beta\mathbb{E}_t\left\{Q_{j,t+1}\frac{U_{C,t+1}}{U_{C,t}}S'\left(\frac{I_{j,t+1}}{I_{j,t}}\right)\left(\frac{I_{j,t+1}}{I_{j,t}}\right)^2\right\}, \text{ and} \tag{3.13}$$

$$\frac{R_{j,t}}{P_{C,t}} = a'(Z_{j,t}), \tag{3.14}$$

where the real value of the existing capital stock, namely, Tobin's Q is $Q_{j,t}$.

Labor market

Households supply homogeneous labor, which monopolistically competitive unions differentiate as in Smets and Wouters (2007) and Iacoviello and Neri (2010). There is one union for each sector and country, where savers govern the unions as in Quint and Rabanal (2014). Unions sell labor services to wholesale labor packers that, ultimately, supply composite labor services to intermediate firms. Building on Erceg et al. (2000), unions face nominal wage rigidities in the form of a Calvo (1983) style lottery, where the fraction of unions receiving a wage setting signal is $\theta_{W,j}$, for $j = C, D$. Moreover, unions partially index wages to last period's price inflation of non-durable consumption goods as in Smets and Wouters (2003), with $\gamma_{W,j}$ measuring the sector-specific degree of indexation.

Unions' wage setting behavior yields the following Phillips curve for sectoral wages

$$\log\left(\frac{\omega_{j,t}}{\Pi_{C,t-1}^{\gamma_{W,j}}}\right) = \beta\mathbb{E}_t\left\{\log\left(\frac{\omega_{j,t+1}}{\Pi_{C,t}^{\gamma_{W,j}}}\right)\right\} - \frac{(1-\theta_{W,j})(1-\beta\theta_{W,j})}{\theta_{W,j}}\log\left(\frac{\mathcal{M}_{j,t}}{\mathcal{M}_j}\right). \tag{3.15}$$

$\Pi_{C,t} = P_{C,t}/P_{C,t-1}$ and $\omega_{j,t} = W_{j,t}/W_{j,t-1}$ are price inflation of non-durable consumption goods and gross wage inflation in sector $j = C, D$, respectively. Nominal, sectoral wages, $W_{j,t}$, include non-competitive wage markups, $\mathcal{M}_{j,t}$, which result from unions' monopoly power over wage setting and read for savers

$$\mathcal{M}_{C,t} = \frac{W_{C,t}}{P_{C,t}} \frac{U_{C,t}}{(1-\varrho)^{-\iota_L} L_t{}^{\eta-\iota_L} L_{C,t}^{\iota_L}} \quad \text{and} \quad \mathcal{M}_{D,t} = \frac{W_{D,t}}{P_{C,t}} \frac{U_{C,t}}{\varrho^{-\iota_L} L_t{}^{\eta-\iota_L} L_{D,t}^{\iota_L}}. \quad (3.16)$$

Thus the markups represent deviations of savers' marginal rate of substitution from sector-wide real wages.

By contrast, borrowing households are merely members of unions with no governing power. Therefore, they only adjust the amount of supplied labor services to the prescribed wage. Their sectoral optimality conditions read

$$\mathcal{M}_{C,t} = \frac{W_{C,t}}{P_{C,t}} \frac{\tilde{U}_{C,t}}{(1-\varrho)^{-\iota_L} \tilde{L}_t^{\eta-\iota_L} \tilde{L}_{C,t}^{\iota_L}} \quad \text{and} \quad \mathcal{M}_{D,t} = \frac{W_{D,t}}{P_{C,t}} \frac{\tilde{U}_{C,t}}{\varrho^{-\iota_L} \tilde{L}_t^{\eta-\iota_L} \tilde{L}_{D,t}^{\iota_L}}. \quad (3.17)$$

Final goods firms

Final goods bundlers operate under perfect competition with fully flexible prices. They buy intermediate goods $i \in [0, n]$ from firms of sector $j = C, D$ and combine them according to aggregator function

$$Y_{j,t} = \left(\frac{1}{n}\right)^{\lambda} \left(\int_0^n Y_{j,t}(i)^{\frac{1}{1+\lambda}} di\right)^{1+\lambda}. \quad (3.18)$$

$Y_{j,t}(i)$ represents type i intermediate goods, which bundlers employ for the production of the final goods, $Y_{j,t}$. λ is the net price markup (see, e.g., Smets and Wouters, 2003). Cost minimization of bundling firms gives rise to the following sector-specific demand Equations

$$Y_{C,t}(i) = \frac{1}{n}\left(\frac{P_{H,t}}{P_{H,t}(i)}\right)^{\frac{1+\lambda}{\lambda}} Y_{C,t} \quad \text{and} \quad Y_{D,t}(i) = \frac{1}{n}\left(\frac{P_{D,t}}{P_{D,t}(i)}\right)^{\frac{1+\lambda}{\lambda}} Y_{D,t}. \quad (3.19)$$

$P_{j',t}(i)$ and $P_{j',t}$, for $j' = H, D$, are domestic prices of sectoral intermediate and final products, respectively. Under zero profits in the final goods market the latter read

$$P_{j',t} = \left(\frac{1}{n}\right)^{-\lambda} \left(\int_0^n P_{j',t}(i)^{-\frac{1}{\lambda}} di\right)^{-\lambda}. \quad (3.20)$$

Intermediate goods firms

Building on Davis and Heathcote (2005) and Iacoviello and Neri (2010), we allow for sectoral heterogeneity of intermediate goods firms, which operate under monopolistic competition. The model introduces endogenous sectoral dynamics as a result of sector-specific production technologies

$$Y_{C,t}(i) = K'_{C,t}(i)^{\mu_C} L_{C,t}(i)^{1-\mu_C},\; Y_{D,t}(i) = \zeta_{AD,t} l(i)^{\mu_l} K'_{D,t}(i)^{\mu_D} L_{D,t}(i)^{1-\mu_l-\mu_D}. \quad (3.21)$$

$K'_{j,t}(i) = Z_{j,t}(i) K_{j,t-1}(i)$ denotes sectoral capital, effectively used in production, i.e., the accumulated stock of productive capital adjusted for time-varying capital utilization (see Smets and Wouters, 2007). μ_j, for $j = C, D$, are sectoral capital shares, and μ_l is the land share in the housing sector. $\zeta_{AD,t}$ is an AR(1) housing technology shock.

Firms in the intermediate goods sector solve a standard cost minimization problem, which results in the following sectoral marginal cost Equations

$$MC_{C,t}(i) = \frac{{R_{C,t}}^{\mu_C} {W_{C,t}}^{1-\mu_C}}{\mu_C^{\mu_C}(1-\mu_C)^{1-\mu_C}},\; MC_{D,t}(i) = \frac{{R_{l,t}}^{\mu_l} {R_{D,t}}^{\mu_D} {W_{D,t}}^{1-\mu_l-\mu_D}}{\mu_l^{\mu_l} \mu_D^{\mu_D} (1-\mu_l-\mu_D)^{1-\mu_l-\mu_D} \zeta_{AD,t}}. \quad (3.22)$$

The stock of land is fixed, i.e., $l_t = l$, and the interest for renting out land, $R_{l,t}$, is

$$R_{l,t} = \frac{\mu_l}{1-\mu_l-\mu_D} \frac{W_{D,t} L_{D,t}(i)}{l}, \quad (3.23)$$

where we choose l to yield equal sectoral wages as in Aspachs-Bracons and Rabanal (2011). Firms in the intermediate products sector earn subsequent profits

$$\Pi_{C,t}(i) = (P_{H,t}(i) - MC_{C,t}(i)) \left(\frac{1}{n}\right) \left(\frac{P_{H,t}(i)}{P_{H,t}}\right)^{-\frac{1+\lambda}{\lambda}} Y_{C,t} \text{ and} \quad (3.24)$$

$$\Pi_{D,t}(i) = (P_{D,t}(i) - MC_{D,t}(i)) \left(\frac{1}{n}\right) \left(\frac{P_{D,t}(i)}{P_{D,t}}\right)^{-\frac{1+\lambda}{\lambda}} Y_{D,t}, \quad (3.25)$$

where they maximize the expected value of these profits. In analogy to unions' wage setting process, intermediate firms face nominal rigidities. Thus in each sector a fraction of firms, $\theta_{P,j}$, is not able to set the profit maximizing price, $\dot{P}_{H,t}(i)$, as in Calvo (1983), but is allowed to partially index prices to sectoral price inflation as in Smets and Wouters (2003). The solution to non-durable sector firms' program is

$$\mathbb{E}_t \left\{ \sum_{k=0}^{\infty} \Lambda_{t,t+k} \theta_{P,C} Y_{C,t+k}(i) \left(\frac{\dot{P}_{H,t}(i)}{P_{H,t}} \frac{P_{H,t-1+k}^{\gamma_{P,C}}}{P_{H,t-1}^{\gamma_{P,C}}} \frac{P_{H,t}}{P_{H,t+k}} - (1+\lambda) \frac{MC_{C,t+k}(i)}{P_{H,t+k}} \right) \right\} = 0, \quad (3.26)$$

where firms discount future profits with factor $\Lambda_{t,t+k} = \beta^k(\lambda_{t+k}/\lambda_t)$, and $\gamma_{P,C}$ denotes the intensity of price indexation. The counterpart optimality condition for housing sector firms is analogue and reads

$$\mathbb{E}_t \left\{ \sum_{k=0}^{\infty} \Lambda_{t,t+k}\theta_{P,D} Y_{D,t+k}(i) \left(\frac{\dot{P}_{D,t}(i)}{P_{D,t}} \frac{P_{D,t-1+k}^{\gamma_{P,D}}}{P_{D,t-1}^{\gamma_{P,D}}} \frac{P_{D,t}}{P_{D,t+k}} - (1+\lambda)\frac{MC_{D,t+k}(i)}{P_{D,t+k}} \right) \right\} = 0. \tag{3.27}$$

Finally, we obtain the law of motion for domestic prices in the non-durable sector

$$P_{H,t}{}^{-\frac{1}{\lambda}} = \theta_{P,C} \left[P_{H,t-1} \left(\frac{P_{H,t-1}}{P_{H,t-2}} \right)^{\gamma_{P,C}} \right]^{-\frac{1}{\lambda}} + (1-\theta_{P,C})\dot{P}_{H,t}(i)^{-\frac{1}{\lambda}}, \tag{3.28}$$

and the housing sector

$$P_{D,t}{}^{-\frac{1}{\lambda}} = \theta_{P,D} \left[P_{D,t-1} \left(\frac{P_{D,t-1}}{P_{D,t-2}} \right)^{\gamma_{P,D}} \right]^{-\frac{1}{\lambda}} + (1-\theta_{P,D})\dot{P}_{D,t}(i)^{-\frac{1}{\lambda}}. \tag{3.29}$$

Market equilibrium

In equilibrium, home country production of non-durables equals borrowers' consumption demand as well as savers' consumption and investment demand

$$\begin{aligned} Y_{C,t} = n &\left(\omega\tilde{C}_{H,t} + (1-\omega)\left(C_{H,t} + I^C_{H,t} + I^D_{H,t} \right) \right) \\ &+ (1-n)\left(\omega^*\tilde{C}^*_{H,t} + (1-\omega^*)\left(C^*_{H,t} + I^{C*}_{H,t} + I^{D*}_{H,t} \right) \right) + \Omega_t, \end{aligned} \tag{3.30}$$

with Ω_t denoting resource cost, which result from time-varying utilization of the capital stock. The housing market clears under the following condition

$$Y_{D,t} = n\left(\omega\tilde{X}_t + (1-\omega)X_t \right). \tag{3.31}$$

With the definitions of housing and non-housing supply at hand, we obtain domestic GDP in real terms, i.e., $Y_t = Y_{C,t} + Y_{D,t}$. Sectoral labor markets clear as follows $\omega\tilde{L}_{j,t} + (1-\omega)L_{j,t} = \int_0^n L_{j,t}(i)di$, for $j = C, D$, and the equilibrium conditions of domestic and international debt markets are

$$\omega\tilde{S}_t = (1-\omega)S_t \quad \text{and} \quad n(1-\omega)B_t + (1-n)(1-\omega^*)B^*_t = 0. \tag{3.32}$$

Ultimately, the evolution of the domestic country's net foreign assets is

$$\begin{aligned} n(1-\omega)B_t = {} & n(1-\omega)R_{t-1}B_{t-1} \\ & + (1-n)P_{H,t}\left[\omega^*\tilde{C}^*_{H,t} + (1-\omega^*)\left(C^*_{H,t} + I^{C*}_{H,t} + I^{D*}_{H,t} \right) \right] \\ & - nP_{F,t}\left[\omega\tilde{C}_{F,t} + (1-\omega)\left(C_{F,t} + I^C_{F,t} + I^D_{F,t} \right) \right]. \end{aligned} \tag{3.33}$$

Monetary policy

The monetary authority perfectly controls the riskless bond rate in the monetary union, R_t^*, and follows an empirically motivated Taylor (1993) type instrument rule

$$\frac{R_t^*}{R^*} = \left(\frac{R_{t-1}^*}{R^*}\right)^{\mu_R} \left(\frac{\Pi_t^*}{\Pi^*}\right)^{\mu_\Pi(1-\mu_R)} \left(\frac{Y_t^*}{Y_{t-1}^*}\right)^{\mu_{\Delta Y}} \left(\frac{\Pi_t^*}{\Pi_{t-1}^*}\right)^{\mu_{\Delta\Pi}} \exp\left(\epsilon_{R,t}^*\right). \tag{3.34}$$

The central bank engages in interest rate smoothing, where μ_R measures the smoothness of interest rate policy. Moreover, the policy instrument reacts to deviations of the union-wide consumer price inflation, from its steady state, Π_t^*/Π^*, and to changes in output as well as the inflation rate as in Christoffel et al. (2008). μ_π, $\mu_{\Delta\pi}$, and $\mu_{\Delta Y}$ are the reaction coefficients. $\epsilon_{R,t}^*$ is a white noise monetary policy shock.

3.3.2 Deriving restrictions

As in Peersman and Straub (2009), we simulate the DSGE model 10,000 times by drawing uniformly distributed, random values for the structural parameters within specified intervals (Table 3.1).[9] Then we present median impulse responses together with 10 and 90 percent percentiles from all draws. For a pairwise comparison of shocks, finding at least one common and one opposed endogenous response that is robustly predicted by the different structural models, yields mutually exclusive restrictions, i.e., orthogonal shocks.

[9]We draw on empirical DSGE models like, e.g., Smets and Wouters (2003), Aspachs-Bracons and Rabanal (2011), in't Veld et al. (2014), and Coenen et al. (2008) to specify parameter ranges.

Table 3.1: Parameter intervals

Parameter	Description	Range
$\theta_{W,C}$	Wage stickiness: non-durable sector	[0.60, 0.90]
$\theta_{W,D}$	Wage stickiness: durable sector	[0.00, 0.30]
$\gamma_{W,C}$	Wage indexation: non-durable sector	[0.50, 0.90]
$\gamma_{W,D}$	Wage indexation: durable sector	[0.00, 0.40]
$\mathcal{M}_C$	Wage markup in steady state: non-durable sector	[1.10, 1.50]
$\mathcal{M}_D$	Wage markup in steady state: durable sector	[1.10, 1.50]
$1+\lambda$	Price markup in steady state	[1.10, 1.50]
h	Habit parameter	[0.40, 0.80]
η	Inverse Frisch elasticity	[1.50, 2.50]
ρ	Adjustment cost: investment	[1.00, 7.00]
υ	Degree of capital utilization	[0.10, 0.50]
$\theta_{P,C}$	Price stickiness: non-durable sector	[0.60, 0.90]
$\theta_{P,D}$	Price stickiness: durable sector	[0.00, 0.30]
$\gamma_{P,C}$	Price indexation: non-durable sector	[0.30, 0.90]
$\gamma_{P,D}$	Price indexation: durable sector	[0.00, 0.40]
μ_{Π}	Reaction coefficient: inflation	[1.15, 3.00]
μ_R	Interest rate smoothing	[0.50, 0.90]
$\mu_{\Delta Y}$	Reaction coefficient: change in output	[0.00, 0.30]
$\mu_{\Delta\Pi}$	Reaction coefficient: change in inflation	[0.00, 0.25]
ρ_B	Persistence: housing bubble shock	[0.95, 0.99]
ρ_{LTV}	Persistence: financial easing shock	[0.95, 0.99]
ρ_β	Persistence: savings glut shock	[0.40, 0.60]
ρ_{RP}	Persistence: risk premium shock	[0.95, 0.99]
ρ_{AD}	Persistence: housing technology shock	[0.95, 0.99]

Notes: The Table displays the parameter ranges employed to simulate the model.

Exogenous processes

We implement the four shocks from Section 3.2 in the DSGE model as follows.

- *Savings glut shock in the rest of the Eurozone.* Rest of union households become more patient compared to home country households. As in Sá and Wieladek (2015), we model the savings glut shock as a positive discount factor shock, $\zeta_{\beta,t}$, in Equations (3.1) and (3.7), describing lifetime utility of borrowers and savers, respectively.

- *Risk premium shock in the rest of the Eurozone.* This disturbance increases preferences of rest of union investors for home country bonds. It corresponds to a negative risk premium shock, $\zeta_{RP,t}$, in the net foreign asset Equation (3.4).

- *Financial easing shock in Spain.* This shock enhances credit availability against housing collateral of domestic borrowers and equals a positive shock, $\zeta_{LTV,t}$, in the collateral constraint Equation (3.3) and the housing Euler Equation (3.6).

- *Housing bubble shock in Spain.* As in In't Veld et al. (2011), this is a shock disturbing the risk premium on housing values and appears as $\zeta_{B,t}$ in domestic borrowers' and savers' housing Euler Equations (3.6) and (3.11).

