W. E. P.

Würzburg Economic Papers

No. 94

Current Account Dynamics and the Housing Boom and Bust Cycle in Spain

Daniel Maas, Eric Mayer, and
Sebastian Rüth[(a)]

[(a)]University of Würzburg

July 2015

University of Würzburg
Department of Economics
Chair for Monetary Poliy and
International Economics
Sanderring 2, D-97070 Würzburg

Corresponding Author

Eric Mayer
University of Würzburg, Department of Economics,
Sanderring 2, 97070 Würzburg, Germany
Email: eric.mayer@uni-wuerzburg.de

Current Account Dynamics and the Housing Boom and Bust Cycle in Spain

Daniel Maas, Eric Mayer, and Sebastian Rüth*

Abstract

We investigate the drivers of the negative correlation between housing markets and the current account in Spain. By employing robust sign restrictions, which we derive from a DSGE model for a currency union, we analyze the effects of domestic pull and foreign push factors in the mixed frequency VAR framework. Savings glut, risk premium, and house price expectations shocks are capable of generating the negative co-movement of housing markets and the current account in the data. In contrast, and counter-factual to the Spanish housing boom, financial easing shocks predict a decline in residential investment. Among the four identified shocks, savings glut shocks have most explanatory power for real house prices, residential investment, and the current account. We also reveal an important role of risk premium and house price expectations shocks for housing markets, whereas financial easing shocks do not explain sizeable fluctuations in the key variables.

Keywords: Current account, housing markets, monetary union.
JEL codes: E32, F32, F45.

*University of Wuerzburg, Department of Economics, Sanderring 2, 97070 Wuerzburg, Germany
Corresponding author: Eric Mayer, eric.mayer@uni-wuerzurg.de, +49 931-31 82948.

1 Introduction

What are the causal drivers of the well-established, negative correlation between house prices and the current account in Spain? Spain witnessed a pronounced boom and bust cycle in housing[1], which coincided with a severe deterioration and subsequent contraction of its current account (see Figure 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 (see Bonhomme and Hospido, 2012) that temporarily accounted for 20% of GDP growth (see Akin et al., 2014). At the peak of the bubble, the current account to GDP ratio recorded minus 10%, followed by a sharp correction after the bust.

The purpose of this paper is to trace back the origins of the joint house price and current account dynamics to domestic, i.e., Spain-specific shocks as well as to external shocks emerging in the rest of the Eurozone. The comparison of "pull" (domestic) versus "push" (foreign) factors, at least, dates back to Calvo et al. (1993) and is still subject to research on the sources of capital flows (see, e.g., Fratzscher, 2012). Locating these factors not only helps to reconstruct the Spanish boom and bust cycle, but is indispensable to design policies, which prevent comparable events in the future (see also Bernanke, 2005).

The pull hypothesis emphasizes the importance of domestic factors as a driver 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; Hristov et al., 2012) as well as house price expectations shocks that fuel markets against the backdrop in belief of ever surging house prices (see, e.g., Shiller, 2005, 2007; Gete, 2015).

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, e.g., 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 quantitatively 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 empirical relevance. We do so by applying a

[1] 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%, which is considerably beyond European average.

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), Iacoviello and Neri (2010), and Aspachs-Bracons and Rabanal (2011) 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 auto-regressive (VAR) model, which allows a discrimination of push and pull forces as in Sá and Wieladek (2015). Due to the short sample size of the Spanish housing cycle, 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 house price expectations 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 a weak drop in the current account and, most notably, a decline in residential investment. Moreover, the savings glut shock outperforms other shocks in terms of explaining variations in all key variables over all forecast horizons, with explained variance shares ranging from 25% to 33%. The risk premium shock, in particular, has explanatory power for residential investment and, to some extent, for the current account. The house price expectations shock explains residential investment in a similar dimension, whereas the financial easing shock fails to explain sizeable fluctuations of any key variable.

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 as there exists country heterogeneity along several dimensions.[2] Most importantly, Spain is member of EMU 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 bubble

[2]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).

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. Confirming their results, we find a prominent role of risk premium and house price expectations shocks for the housing market, in particular, for residential investment, where the former shocks explain the current account to a similar extent as in their analysis. We find little support for financial easing shocks in explaining movements in the Spanish housing market, which is in line with in't Veld et al. (2014). Third, in a recent paper, Gete (2015) applies sign restrictions to a sample of OECD countries to study the quantitative importance of housing demand factors. In addition to his paper, we derive sign restrictions for Spain by explicitly taking into account the specifics of a monetary union. We further extend his analysis by sheding light on the importance of capital push shocks via the open-economy VAR approach. Fourth, 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 like, e.g., Gete (2015) resort to data samples that extend the relevant time period for the same reason.[3] 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 paper is structured as follows. In Section 2, we explain the different hypotheses that we empirically test in detail. Section 3 discusses the model employed to derive the sign restrictions. Section 4 describes the econometric framework and presents the results. Section 5 concludes.

2 Competing Hypotheses

To motivate the quantitative analysis, we characterize the different sources that potentially explain the nexus of the housing boom and bust cycle and the current account in Spain more detailed.[4]

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% of the collateral value. In practice, banks placed 40% of all mortgage loans exactly on the limit of 80%. Furthermore, financial institutions owned appraisal firms, which systematically overstated property values (Akin et al., 2014). This gave scope for more credit, since mortgage volumes are conditioned on appraisals. As the fraction of collateral constrained households is sizeable in Spain (see, e.g., Hristov et al., 2014), these banking practices are of

[3]For Spain, the analysis by Gete (2015) starts in 1980 including two different currency regimes as well as distinct phases of the housing market.

[4]As argued in Shin (2012) and Acharya and Schnabl (2010), gross financial flows are more crucial for overall financing conditions than net financial flows as reflected by the current account. They propose a "banking" rather than a "savings" glut perspective. Yet