Calibration strategy

For parameters governing nominal rigidities in goods and labor markets, we draw on the 90 percent posterior intervals of Smets and Wouters (2003). Calvo parameters, $\theta_{W,C}$ and $\theta_{P,C}$, range from 0.6 to 0.9.[10] Parameters capturing wage and price indexation, $\gamma_{W,C}$ and $\gamma_{P,C}$, vary from 0.5 to 0.9 and 0.3 to 0.9, repectively (see Aspachs-Bracons and Rabanal, 2011). We draw wage and price markups from 1.1 to 1.5, corresponding to elasticities of substitutions for differentiated goods and labor services ranging from 3 to 11 (Coenen et al., 2008). Following Sá and Wieladek (2015), Calvo housing parameters, $\theta_{P,D}$ and $\theta_{W,D}$, vary from 0 to 0.3 and indexation

[10] We expand the lower bound to 0.6 as the posterior intervals in Smets and Wouters (2003) do not include the popular values of $\theta_{W,C} = \theta_{P,C} = 0.75$.

parameters, $\gamma_{P,D}$ and $\gamma_{W,D}$, from 0 to 0.4, implying a more flexible housing compared to the non-durables sector. The degree of habit formation, h, ranges from 0.4 to 0.8 (see Smets and Wouters, 2003; in't Veld et al., 2014). For the inverse Frisch elasticity, η, we allow for variations from 1.5 to 2.5 (Coenen et al., 2008), while we set discount factors of savers, β, to 0.99 and borrowers, $\tilde{\beta}$, to 0.98. We rely on Smets and Wouters (2003) and Aspachs-Bracons and Rabanal (2011) for the capital bloc. Investment and capital utilization adjustment cost coefficients, ρ and ν, range from 1 to 7 and 0.1 to 0.5, respectively. The annual depreciation rate in the housing sector is 1 percent, and 10 percent in the non-durables sector. The capital share is 30 percent in the non-durables and 20 percent in the housing sector, while the land share is 10 percent in the housing sector. As in Aspachs-Bracons and Rabanal (2011), the cost coefficient of labor reallocation, ι, is 1.28, and the construction sector accounts for 10 percent of GDP in steady state. The LTV ratio, $1 - \chi$, is 0.8 (Akin et al., 2014) and the share of borrowing households, ω, is 0.4 (Hristov et al., 2012). The GDP weight of Spain in the Eurozone, n, is 0.1. Consistently, the fraction of Eurozone imports, $1 - \tau$, is 0.15, while the fraction of imports from Spain, τ^*, is 0.0167. Domestic bonds' risk premium elasticity with respect to the net foreign asset position, κ, varies from 0.002 to 0.007 (Quint and Rabanal, 2014) and the Taylor coefficients intervals encompass 90 percent of the posterior distributions from the ECB's New Area-Wide Model (Christoffel et al., 2008). As in Sá and Wieladek (2015), AR shock coefficients vary in persistent regions (Table 3.1), with standard deviations as in Aspachs-Bracons and Rabanal (2011).

Shock propagation

Figure 3.5 displays a financial easing shock.[11] A shock to the collateral constraint allows home country borrowers to increase credit against the expected value of housing, which raises borrowers' demand. Additionally, a relaxation of borrowing constraints fuels domestic absorption, in particular, in the non-durables sector.[12] Thus imports

[11] We calculate home country bond rates as a geometric average of short-term interest rates over a 10-year horizon as in Sá and Wieladek (2015).

[12] We analyze the dynamics for consumption instead of GDP, which allows us to isolate the impact on net exports – reflected by the current account – as well as on domestic absorption.

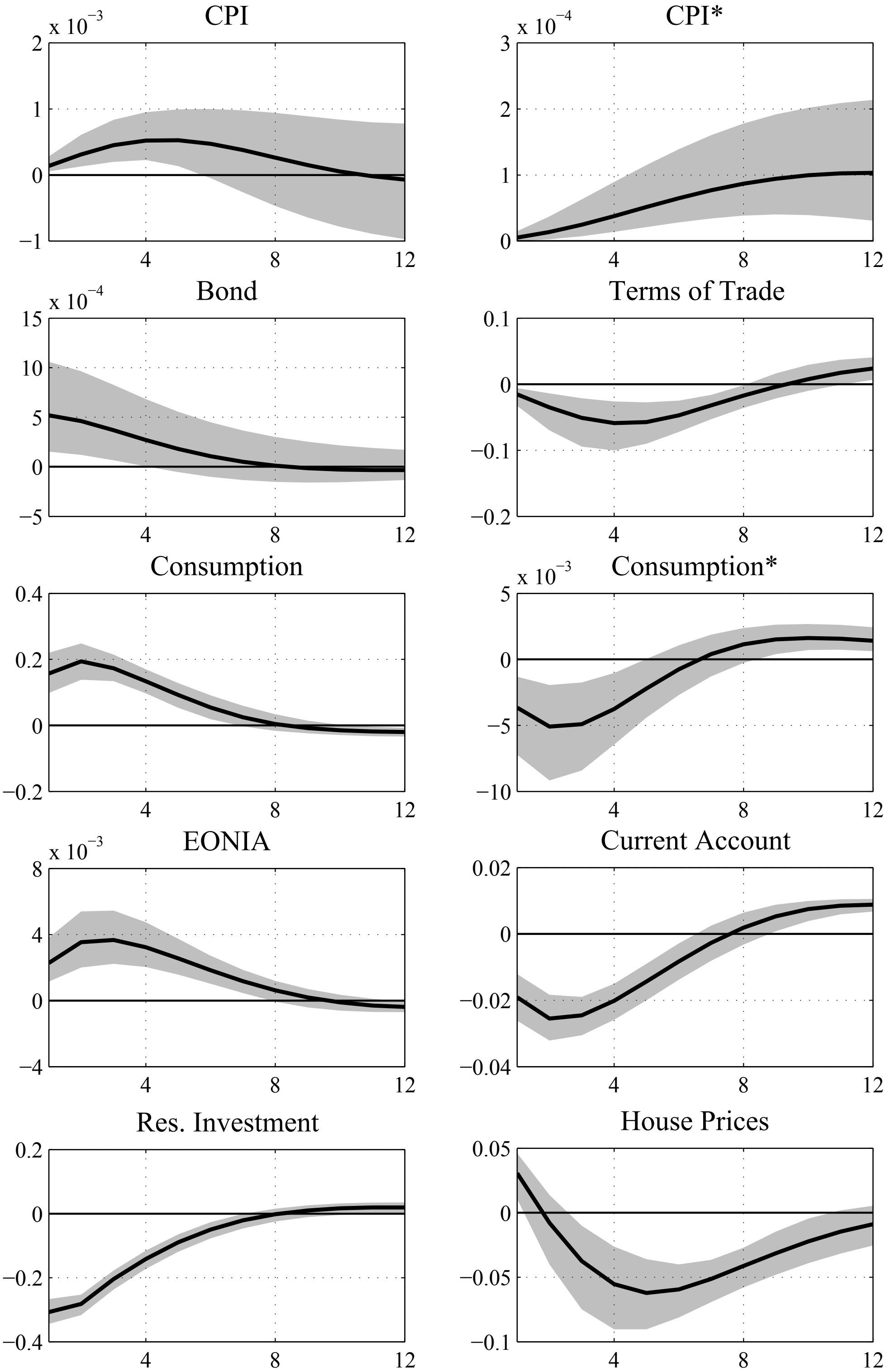
from the union increase, while exports shrink due to adverse terms of trade effects, i.e., the current account turns negative. A financial easing shock does not predict a boom in residential investment as enhanced borrowing capacities, predominantly, cause purchases of non-durables. Beyond, savers invest in housing, when prices are low. As house prices increase at short horizons due to the enhanced housing demand by borrowers, savers' residential investment drops, which overcompensates borrowers' investment in housing and, ultimately, also the house price increase. Furthermore, the central bank reacts to the financial easing shock by raising the policy rate, which translates into an increase of long-term bond rates in the home country.

In contrast, a housing bubble shock can account for a positive co-movement of residential investment and real house prices (see Figure 3.6). Furthermore, while the ratio of consumption to residential investment increases following a financial easing shock, it decreases after a housing bubble shock. We use this feature to disentangle the two shocks (see Figure 3.7). Overall, both pull shocks imply an increase in consumer price inflation and, accordingly, an increase in the policy instrument, which depresses consumption demand in the rest of the monetary union.

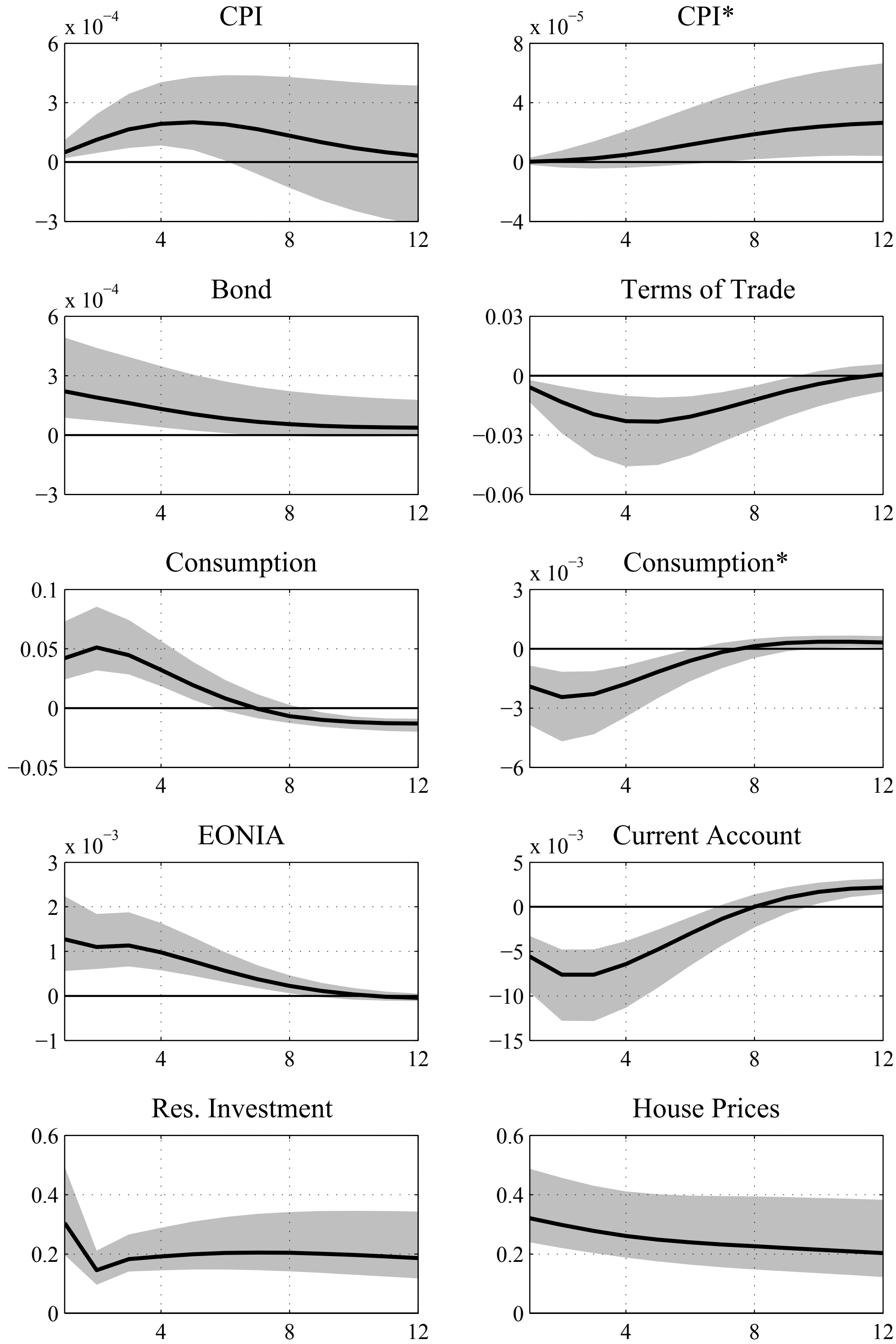
Figure 3.8 traces out the adjustment patterns following a risk premium shock. Rest of union investors have greater preferences for home country assets and invest to a larger extend into these bonds. Capital inflows cause bond rates to fall, which distinguishes the risk premium shock from the alternative pull disturbances. Lower interest rates, in turn, increase domestic absorption as savers and borrowers increase consumption and housing demand. The central bank responds to the home country boom with higher interest rates, which mildly depresses rest of union consumption.

Closely related to the risk premium shock is the savings glut shock (see Figure 3.9). However, in contrast to the risk premium shock, the simulations robustly predict a decline of short-term interest rates in the face of a savings glut shock. The surge of the discount factor in the rest of the union implies higher saving rates that in turn depress current economic activity, i.e., the savings glut shock represents a recessionary shock in the rest of the monetary union associated with a significant fall in consumer prices, and thus calls upon the central bank to decrease the policy instrument. As a consequence of the recession in the rest of the union with pronounced dis-inflationary effects, the CPI in the home country falls due to lower prices of imported goods – a facet of the savings glut shock, which further distinguishes this shock from the risk premium shock. Overall, due to asymmetric business cycles in the union, domestic interest rates are 'too low' triggering a boom in this economy. Lower interest rates, in addition, decrease borrowers' cost of financial services and relax borrowing constraints. This effect supports domestic absorption and reinforces a deterioration of the home country's current account.

As a robustness check, we consider two further disturbances to ensure orthogonality of the analysis with respect to these shocks. First, we simulate a monetary policy stimulus as a negative $\epsilon^*_{R,t}$ shock in Taylor rule Equation (3.34). As we calibrate deep parameter intervals in the currency union symmetrically, a cut in interest rates triggers no net capital flows. Moreover, the decline in interest rates leads to a consumption boom in both parts of the union as well as to higher union-wide consumer price inflation (Figure 3.10). Thus a monetary policy shock is inconsistent with the qualitative dynamics of the other disturbances. Second, we study an increase in home country's housing sector-specific technology, $\zeta_{AD,t}$, in Equation (3.21). Again, all considered structural models robustly predict an increase in domestic and foreign consumption making this shock orthogonal to the shocks under consideration.

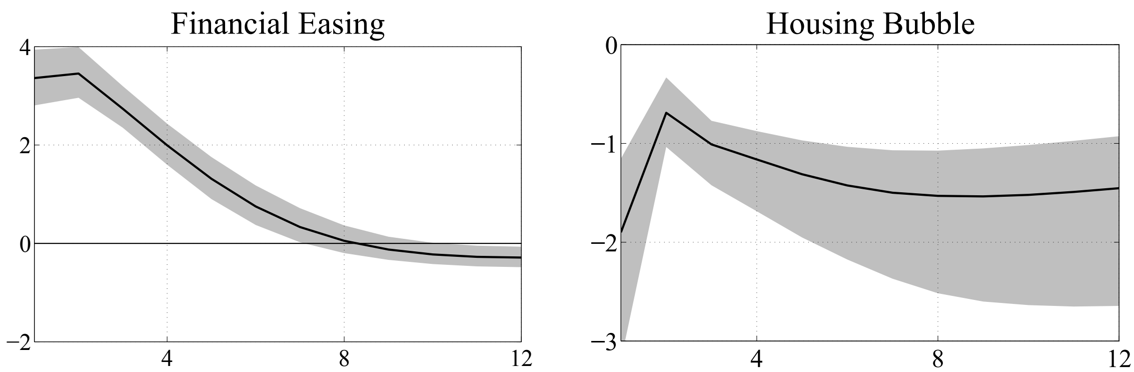


Notes: The x-axis is in quarters. The y-axis measures percent deviations from steady state. The solid line represents the median impulse response. Shaded areas display 10% and 90% percentiles of the simulated impulse responses.



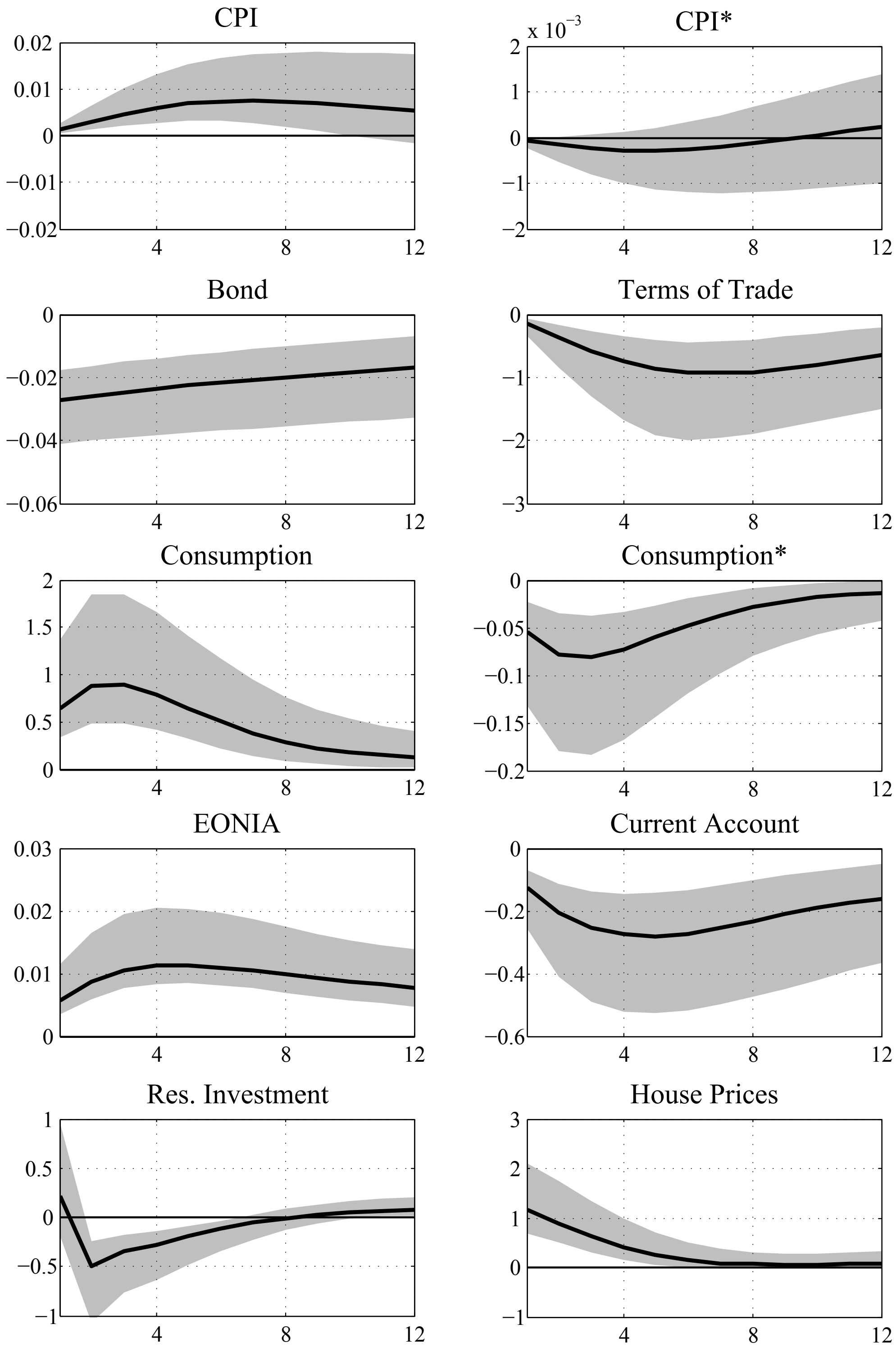
Notes: The x-axis is in quarters. The y-axis measures percent deviations from steady state. The solid line represents the median impulse response. Shaded areas display 10% and 90% percentiles of the simulated impulse responses.

Figure 3.7: Consumption to residential investment ratio

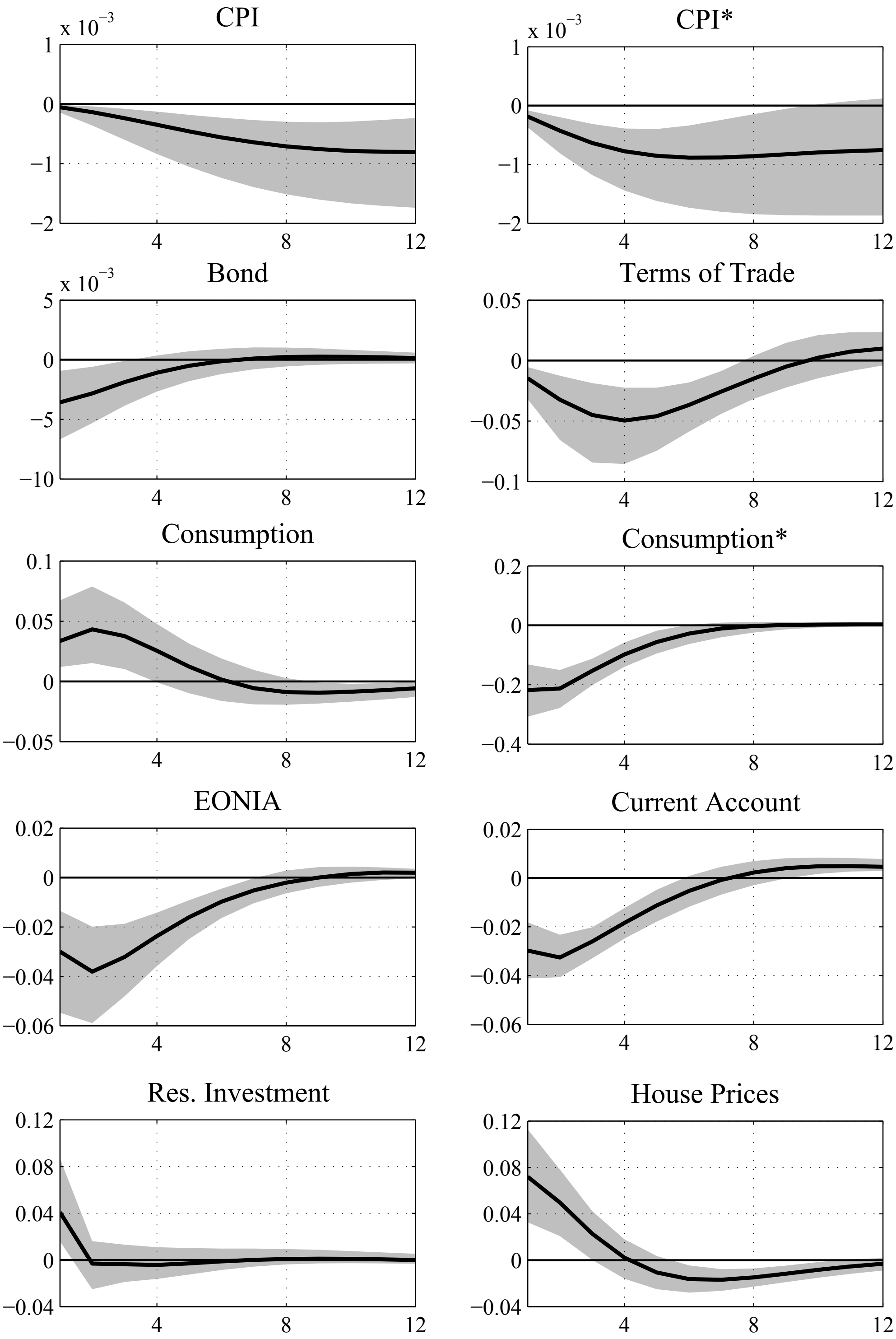


Notes: The x-axis is in quarters. The y-axis measures percent deviations from steady state. The solid line represents the median impulse response. Shaded areas display 10% and 90% percentiles of the simulated impulse responses.

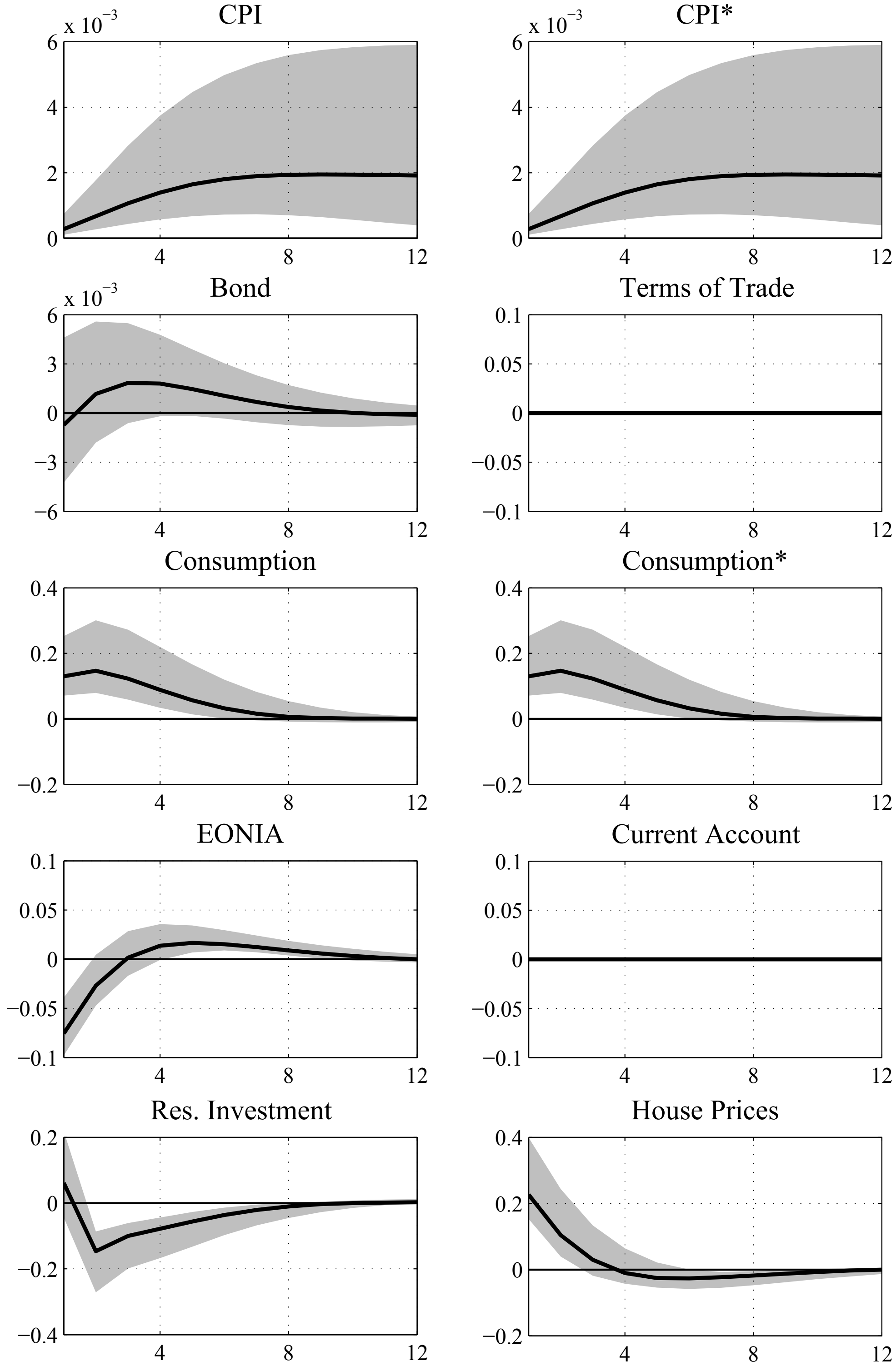
Figure 3.8: Risk premium shock



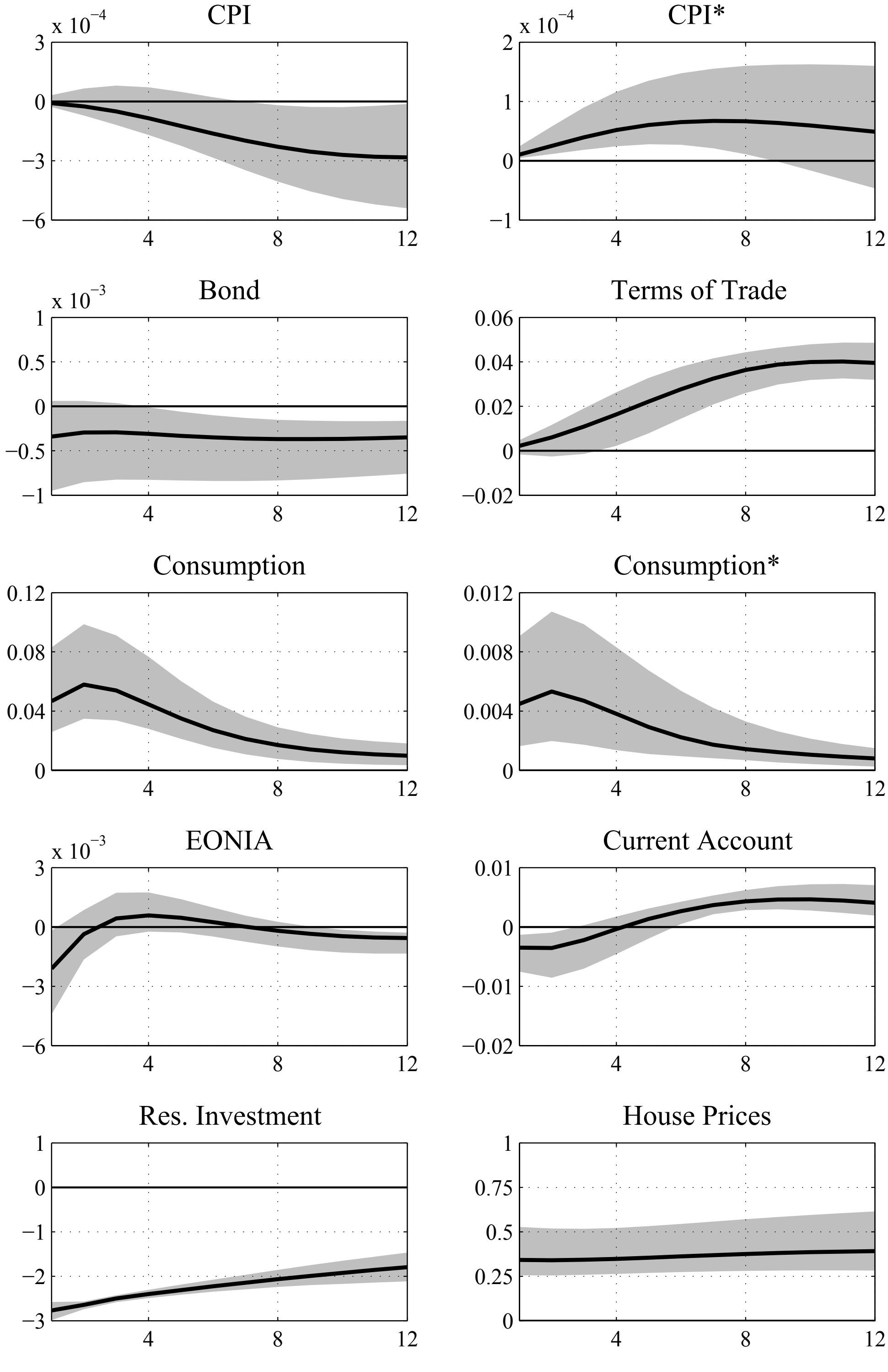
Notes: The x-axis is in quarters. The y-axis measures percent deviations from steady state. The solid line represents the median impulse response. Shaded areas display 10% and 90% percentiles of the simulated impulse responses.



Notes: The x-axis is in quarters. The y-axis measures percent deviations from steady state. The solid line represents the median impulse response. Shaded areas display 10% and 90% percentiles of the simulated impulse responses.



Notes: The x-axis is in quarters. The y-axis measures percent deviations from steady state. The solid line represents the median impulse response. Shaded areas display 10% and 90% percentiles of the simulated impulse responses.



Notes: The x-axis is in quarters. The y-axis measures percent deviations from steady state. The solid line represents the median impulse response. Shaded areas display 10% and 90% percentiles of the simulated impulse responses.

In summary, Table 3.2 displays the set of robust sign restrictions that assure orthogonality between the considered shocks and which we employ in the empirical analysis. As, e.g., in Sá and Wieladek (2015), we impose the restrictions for three quarters and do not impose restrictions on both housing market variables. For the current account, we only restrict the impact quarter in line with the DSGE model predictions to ensure that we isolate shocks, which coincide with a current account deterioration. In Section 3.5, we relax the restriction on the current account and test how our results are affected by this identification assumption.

Table 3.2: Benchmark sign restrictions

	Savings Glut	Financ. Easing	Risk Premium	Housing Bubble
Real Consumption	↑	↑	↑	↑
Real Consumption*	↓	↓	↓	↓
Prices	↓	↑	↑	↑
EONIA	↓	↑	↑	↑
Bond Rate	↓	↑	↓	↑
Bond Rate*				
Loans				
Current Account/GDP	↓	↓	↓	↓
Real House Prices				
Real Res. Investment				
Cons.-to-Investm.		↑		↓

Notes: Except for the current account, where we only restrict the impact quarter, we impose the restrictions for three quarters, i.e., 9 months as ≤ 0 or ≥ 0 (see, e.g., Sá and Wieladek, 2015).

3.4 Empirical methodology

In this section, we empirically analyze the effects of savings glut, risk premium, financial easing, and housing bubble shocks on the current account and the housing market in Spain. We begin with a description of the data and the estimation strategy. Using a Gibbs sampler, we estimate a mixed frequency VAR and draw efficient likelihood inference as in Eraker et al. (2014). In particular, the mixed frequency VAR approach is helpful given the short period of the housing cycle in Spain. Then we present the identification of structural shocks via sign restrictions as proposed in Uhlig (2005) and summarize the empirical findings.

3.4.1 Estimation, data, and inference

The analysis builds on the following reduced form open-economy VAR model

$$\boldsymbol{y}_t = \boldsymbol{c} + \sum_{l=1}^{p} \boldsymbol{\Phi}_l \boldsymbol{y}_{t-l} + \boldsymbol{\varepsilon}_t, \text{ where } \mathbb{E}[\boldsymbol{\varepsilon}_t] = 0 \text{ and } \mathbb{E}[\boldsymbol{\varepsilon}_t \boldsymbol{\varepsilon}_t{}'] = \boldsymbol{\Sigma}_\varepsilon. \tag{3.35}$$

$\boldsymbol{c}$ is a vector of intercepts, $\boldsymbol{\Phi}_l$ is a $n \times n$ matrix including AR coefficients at lag $l = 1, ..., p$, and $\boldsymbol{\Sigma}_\varepsilon$ is a $n \times n$ variance-covariance-matrix. $\boldsymbol{\varepsilon}_t$ represents one step ahead forecasting errors, and $\boldsymbol{y}_t$ comprises the following n endogenous variables

$$\boldsymbol{y}_t = [\ c_t\ c_t^*\ cpi_t\ eonia_t\ bond_t\ bond_t^*\ loans_t\ ca_t\ rinv_t\ cpih_t\ ]'. \tag{3.36}$$

The open-economy VAR framework is increasingly employed to study spillover effects from domestic shocks into foreign country aggregates, et vice versa (see, e.g., Fratzscher et al., 2010; Sá and Wieladek, 2015). Accordingly, we include Spanish data and time series for the rest of the Euro Area in $\boldsymbol{y}_t$.[13] cpi_t is the log level of the Harmonized Index of Consumer Prices (HICP). c_t denotes the cpi_t deflated log level of private consumption expenditures, and $bond_t$ measures nominal 10-year sovereign bond yields in percent. To calculate rest of Euro Area counterparts (indicated with *), we apply the household expenditure weights used by the HICP. These weights are updated annually and range from a share of 8.8 percent to 12.7 percent for Spain at

[13] An alternative is to specify data as country differentials by assuming symmetry across countries.

Euro Area expenditures.[14] $eonia_t$ represents interest rates in percent for unsecured, overnight lending in Euro Area interbank markets. As in Ciccarelli et al. (2015), we use $eonia_t$ instead of the interest rate on the ECB's main refinancing operations as proxy for the monetary policy stance. Following the financial turmoil of 2008 the ECB adopted various credit enhancing policies for banks, e.g., liquidity provisions with fixed interest rates and full allotment as well as longer-term refinancing operations, which temporarily pushed $eonia_t$ toward the ECB's deposit facility interest rate (see Lenza et al., 2010). Therefore, $eonia_t$, in contrast to the official policy rate, implicitly accounts for these liquidity management programs making it a reasonable policy measure especially since the financial crisis (Ciccarelli et al., 2015). As a measure of bank lending by Spanish banks, we include $loans_t$, which represents the outstanding stock of Euro denominated bank loans to the non-financial private sector in real terms. ca_t stands for the Spanish current account to GDP ratio in percent. $rinv_t$ and $cpih_t$ are log levels of real residential investment and a real house price index measuring residential property prices of all Spanish dwellings, respectively. Except for $cpih_t$, which we obtain from the BIS, all data come from Eurostat, the Bank of Spain, or the ECB. Consumption, price, and interest rate series primarily enter the VAR due to the identification of shocks, while we include the current account, loan volume, and housing variables to study the effects of capital inflows on the Spanish housing market. To pick up the EMU convergence period, we start the sample in 1995 M1 (see Crespo-Cuaresma and Fernández-Amador, 2013). We confine the estimation to 2013 M12 to avoid non-linearities caused by the zero lower bound on the nominal interest rate and provide robustness for the sample choice in Section 3.5.

Since the data sample is short, we employ a Bayesian mixed frequency approach for estimation and inference. In particular, for the case of short samples, Eraker et al. (2014) demonstrate that combining high frequency with low frequency time series yields efficiency gains compared to an estimator that discards high frequency information by relying on the coarsest data frequency for all variables. Thus we use

[14]See, e.g., Dees et al. (2007), who compare fix country weights with continuously varying weighting schemes in a GVAR analysis.

n_z quarterly series, $\boldsymbol{z}_t$, and, provided that they are available, n_x monthly series, $\boldsymbol{x}_t$, where $n_z + n_x = n$. Concretely, the subsets of $\boldsymbol{y}_t$ read

$$\boldsymbol{x}_t = [\ cpi_t\ eonia_t\ bond_t\ bond_t^*\ loans_t\ ]' \text{ and } \boldsymbol{z}_t = [\ c_t\ c_t^*\ ca_t\ rinv_t\ cpih_t\ ]' . \quad (3.37)$$

Following the Bayesian mixed frequency approach, we assume high frequency elements in $\boldsymbol{z}_t$ to be latent and hence consider them as missing realizations.[15] Using Markov-Chain-Monte-Carlo methods, the estimator alternately samples from latent observations and model parameters. Let $\widehat{\boldsymbol{z}}^i$ include low frequency data, observed as well as latent, for Markov-Chain-Monte-Carlo iteration i, where the sampled data are $\widehat{z}_1^i, \widehat{z}_2^i, \widehat{z}_4^i \ldots \widehat{z}_{T-1}^i$. Furthermore, let $\widehat{\boldsymbol{z}}_{-t}^i$ represent the complete vector $\widehat{\boldsymbol{z}}^i$ except for element $\widehat{z}_t^i$. As in Eraker et al. (2014), we proceed as follows. First, given initial values and using a conjugate Normal inverse Wishart prior for the parameters, we draw $\widehat{z}_t^i$ from a multivariate normal density, while conditioning on $\boldsymbol{x}_t$, $\widehat{\boldsymbol{z}}_{-t}^{i-1}$, $\boldsymbol{c}^{i-1}$, $\boldsymbol{\Phi}_l^{i-1}$, and $\boldsymbol{\Sigma}_\varepsilon^{i-1}$. Second, we draw $\boldsymbol{c}^i$ and $\boldsymbol{\Phi}_l^i$ for given $\boldsymbol{x}_t$, $\widehat{\boldsymbol{z}}^i$, and $\boldsymbol{\Sigma}_\varepsilon^{i-1}$, and third, we obtain $\boldsymbol{\Sigma}_\varepsilon^i$ by conditioning on $\boldsymbol{x}_t$, $\widehat{\boldsymbol{z}}^i$, $\boldsymbol{c}^i$, and $\boldsymbol{\Phi}_l^i$. Taking the temporal aggregation structure of low frequency variables in the VAR(p) into account, we computationally follow Qian (2013) and draw blocks of latent observations (aggregation cycle). We estimate the VAR with $p = 6$ lags, i.e., 2 quarters after linearly de-trending all series and provide robustness on the VAR specification in Section 3.5.[16] Note that the de-trending is motivated by the mixed frequency approach, which requires non-trending data. We experimented with different de-trending procedures, where results are robust to the concrete choice of methods.

3.4.2 Identification

From the VAR model in Equation (3.35), we derive impulse response functions to structural shocks by imposing sign restrictions (see, e.g., Faust, 1998; Canova and de Nicolo, 2002; Uhlig, 2005). Reduced form forecasting errors, $\boldsymbol{\varepsilon}_t$, linearly map

[15]See Ghysels (2015) for an alternative method of estimating mixed frequency VAR models within the mixed data sampling regression framework. In addition, Foroni and Marcellino (2013) offer a survey of mixed frequency data methods, in general.

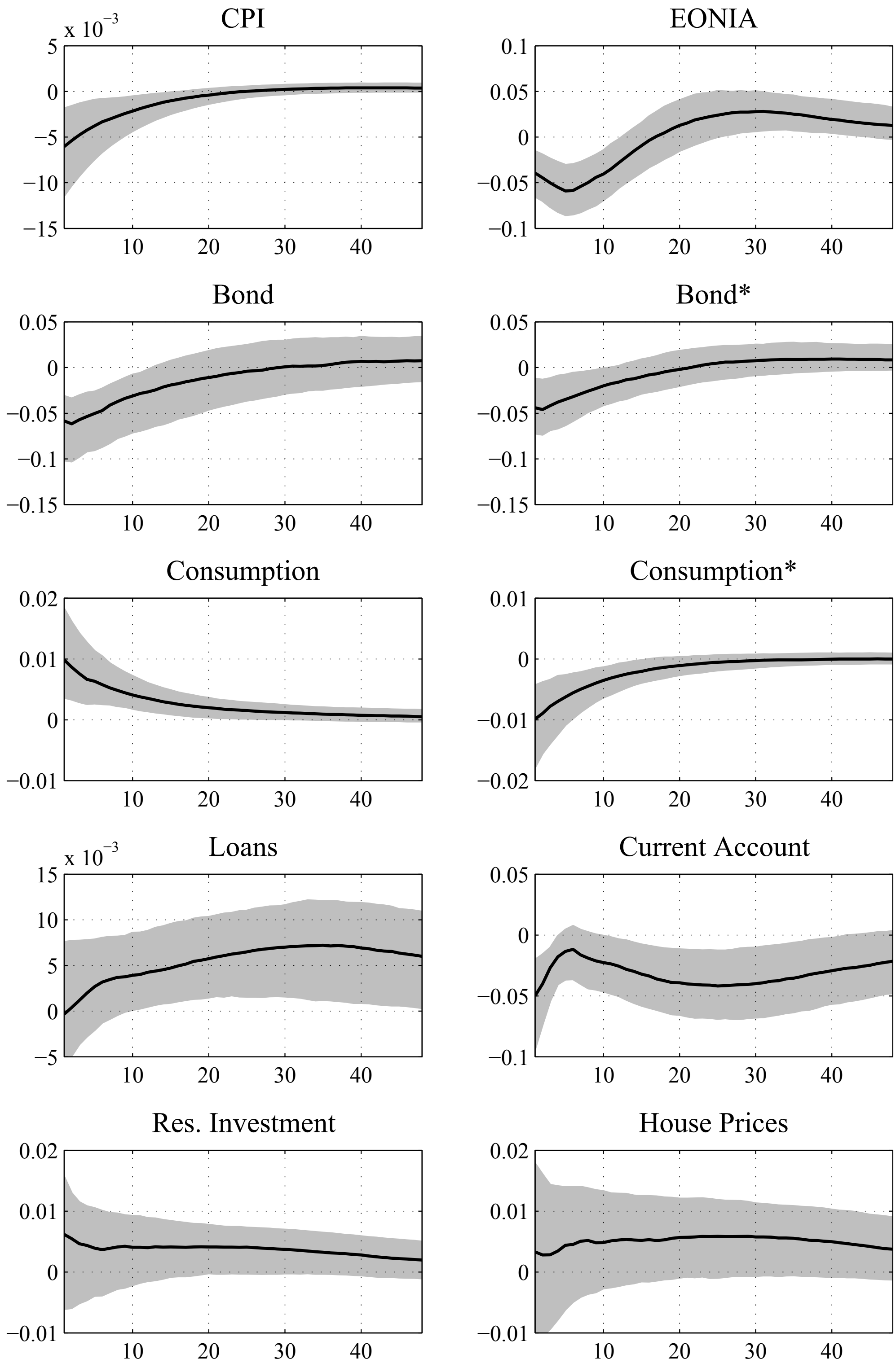
[16]We analyzed the sensitivity of the results with respect to the choice of the lag length by running the estimation for $p = 3, 9$, and 12 lags. For these specifications, the qualitative behavior of the key variables is not markedly affected. The results of these exercises are available upon request.

structural shocks, $\boldsymbol{\eta}_t$, through $\widetilde{\boldsymbol{P}}\boldsymbol{\eta}_t = \boldsymbol{\varepsilon}_t$, with $\mathbb{E}[\boldsymbol{\eta}_t] = 0$ and $\mathbb{E}[\boldsymbol{\eta}_t\boldsymbol{\eta}_t'] = \boldsymbol{\Sigma}_\eta$. $\boldsymbol{\Sigma}_\eta$ is diagonal ensuring orthogonality of the structural shocks. Furthermore, $\widetilde{\boldsymbol{P}} = \boldsymbol{PQ}$, where $\boldsymbol{P}$ represents one Cholesky factor from the Bayesian estimation. Hence, we can rewrite the variance-covariance-matrix of the reduced form model as $\mathbb{E}[\boldsymbol{\varepsilon}_t\boldsymbol{\varepsilon}_t'] = \boldsymbol{\Sigma}_\varepsilon = \boldsymbol{PQQ'P'}$, where $\boldsymbol{Q}$ is an ortho-normal matrix, i.e., $\boldsymbol{QQ'} = \boldsymbol{I}$. We obtain $\boldsymbol{Q}$ by applying the QR decomposition to a matrix $\boldsymbol{Z}$, which is sampled from a $\mathcal{N}(0,1)$ density. Each $\boldsymbol{Q}$ determines a different structural model and thus different impulse response functions. According to the sign restrictions approach, we derive impulse response functions for various structural models saving only those draws that are consistent with the imposed restrictions. As summary statistics, we then present the 16th, 50th, and 84th percentile of all accepted draws as in, e.g., Peersman (2005), Uhlig (2005), and Fratzscher et al. (2010).

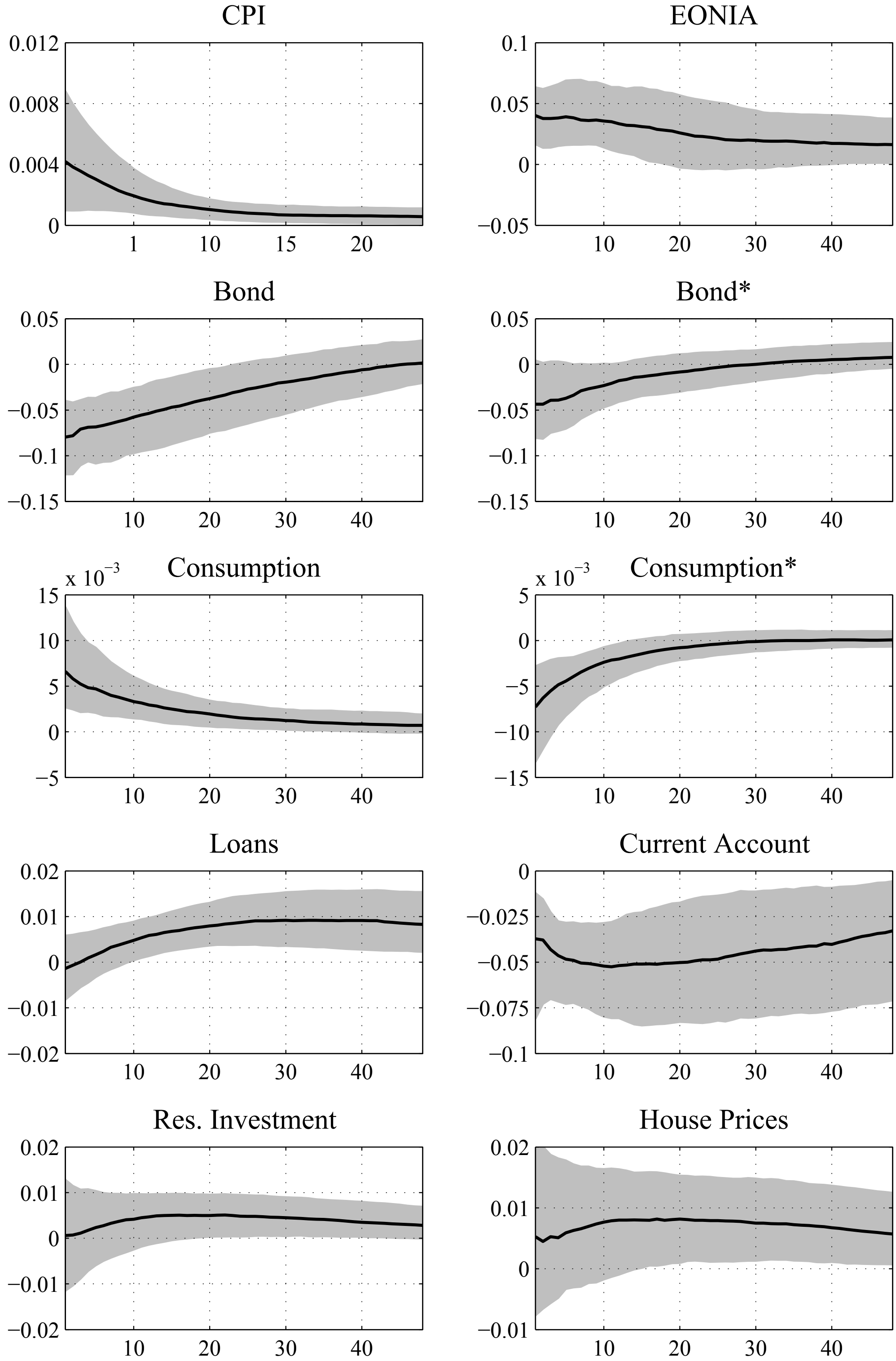
We simultaneously identify four types of macroeconomic shocks by imposing sign restrictions as summarized in Table 3.2 for nine months, i.e., three quarters (see, e.g., Sá and Wieladek, 2015). A broad class of open-economy DSGE models robustly predicts these restrictions. They are sufficient to disentangle the four shocks, and they ensure orthogonality to other disturbances (Section 3.3). As demonstrated in Paustian (2007) and Canova and Paustian (2011), we sharpen the identification by imposing more than the minimum set of sign restrictions, which increases the probability to isolate the shocks of interest. However, by leaving the responses of the housing market variables unrestricted, we remain agnostic about their dynamics.

3.4.3 Results

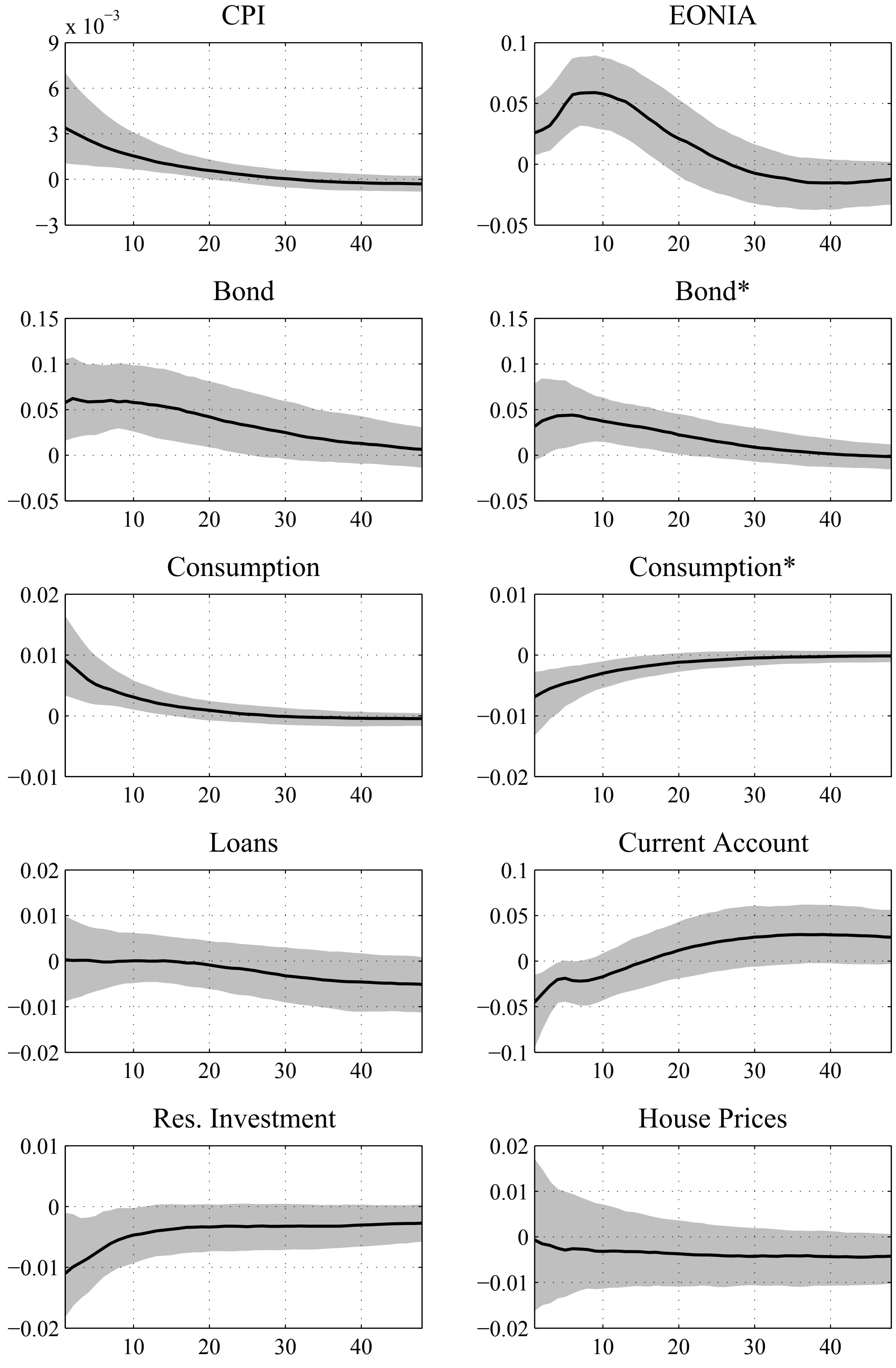
Figures 3.12 to 3.15 trace out the propagation of the identified shocks through the variables in $\boldsymbol{y}_t$. The shaded area denotes the 68 percent credible set from the Bayesian estimation and the solid line represents the median impulse response function. We report the dynamics for 48 months, i.e., for four years. We define all monthly shocks to reduce consumption in the rest of the Euro Area, i.e., c_t^* falls, as well as to incur a Spanish consumption boom, i.e., c_t increases.



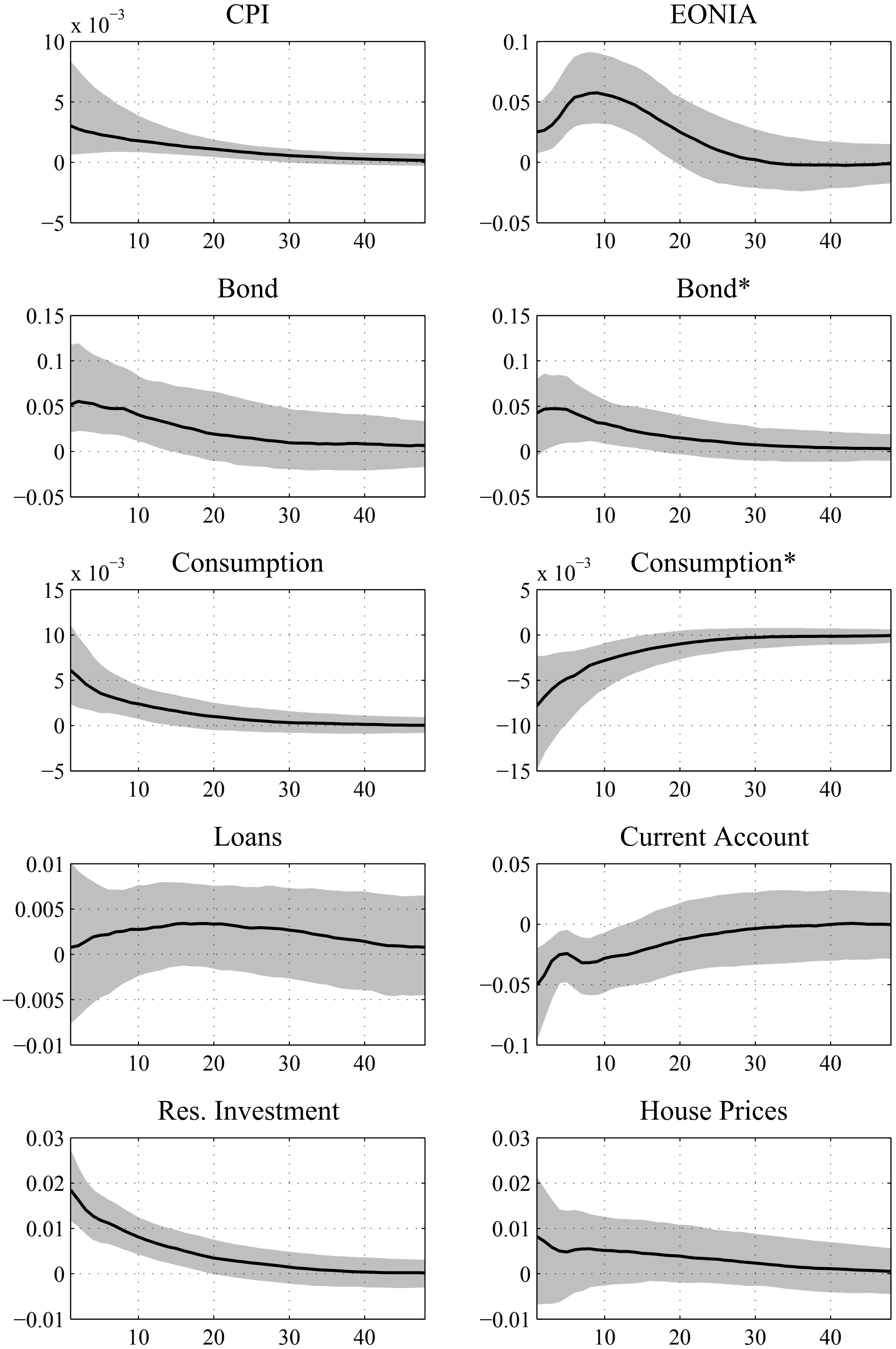
Notes: The x-axis is in months. The solid line represents the median impulse response functions from the BVAR. Shaded areas display 16% and 84% percentiles of the posterior distribution.



Notes: The x-axis is in months. The solid line represents the median impulse response functions from the BVAR. Shaded areas display 16% and 84% percentiles of the posterior distribution.



Notes: The x-axis is in months. The solid line represents the median impulse response functions from the BVAR. Shaded areas display 16% and 84% percentiles of the posterior distribution.



Notes: The x-axis is in months. The solid line represents the median impulse response functions from the BVAR. Shaded areas display 16% and 84% percentiles of the posterior distribution.

After a savings glut shock, the current account is significantly negative in the impact quarter in line with the imposed restriction (see Figure 3.12). Then, the response is insignificant for four months before significantly falling again for three and a half years. The unrestricted housing variables follow a sluggish increase, with median impulse responses being positive over the whole forecast horizon. Residential investment and house prices are, however, only significant at the margin. Furthermore, the unrestricted bank loans feature a slowly building rise, which remains significant from the second year onwards. Figure 3.13 displays adjustment patterns after a risk premium shock on Spanish bonds. This macroeconomic disturbance produces housing market and current account dynamics quantitatively similar to the savings glut shock. Though, this shock reveals more inertia with respect to the current account, which stays significantly different from zero over the entire forecast horizon. Beyond, the risk premium shock predicts significant increments for both housing variables after one and a half years. The bank loan response largely mirrors the dynamics after a savings glut shock, i.e., loans slowly build up and remain significantly positive after a year. The financial easing shock from Figure 3.14 only forces the current account into negative territory as long as we impose the sign restrictions, i.e., for three months. Then, the current account reponse is insignificant, with the median impulse response even overshooting the pre-shock level after one and a half years. Interestingly, the shock does not predict a boom with respect to both housing variables. While house prices are insignificant over the entire impulse horizon, residential investment even falls significantly in the first year after the shock and bank loans do not exhibit a significant reaction. Nevertheless, the DSGE model predicts this impact on housing markets (see Figure 3.5). From a theoretical perspective, financial easing shocks generally need not entail a housing boom as savers consume less housing, whereas borrowers increase the demand for housing. The overall impact on housing markets thus crucially hinges on the composition of households and their discount factors (see Justiniano et al., 2015). Altogether, the negative impulse response dynamics of residential investment together with the negative reaction of house prices (albeit insignificant) after a financial easing shock are hard to reconcile with the Spanish housing boom. As opposed to the financial easing shock, the housing bubble shock is capable of generating a negative correlation between the current account and all housing market variables in the VAR (see Figure 3.15). Most notably, residential investment immediately builds up in a statistically significant fashion for 20 months after the shock and house prices also increase at short horizons. Bank loans feature a hump-shaped increase which, however, is statistically insignificant.

Finally, we evaluate the relative importance of the shocks through the lens of a forecast error variance decomposition, which considers the estimated magnitude of the structural disturbances. For the variables of interest, entries in Table 3.3 reveal the fractions of the forecast error variance, which can be attributed to the respective shocks over various forecast horizons in percent. We present all k-step ahead forecast revisions for the median draw and report 68 percent credible sets. Overall, in terms of explanatory power for the housing market variables and the current account, we find fairly homogeneous results for the four identified shocks, with explained variance shares ranging in orders of magnitude similar to Sá and Wieladek (2015), who employ US data and use a similar identification scheme. With a share of 7.5 percent explained variation in real house prices, the savings glut shock has most explanatory power compared to the other disturbances for this variables, where its effect, primarily, is operative for the impact period. Furthermore, the savings glut shock explains more than 6 percent of the current account after 6 months and more than 5 percent of residential investment at longer horizons. The risk premium shock accounts for a similar share of fluctuations in house prices as the savings glut shock, however, the risk premium shock exerts its influence on house prices predominantly over longer forecast horizons. With respect to residential investment, the risk premium shock explains most variations of all shocks considered, with a maximum share of explained forecast revisions of more than 8.5 percent after four years. Albeit, the financial easing shock falls short in explaining the negative correlation of housing markets and the current account in Spain, this shock reveals some explanatory power for the key variables. With explanatory power of up to 8 percent for the current account, more than 7 percent for real house prices, and nearly 8 percent for residential investment, the magnitude of this shock is similar to both push disturbances. Ultimately, the housing bubble shock accounts for roughly 7 percent of the variation in real house prices after 6 months, while its explanatory power for the remaining variables and horizons is somewhat smaller than for the other shocks considered.

Table 3.3: Forecast error variance decomposition

	Horizon	Current Account	House Prices	Res. Investment
Savings Glut	Impact	3.88	7.49	5.00
Shock		(0.34, 17.60)	(0.79, 26.54)	(0.37, 21.8)
	6 Months	6.34	5.35	4.39
		(1.91, 16.68)	(1.27, 15.14)	(0.87, 18.90)
	12 Months	6.01	5.04	4.69
		(1.51, 18.17)	(1.22, 16.38)	(0.94, 18.73)
	24 Months	5.61	5.28	4.88
		(1.19, 17.34)	(1.18, 16.47)	(1.16, 18.32)
	48 Months	5.05	4.97	5.07
		(1.18, 16.40)	(1.30, 16.17)	(1.28, 17.62)
Risk Premium	Impact	5.63	4.79	5.84
Shock		(0.69, 22.49)	(0.55, 19.92)	(0.52, 21.65)
	6 Months	7.78	5.97	7.98
		(2.51, 19.57)	(1.05, 24.67)	(1.56, 22.65)
	12 Months	7.50	6.48	7.63
		(1.98, 19.89)	(1.16, 23.76)	(1.76, 22.80)
	24 Months	6.63	6.76	8.16
		(1.42, 21.53)	(1.16, 22.56)	(2.13, 22.43)
	48 Months	6.76	7.04	8.53
		(1.53, 21.54)	(1.37, 21.04)	(2.31, 21.94)
Financial Easing	Impact	6.72	6.23	3.40
Shock		(0.53, 18.59)	(0.61, 21.77)	(0.29, 16.63)
	6 Months	7.84	6.09	7.44
		(1.98, 28.52)	(1.04, 20.36)	(1.45, 25.48)
	12 Months	7.88	6.19	7.52
		(2.04, 17.88)	(1.08, 20.29)	(1.40, 24.94)
	24 Months	7.02	6.65	7.93
		(1.81, 19.00)	(1.25, 20.62)	(1.75, 23.14)
	48 Months	7.29	7.35	7.79
		(1.72, 20.90)	(1.32, 20.96)	(2.16, 22.59)
Housing Bubble	Impact	3.84	4.98	6.21
Shock		(0.31, 17.35)	(0.54, 19.68)	(0.76, 23.17)
	6 Months	5.92	5.76	4.72
		(1.66, 15.04)	(1.34, 19.91)	(0.81, 18.97)
	12 Months	5.88	5.61	4.83
		(1.70, 16.11)	(1.55, 18.82)	(0.82, 18.14)
	24 Months	5.79	5.65	5.23
		(1.29, 17.63)	(1.37, 18.70)	(1.07, 17.74)
	48 Months	5.88	6.07	5.33
		(1.19, 18.56)	(1.64, 18.93)	(1.41, 17.95)

Notes: Results are in percent for the median draw and we report the 68 percent credible sets in brackets.

In general, the analysis leaves substantial fractions of the forecast revisions in the key variables undeclared, i.e., explained by structural shocks that we do not identify. Our analysis, for instance, is orthogonal to macroeconomic disturbances emerging from, e.g., asymmetric housing technology dynamics or monetary policy shocks (see Section 3.3).[17]

3.5 Robustness

In this section, we extend the empirical analysis along several dimensions and review the robustness of our findings with respect to the modifications considered. First, we compare the findings with the so-called median target solution proposed in Fry and Pagan (2011). Second, for the identification of structural shocks, we allow for a different set of sign restrictions, which leaves the response of the current account unrestricted and, third, we analyze the shock propagation for a different data sample that stops in 2012 to filter out possible effects of Mario Draghi's *"whatever it takes"* speech in London on 26 July 2012.

3.5.1 Median target solution

Until this point, we rely on the median of all accepted impulse responses in the VAR to draw inference on. Yet, since the impulse responses of these point-wise posterior statistics need not necessarily be generated by the same structural model, we now calculate the median target solution as in Fry and Pagan (2011) and study the model dynamics for this particular model. The median target hereby refers to a single model producing impulses, which minimize the weighted distance to the median. Consequently, this model renders an interpretation feasible from a structural perspective.

Figure 3.16 displays the adjustment patterns of the key variables (columns) following the four identified shocks (rows) for the median model (solid line) together

[17] Furthermore, our modeling device of, e.g., the financial easing shock as a LTV shock in the DSGE model represents a lending shock in terms of quantities and thus excludes an also conceivable relaxation of bank lending standards in terms of prices, i.e., mortgage rates. Therefore, the easing shock is not able to explicitly capture all facets of eased lending standards emerging from, e.g., stronger competition within the banking sector.

with the median target solution (broken line), where we – here and in what follows – omit the remaining variables to conserve space. The impulse response functions of the median target model resemble the dynamics of the posterior median fairly close. Only for a small number of months, the median target response lies outside the 68 percent credible set (shaded regions) and thus is different from the median model in a statistical sense. Therefore, we conjecture that our inference as well as our main findings are not materially affected by considering the median instead of the median target solution.

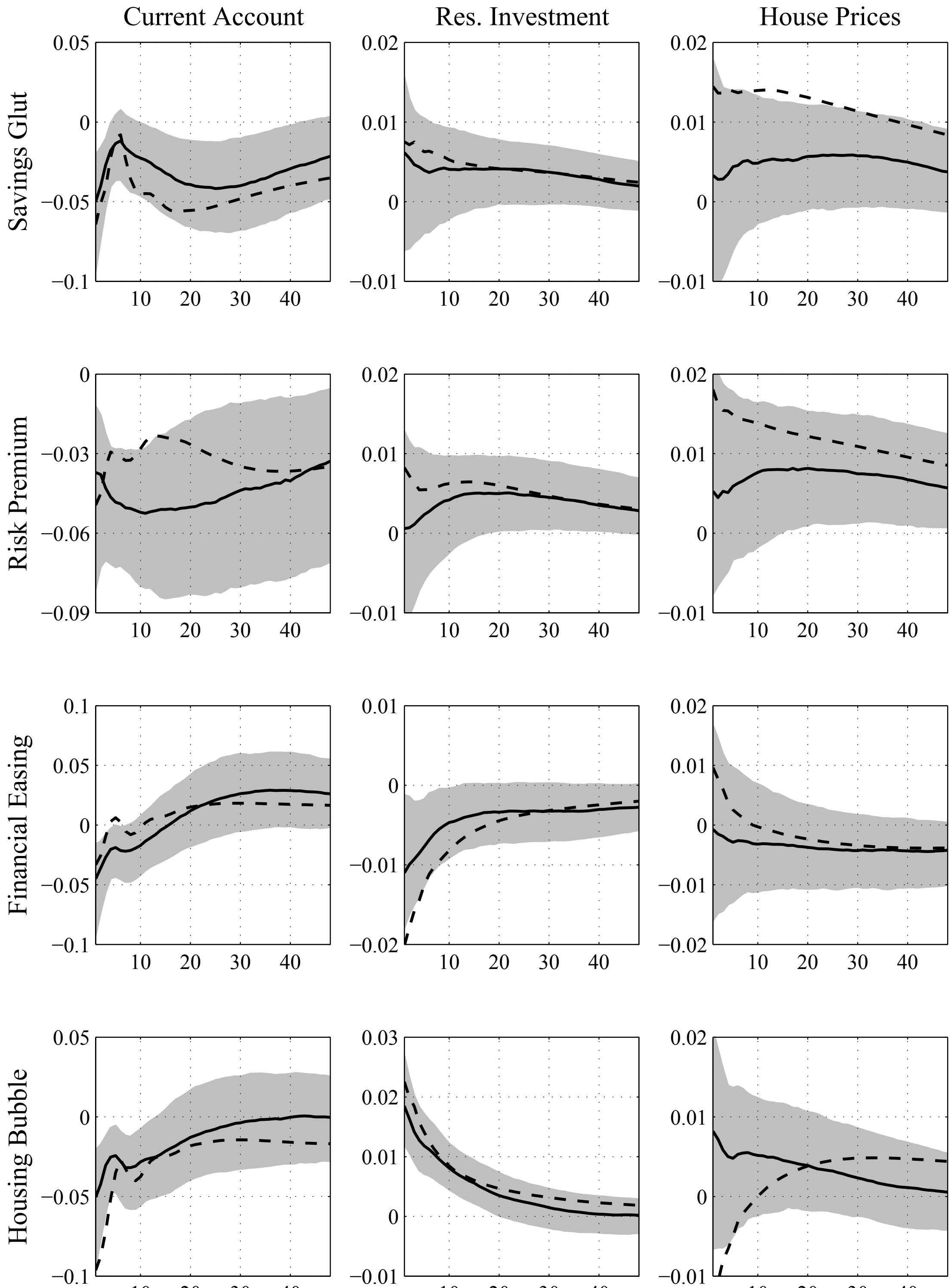
3.5.2 Alternative set of sign restrictions

The benchmark set of sign restrictions from Table 3.2 imposes restrictions on the current account, while leaving both housing market variables unrestricted. As a consequence, only the magnitude of the response is informative for the current account in the restricted impact quarter. In line with the DSGE model simulations from Figures 3.5 to 3.9, we can allow for an alternative identification strategy, which leaves the current account unrestricted over the whole forecast horizon. Instead, we impose a positive reaction on real house prices for all shocks in the impact quarter. We refrain from restricting higher impulse response horizons as such a restriction does not hold for the financial easing shock. The set of sign restrictions for this alternative identification scheme is summarized in Table 3.4.

Table 3.4: Alternative sign restrictions

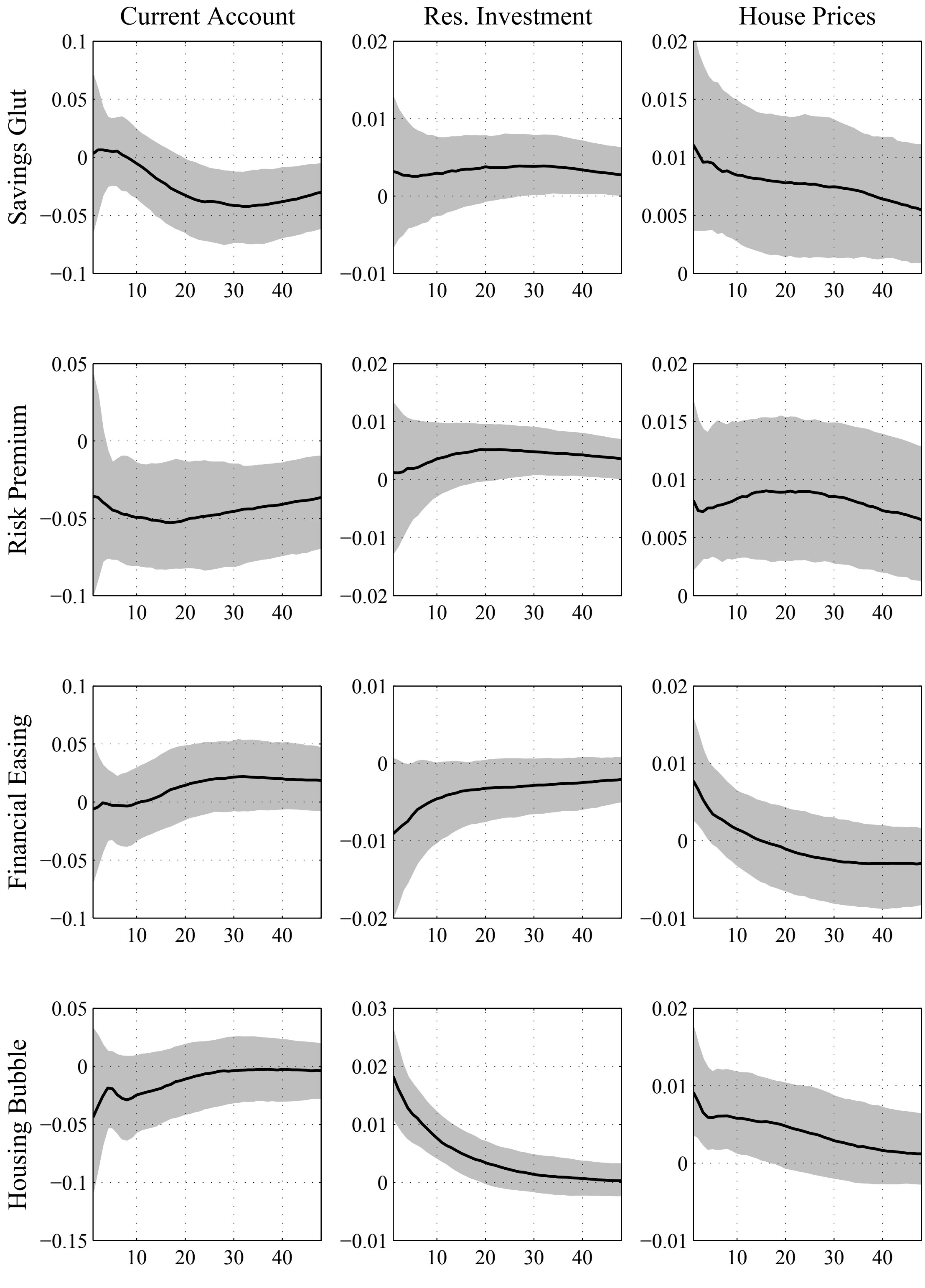
	Savings Glut	Financ. Easing	Risk Premium	Housing Bubble
Real Consumption	$\uparrow$	$\uparrow$	$\uparrow$	$\uparrow$
Real Consumption*	$\downarrow$	$\downarrow$	$\downarrow$	$\downarrow$
Prices	$\downarrow$	$\uparrow$	$\uparrow$	$\uparrow$
EONIA	$\downarrow$	$\uparrow$	$\uparrow$	$\uparrow$
Bond Rate	$\downarrow$	$\uparrow$	$\downarrow$	$\uparrow$
Bond Rate*				
Loans				
Current Account/GDP				
Real House Prices	$\uparrow$	$\uparrow$	$\uparrow$	$\uparrow$
Real Res. Investment				
Cons.-to-Investm.		$\uparrow$		$\downarrow$

Notes: Except for real house prices, where we only restrict the impact quarter, we impose the restrictions for three quarters, i.e., 9 months as ≤ 0 or ≥ 0 (see, e.g., Sá and Wieladek, 2015).



Notes: The x-axis is in months. The solid line represents the median impulse response functions from the BVAR. Shaded areas display 16% and 84% percentiles of the posterior distribution and the dashed line denotes the median target (see Fry and Pagan, 2011).

Figure 3.17 plots the VAR dynamics for the new identification scheme. Confirming the findings of the benchmark identification restrictions, the evidence of Figure 3.17 closely resembles the impulse responses of Figure 3.16. Due to the positive impact restriction on house prices, though, the latter rise significantly following a savings glut shock and remain significantly positive over the whole forecast horizon, while the residential investment response is little affected by the different identification scheme. Interestingly, for the savings glut shock, the unrestricted current account falls significantly after some quarters – a finding that, in particular, holds for the risk premium shock, which forces the current account to fall already in the impact period. The house price reaction is significant over a longer horizon for the risk premium shock, while the residential investment response, again, largely mirrors the results of the benchmark identification. For the housing bubble shock, the qualitative behavior of the impulse response functions does not change substantially for any of the key variables. Yet, for the financial easing shock, we now impose a positive impact reaction on house prices – a restriction for which we find no support in the benchmark specification where house prices are unrestricted and tend to fall in the data following the financial easing shock. Consistently, the imposed house price increase in the new identification scheme only holds for the impact quarter. Then, house prices overshoot the pre-shock level and turn negative, albeit insignificant. Residential investment still declines and, most notably, the financial easing shock predicts no current account deterioration, which resonates with the benchmark identification, where the current account turns positive after some quarters. In summary, our main findings are robust to this alternative identification assumptions.

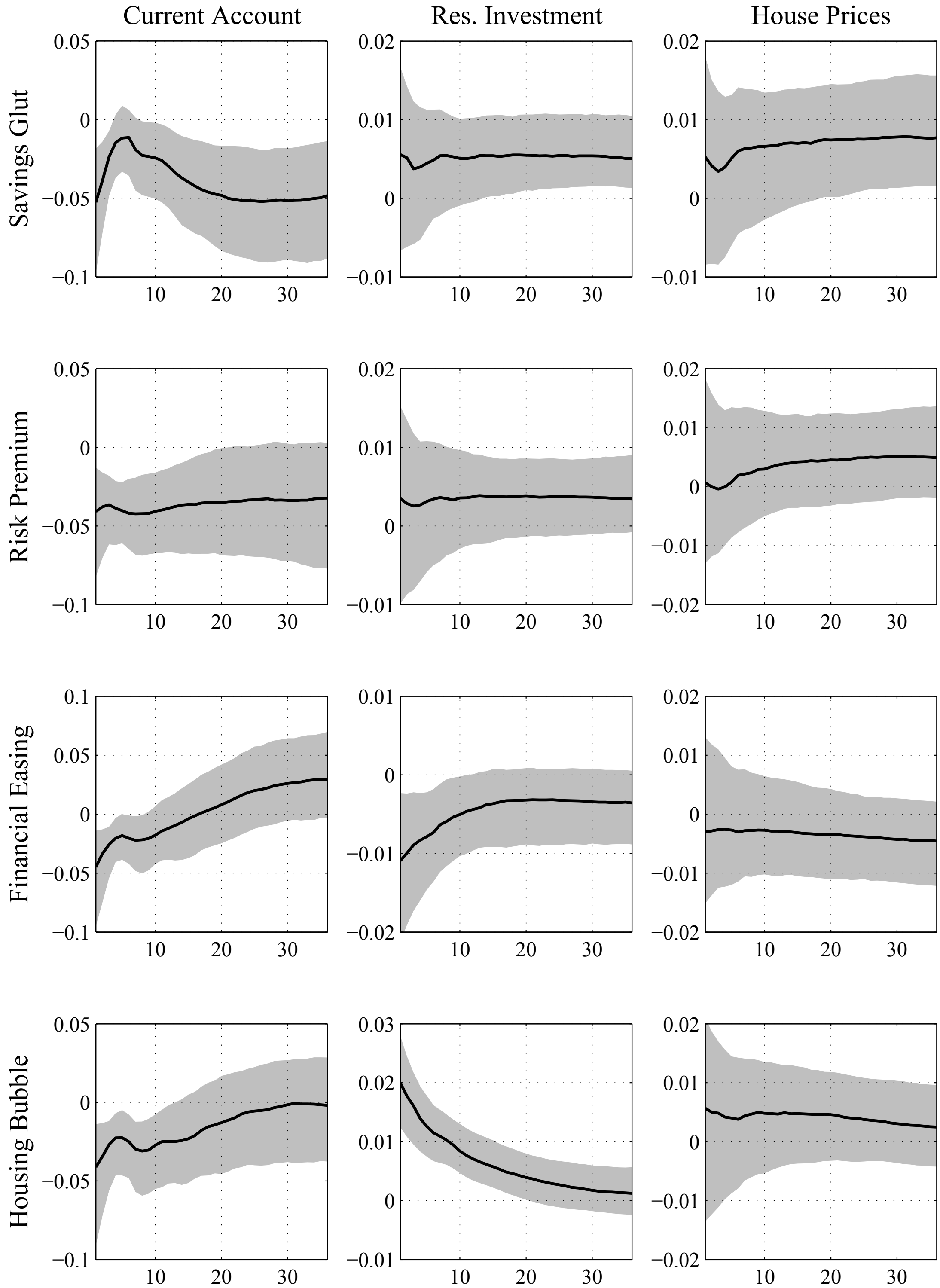


Notes: The x-axis is in months. The solid line represents the median impulse response functions from the BVAR. Shaded areas display 16% and 84% percentiles of the posterior distribution.

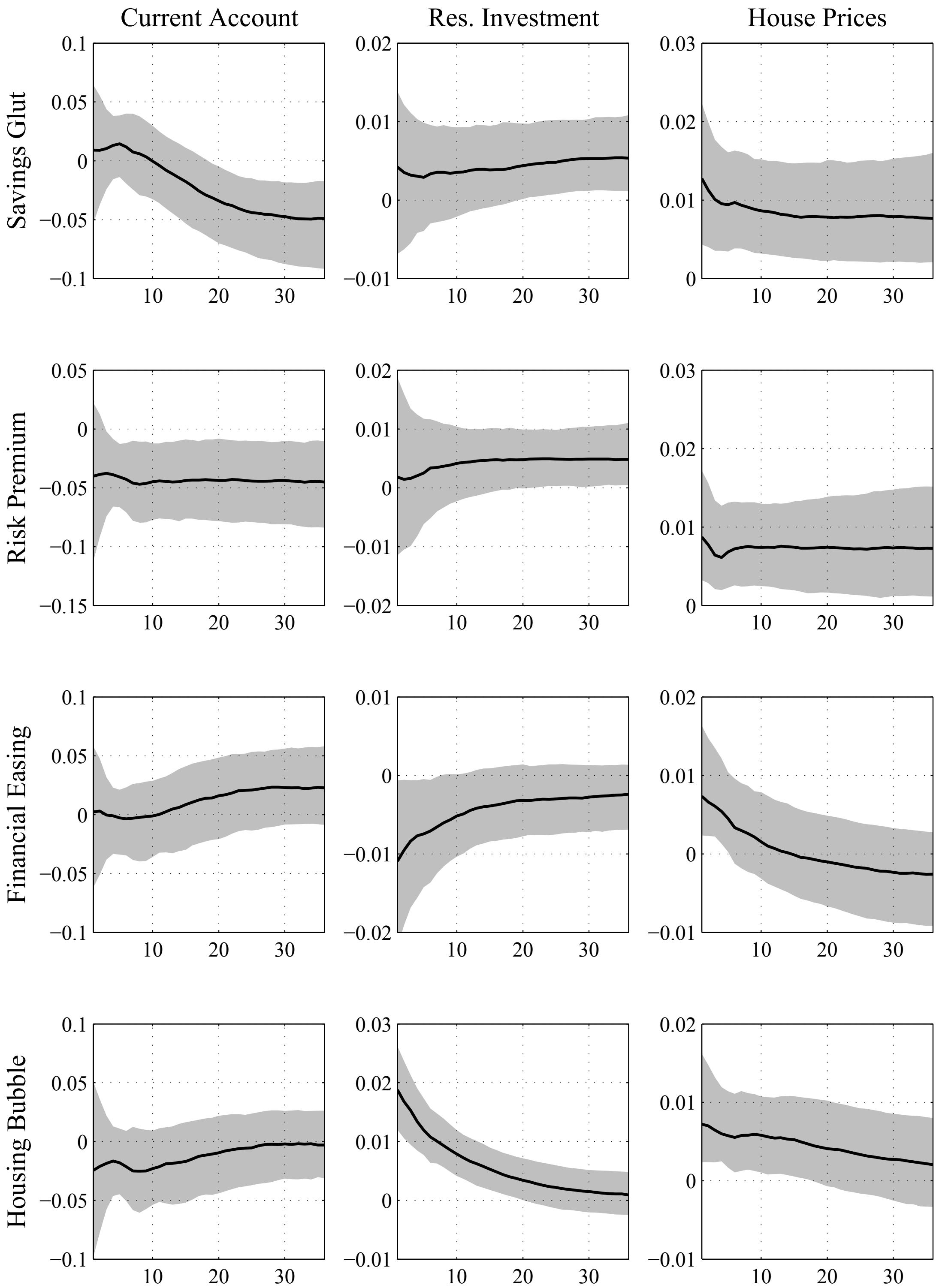
3.5.3 Different sample: Excluding Draghi's speech

Ultimately, we assess the main findings of the chapter against a modification of the data sample. During the Euro crisis, one particular moment stands out as being a game changer within the crisis. Namely, ECB president Draghi's speech at the Global Investment Conference in London on 26 July 2012. Explicitly stating that the Euro was *"irreversible"*, Draghi tended to convince market participants to view the Euro as an economic restriction, which should no longer be called into question, i.e., a given restriction around which market participants should optimize. Says Draghi: *"Within our mandate, the ECB is ready to do whatever it takes to preserve the euro. And believe me, it will be enough"*.[18] Following the speech, among others, risk premia on sovereign bonds of countries in the southern periphery of the Euro Area tumbled (see also Figure 3.3) contributing to a temporary calming of financial markets. We test to what extend our results are affected by the post-Draghi-speech sub-sample and re-run the regression with an ending date in June 2012, both, for the benchmark identification (Figure 3.18) and the identification scheme with an unrestricted current account and a positive impact restriction on house prices (Figure 3.19). Observing both figures, no notable differences compared to the benchmark sample emerge.

[18] See Draghi (2012) for the complete speech.



Notes: The x-axis is in months. The solid line represents the median impulse response functions from the BVAR. Shaded areas display 16% and 84% percentiles of the posterior distribution.



Notes: The x-axis is in months. The solid line represents the median impulse response functions from the BVAR. Shaded areas display 16% and 84% percentiles of the posterior distribution.

3.6 Conclusion

Since the late 1990's, two macroeconomic cycles, which hampered policy makers and attracted great interest of academics and the news media, have been characterizing the Spanish economy: the persistent buildup of a housing bubble and the pronounced deterioration of the current account. With the onset of the Great Recession, both developments reverted sharply. To our knowledge, we are the first to put different hypotheses to a test by quantitatively studying this joint co-movement in the data through the lens of an open-economy VAR that explicitly takes into account the specifics of a monetary union by deriving robust sign restrictions from a single currency area DSGE model. Savings glut, risk premium, and housing bubble shocks are able to generate the imbalances of Spain vis-à-vis the rest of the Eurozone and, at the same time, a housing boom in Spain. In contrast, financial easing shocks are neither capable of generating a distinct deterioration of the Spanish current account, nor of triggering a housing boom in Spain. In contrast, financial easing shocks are counterfactual to the housing boom, as a loosening of lending standards coincides with a cooling down in housing markets in our structural VAR analysis.

Chapter 4

State-Dependent Monetary Policy Transmission and Financial Market Tensions

4.1 Introduction

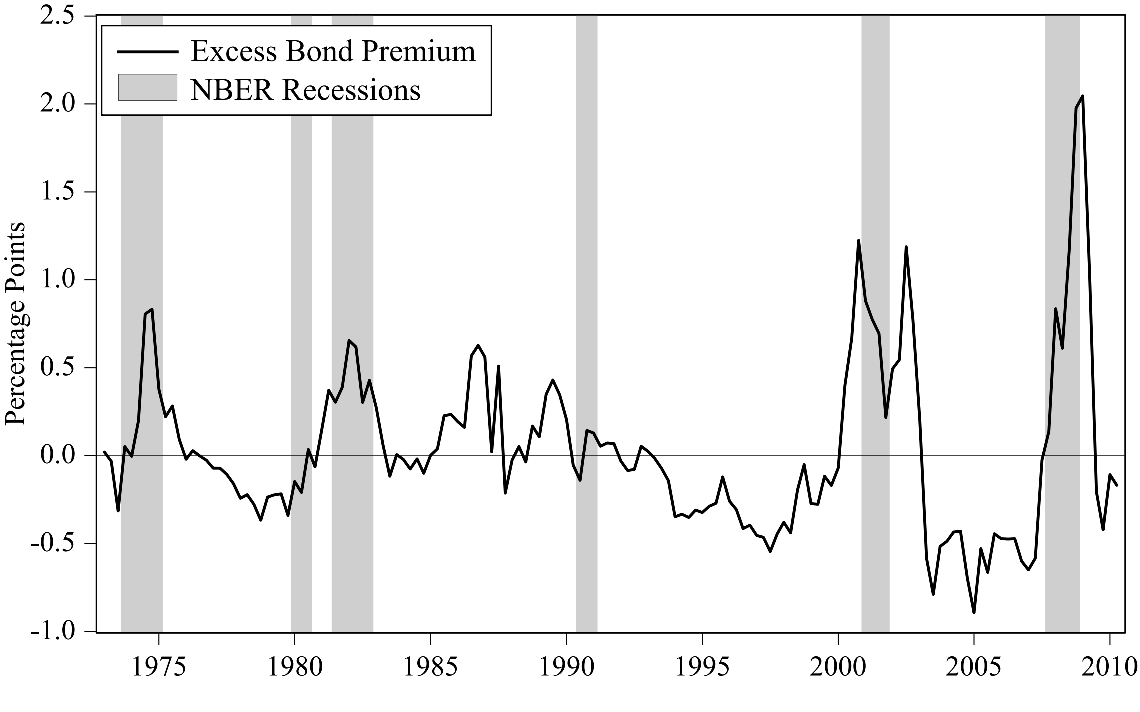
Existing empirical and theoretical models of the monetary policy transmission mechanism predominantly rely on linear frameworks assuming the impact of monetary policy on the broader economy to be a time-invariant, structural constant (e.g., Christiano et al., 2005; Smets and Wouters, 2007). Yet, a small strand of empirical contributions studies the effects of monetary policy by conditioning on the *business cycle*. These studies, however, do not appear to converge to a coherent view on whether monetary policy is more potent in expansions or recessions.[1] The divergence of results, among others, may be driven by the fact that the business cycle

[1] See the discussion and references in Tenreyro and Thwaites (forthcoming), who document that – due to, inter alia, a reinforcing fiscal policy stance – monetary policy might be more powerful in times of strong economic growth.

captures various interfering developments, which indeed affect the monetary policy pass-through, but which potentially inversely relate to and counteract each other. Candidates are time-variation in, e.g., the fiscal policy stance, price rigidities, or labor market frictions.

We approach the hypothesis of state-dependent monetary policy from a different perspective by conditioning monetary policy shocks and its dynamic effects, exclusively, on the degree of *financial market tensions* as measured by the Gilchrist and Zakrajsek (2012) excess bond premium (EBP) plotted in Figure 4.1.[2]

Figure 4.1: Excess bond premium



Notes: The figure plots the quarterly EBP as in Gilchrist and Zakrajsek (2012). Shaded areas represent NBER-dated recessions.

[2]Similarly, Born et al. (2015) study the impact of government consumption shocks conditional on the sovereign default premium, i.e., a measure of fiscal stress.

Gilchrist and Zakrajsek (2012) provide evidence that the EBP, which represents deviations of corporate versus Treasury bond spreads from expected default risks, constitutes a clean measure of risk attitudes and strains in financial markets as a whole. Using the Jordà (2005) local projections method and identifying monetary policy surprises in a non-linear equivalent to Romer and Romer (2004), we find that monetary policy impacts macroeconomic, housing, and financial variables several times stronger and more persistently in *high* relative to *low* financial frictions environments. In addition, positive dynamic responses of the EBP to a monetary tightening indicate a monetary-policy-induced amplification of strains in financial markets, in particular, when financial strains are *already high*.[3]

The excess sensitivity of macroeconomic aggregates to monetary policy innovations during high financial frictions regimes aligns with the *credit channel theory* of monetary policy (Bernanke and Gertler, 1995). A monetary policy contraction may raise private borrowing costs relative to government bond rates via tighter borrowing constraints in an environment of financial market imperfections – reflecting departures from the Modigliani-Miller axioms. In the presence of such frictions, the financial accelerator literature predicts the macroeconomic effects of monetary policy to be both, larger in magnitude *and* of higher persistence (Bernanke et al., 1999). Given time-variation in financial market frictions, as suggested by the EBP, the notion of state-, i.e., financial-strains-dependent monetary policy is consistent with this literature.

[3]These results are fully consistent with Dahlhaus (2014), who studies monetary policy transmission in normal versus financial crises episodes in a factor model with Cholesky identified monetary policy shocks.

4.2 Econometric framework

This section presents the econometric framework. Section 4.2.1 describes the empirical model specification and the approach to draw inference. Section 4.2.2 characterizes the state variable, summarizes the data, and explains the identification strategy.

4.2.1 Model specification and inference

To condition the dynamic effects of exogenous monetary policy surprises on the degree of strains in the financial system – proxied by the EBP – we apply local projections as in Jordà (2005) and allow the transition between states to smoothly evolve as in Granger and Teräsvirta (1994). Regarding the model specifications, we closely follow Tenreyro and Thwaites (forthcoming), which is the first study using local projections for the analysis of monetary policy.

We estimate the subsequent equation for variable, y_{t+s}, at horizons $s \in \{0, ..., S\}$, as a set of seemingly unrelated regressions

$$y_{t+s} = \delta t + (\alpha_s^h + \beta_s^h \varepsilon_t + \gamma(L)_s^{h\prime} x_{t-1})\Omega(z_t) + (\alpha_s^l + \beta_s^l \varepsilon_t + \gamma(L)_s^{l\prime} x_{t-1})[1 - \Omega(z_t)] + u_{t+s}. \tag{4.1}$$

δ denotes the slope of a linear trend, α_s^i are intercepts for regime $i \in \{h, l\}$, where h indicates a *high* and l a *low* degree of financial market tensions. γ_s^i and β_s^i measure the impact of a set of controls, x_t, and an identified monetary policy shock, ε_t, on y_{t+s}, i.e., the $s = 0, ..., S$ sequence of β_s^i-coefficients directly yields state-dependent impulse response functions to a time t shock at horizon $t + s$.[4] The weights assigned to each regime are determined by $\Omega(\cdot)$, which smoothly varies between 0 and 1 with the state variable, z_t, according to

$$\Omega(z_t) = \exp\left(\frac{\theta(z_t - \nu)}{\sigma_z}\right)\left[1 + \exp\left(\frac{\theta(z_t - \nu)}{\sigma_z}\right)\right]^{-1}, \tag{4.2}$$

where σ_z is the standard deviation of z_t. ν determines the fraction of periods characterized by state $i \in \{h, l\}$, and θ scales the propensity for regime shifts, with $\theta \to \infty$

[4] We set the lag length, L, to one according to the Schwartz Bayesian Criterion.

pushing $\Omega(\cdot)$ toward a binary weighting function. To draw inference on the moments of interest, we account for auto-correlation within and across regression residuals, u_t, arising from sequential leads of y_{t+s}. We follow Driscoll and Kraay (1998) to analytically correct the standard errors by, first, estimating the parameters in (4.1), and, second, calculating the average over moment conditions for s when computing the HAC robust variance-covariance matrix.

4.2.2 State variable, data, and identification

The state variable, z_t, is the seven quarter lagged moving average (see Tenreyro and Thwaites, forthcoming) of the EBP developed in Gilchrist and Zakrajsek (2012). Based on individual secondary market corporate bond prices, these authors calculate a credit spread index representing the difference between corporate bond and synthetic Treasury bond yields with identical cash flow and maturity characteristics. Applying a bond-spread pricing approach to these spreads purges the index from movements that capture the well-known counter-cyclical fluctuations in default expectations. This exercise, ultimately, yields the EBP, which reflects deviations of corporate bond pricing from ex-ante default risks and thus a gauge of effective risk attitudes or, more general, frictions in the financial system. We use the EBP to proxy financial market tensions as it, first, forecasts economic conditions more accurately relative to other spreads and, second, is purged from default risks. The latter feature allows for a distinction between classical interest rate pass-through of monetary policy, putatively also operating through default expectations, and pure financial accelerator effects (Gertler and Karadi, 2015). Third, being based on corporate relative to sovereign bond yields, the EBP is purified from the term premium and thus from expectations on future policy rates.

The core set of variables follows Gilchrist and Zakrajsek (2012) by comprising GDP, consumption, non-residential investment, inflation (GDP deflator), the EBP, the Federal Funds rate, 10-year Treasury bond yields, and excess stock market returns. In addition, we include residential investment, house price inflation (real Shiller index), and mortgage debt of households to study the monetary policy propagation into housing markets. All quantity series enter as log-levels (times 100), and we deflate these data with the GDP deflator and divide them by the civilian non-institutional population. Interest rates and spreads are in percent, while inflation rates and excess market returns are percentage growth rates. We measure the data at the quarterly frequency and start the sample, due to the availability of EBP, in 1973Q1, while restricting the analysis to 2008Q4, whereupon the Federal Funds rate sticks to its zero lower bound.

We follow Romer and Romer (2004) to uncover structural monetary policy innovations by regressing changes in the Federal Funds rate, Δi_t, on an updated set of the Romer and Romer (2004) controls, $x_t^{\mathrm{R\&R}}$, and obtain putatively orthogonal monetary policy shocks, ε_t, as the regression errors. Yet, as in Tenreyro and Thwaites (forthcoming), we allow the FED's reaction function, i.e., the slopes of the controls, χ^i, with $i \in \{h, l\}$, to be state-dependent and – consistent with (4.1) – to smoothly alternate with z_t as follows[5]

$$\Delta i_t = \chi^h x_t^{\mathrm{R\&R}} \Omega(z_t) + \chi^l x_t^{\mathrm{R\&R}} [1 - \Omega(z_t)] + \varepsilon_t. \tag{4.3}$$

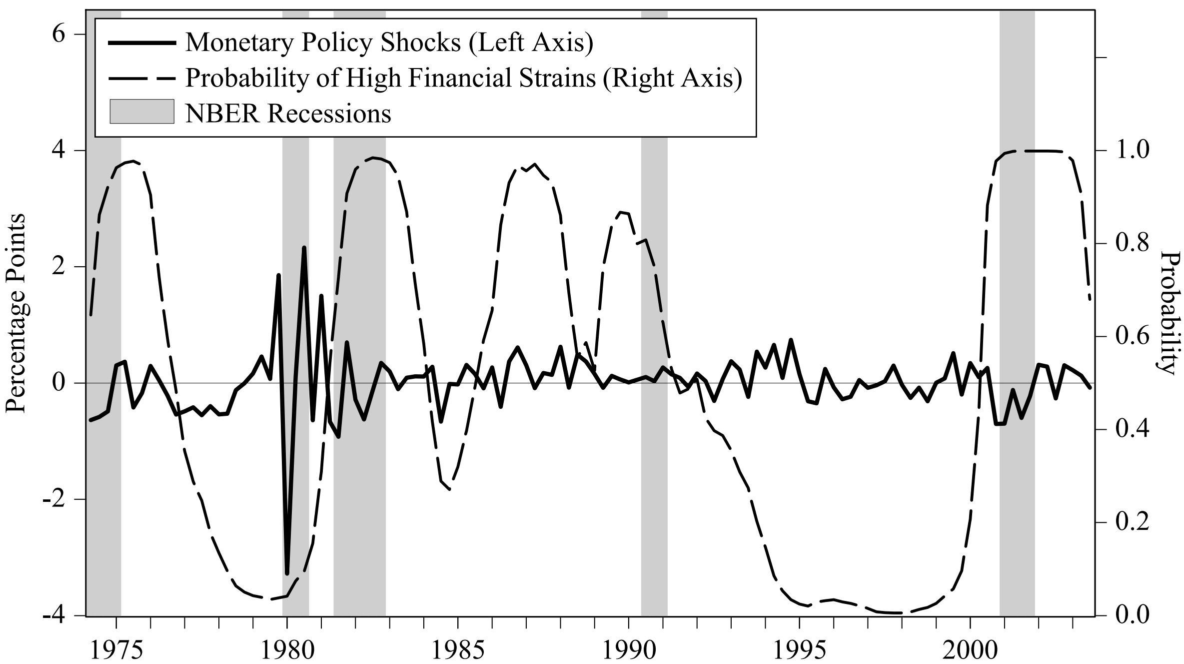
4.3 Results

Figure 4.2 plots $\Omega(z_t)$ and ε_t together with NBER-dated recessions across time, where we calibrate $\theta = 3$ (Tenreyro and Thwaites, forthcoming) and ν such that 50% of quarters are characterized by regime h and l, respectively, which is in line with the observation of roughly half of the EBP realizations in the sample being positive (negative).[6]

[5]We use the data set and code from Tenreyro and Thwaites (forthcoming), which is available on the publisher's webpage (AEJ: Macroeconomics).

[6]Even proportions of regimes, in addition, yield higher accuracy for parameter estimates (e.g., Bernardini and Peersman, 2015).

Figure 4.2: Indicator function and monetary policy shocks

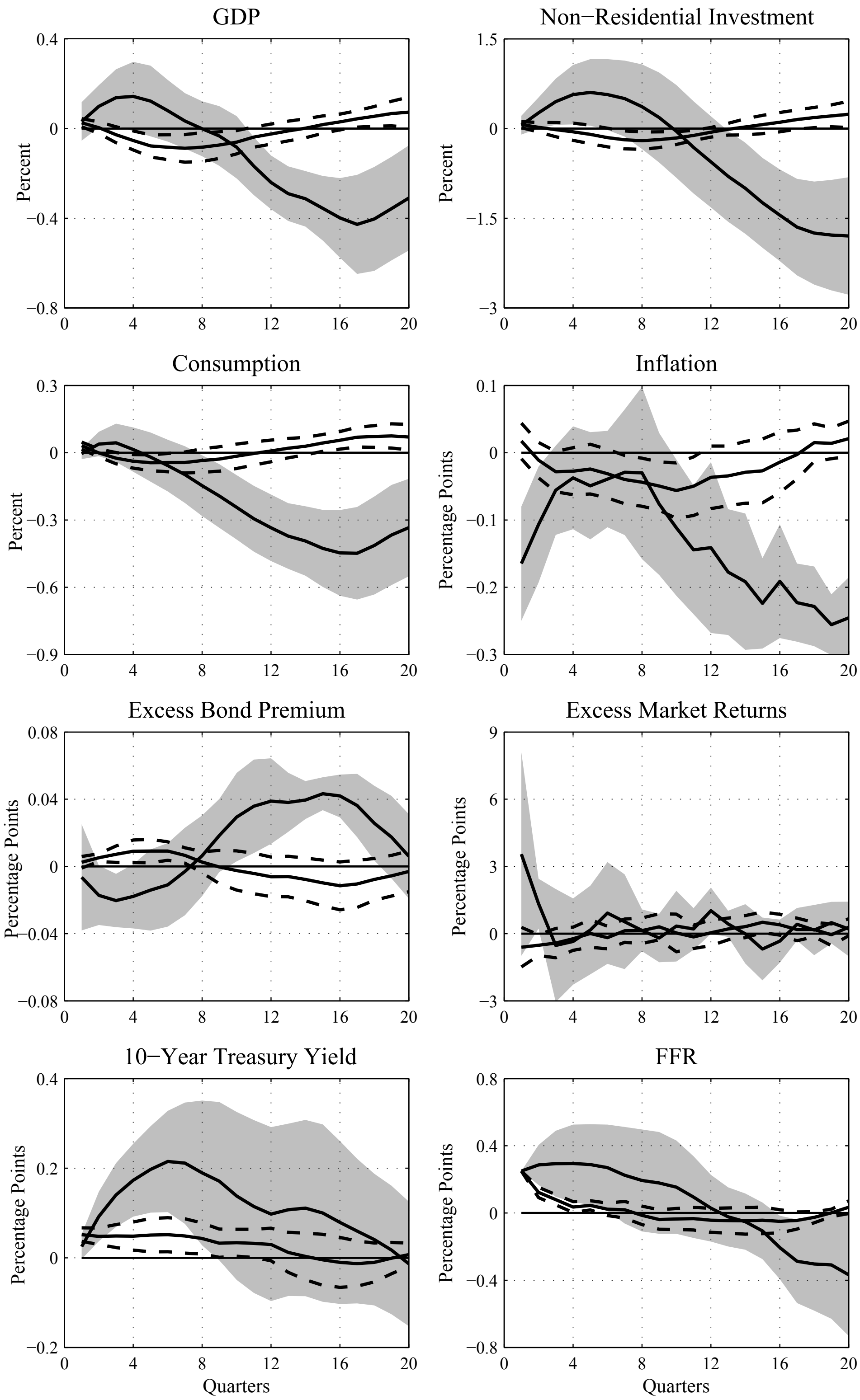


Notes: The figure plots $\Omega(z_t)$ and ε_t. The shortened time horizon reflects lost observations due to L and s from (4.1) thus constituting the effective sample, in which shocks occur. Shaded areas represent NBER-dated recessions.

Figure 4.3 traces out the adjustment patterns of the core variables after a contractionary 25 basis points monetary policy innovation together with Newey-West adjusted 95% confidence intervals. Monetary policy impacts GDP, non-residential investment, and consumption several times stronger and more persistently in a *high* (shaded error bands) relative to a *low* (dashed error bands) financial frictions environment (see, e.g., Dahlhaus, 2014). Statistically significant differences between regimes, primarily, emerge over the medium run. Price inflation also declines more in the high financial tensions state, whereas differences, in this case, are significant at short *and* medium horizons. These dynamics are paralleled by a muted, short-lived surge of the EBP in the low and a slowly building, pronounced increase of the EBP in the high financial frictions state.[7] In fact, the monetary tightening appears to amplify strains in financial markets, in particular, when frictions are *already high*, which is consistent with, inter alia, Bernanke and Gertler (1995), Bernanke et al. (1999), and subsequent financial accelerator models.

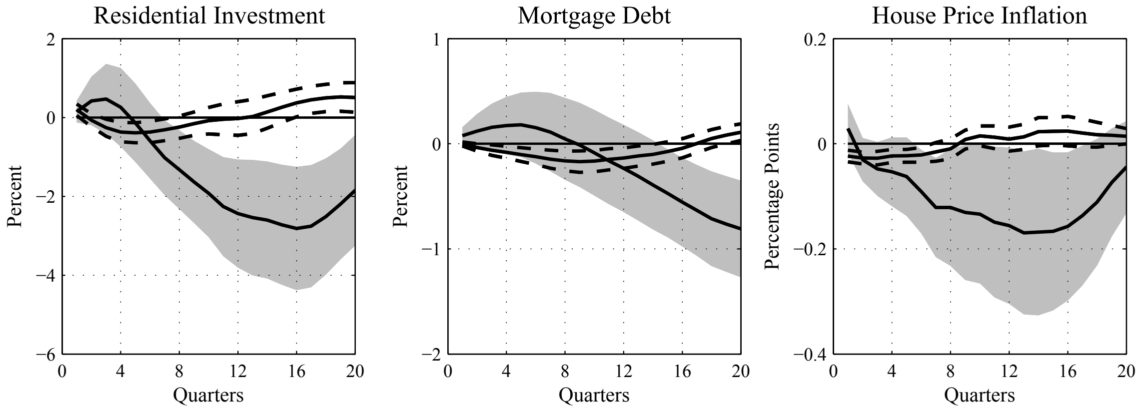
While the reaction of the policy instrument itself is somewhat more sluggish in the high tensions regime, differences between regimes are not significant for the majority of horizons, i.e., the systematic monetary policy component does not drive the results. Yet, we find some evidence for a stronger steepening of the yield curve in the high financial frictions state by documenting a more pronounced increase in 10-year sovereign bond yields, whereas the reaction of excess stock market returns is insignificant in both states. In addition, the main findings extend to the monetary policy propagation into housing markets (Figure 4.4), with housing market variables displaying excess sensitivity to monetary policy surprises in high relative to low financial frictions regimes.

[7]We also find an increase in the EBP, when estimating a state-invariant version of (4.1), with results presented in the Appendix of this chapter. This EBP reaction is in line with Gertler and Karadi (2015), who report a similar result using a (hybrid) high frequency identification approach in a linear VAR.



Notes: Impulse responses after a monetary policy shock in the low (dashed error bands) and high (shaded error bands) financial tensions regime.

Figure 4.4: Monetary policy propagation into housing markets



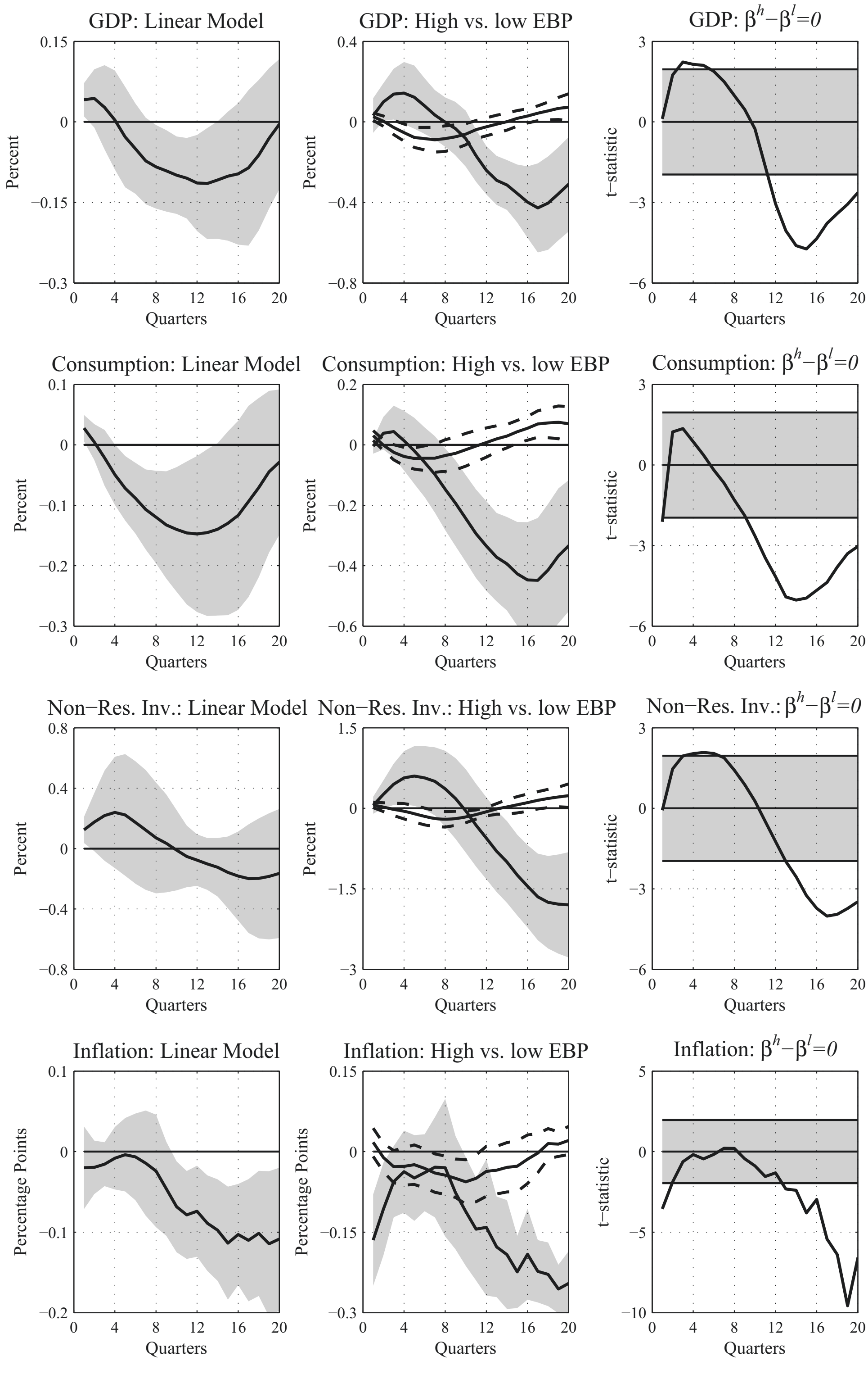
Notes: Impulse responses after a monetary policy shock in the low (dashed error bands) and high (shaded error bands) financial tensions regime.

4.4 Conclusion

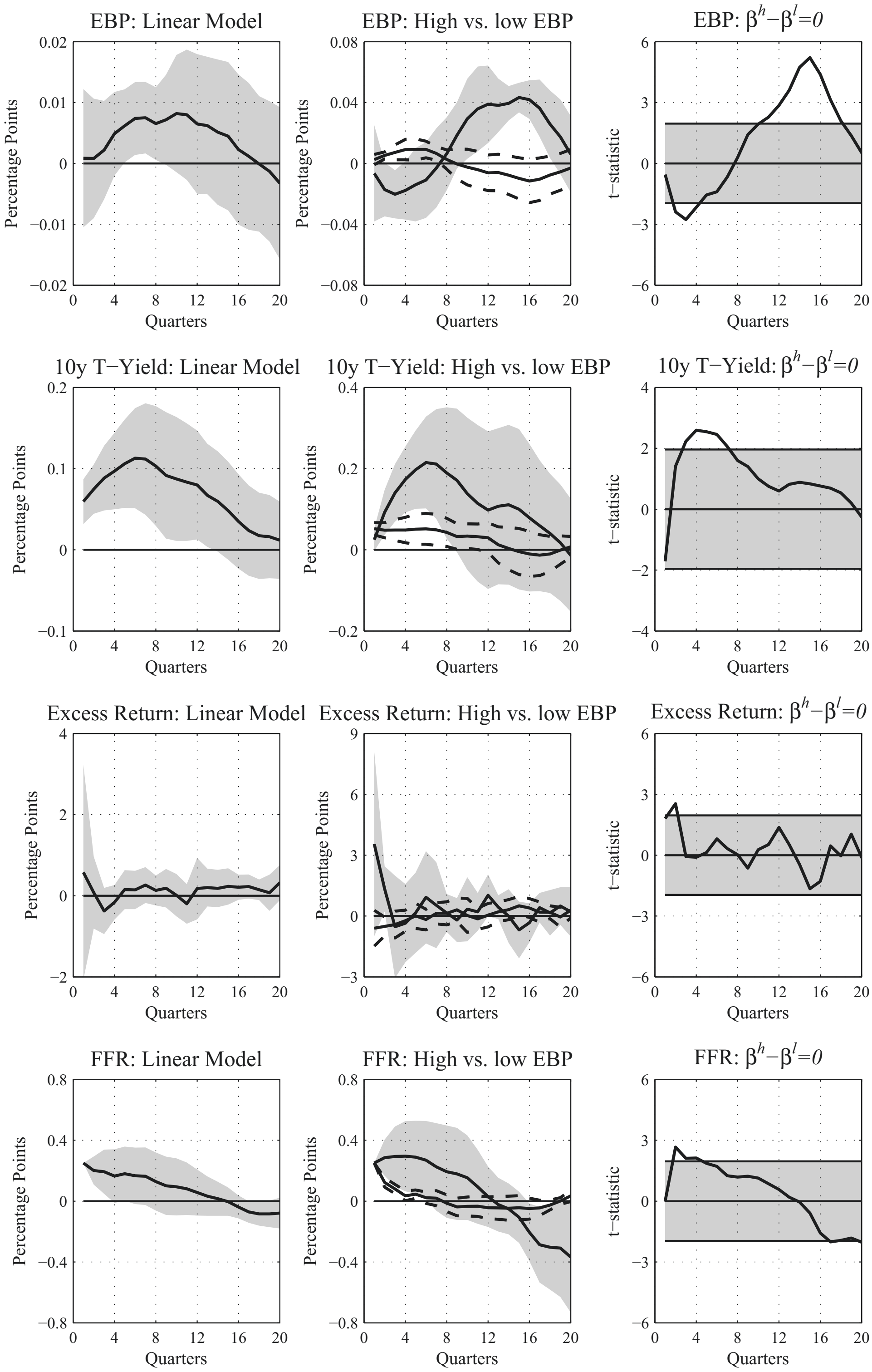
Based on local projections, we provide empirical evidence that US monetary policy affects the broader economy stronger and with more persistence, when strains in financial markets are high. A positive EBP reaction – signaling an increase in financial market frictions – hereby appears to amplify contractionary monetary policy shocks (e.g., Gertler and Karadi, 2015), in particular, in high financial tensions regimes.

Although the data sample in this analysis ends in 2008Q4, as conventional monetary policy reaction functions might not apply to the subsequent financial turmoil episode, our results have implications for the Great Recession, when stress in financial markets was severe. Indeed, with 200 basis points, the EBP reached levels during the financial crisis not seen before (Figure 4.1). Interpreted through the lens of this analysis, monetary policy might have been particularly effective in this episode. However, given the binding zero lower bound restriction on nominal interest rates, the FED's inability to reduce rates might – ceteris paribus – have exacerbated the Great Recession more than previous studies would predict.

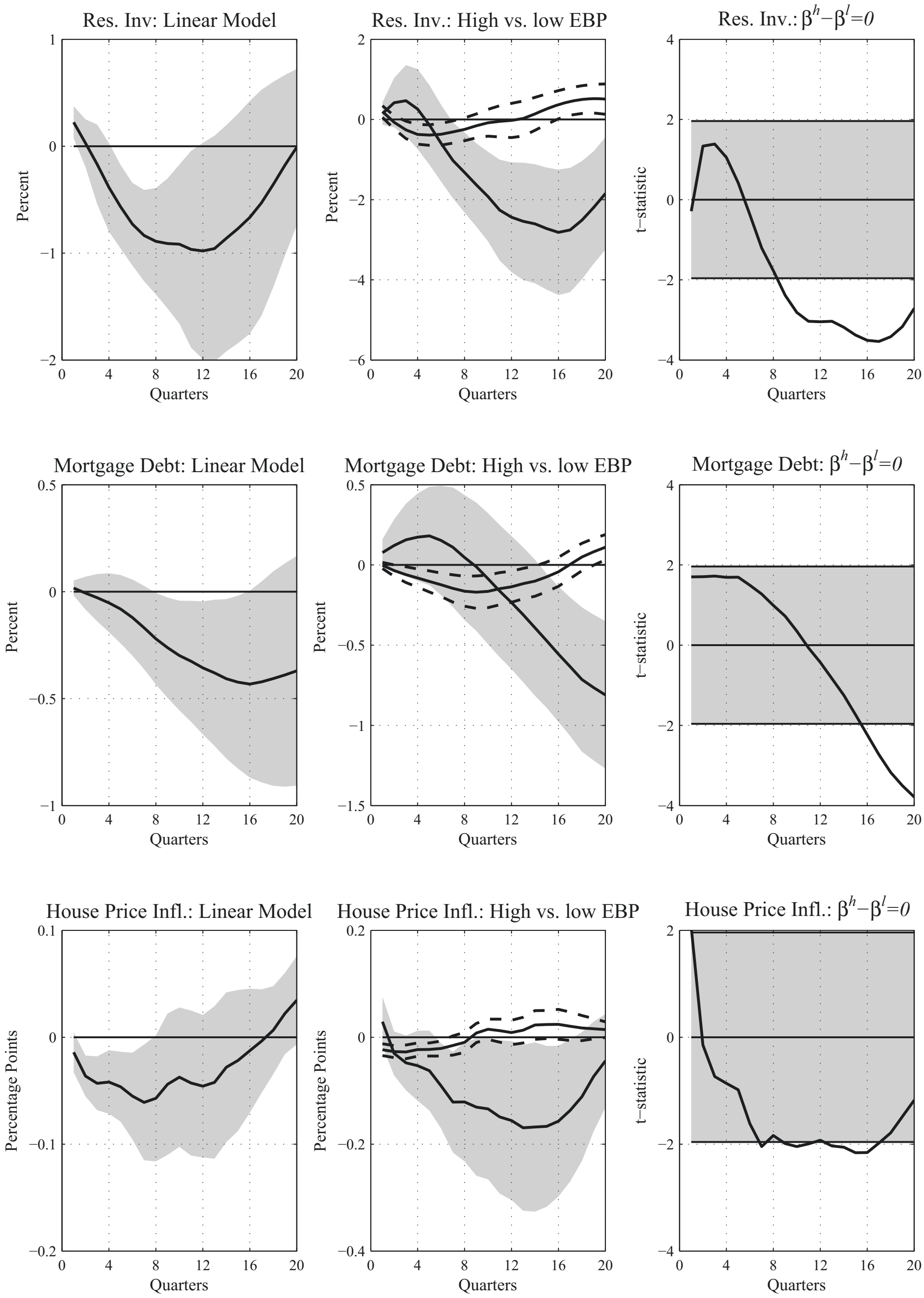
4.5 Appendix to chapter 4



Notes: Impulse responses after a monetary policy shock in the linear model (first column) and the low (dashed error bands) and high (shaded error bands) financial tensions models (second column), together with t-statistics from testing the null of no differences between state-dependent regimes (third column).



Notes: Impulse responses after a monetary policy shock in the linear model (first column) and the low (dashed error bands) and high (shaded error bands) financial tensions models (second column), together with t-statistics from testing the null of no differences between state-dependent regimes (third column).



Notes: Impulse responses after a monetary policy shock in the linear model (first column) and the low (dashed error bands) and high (shaded error bands) financial tensions models (second column), together with t-statistics from testing the null of no differences between state-dependent regimes (third column).

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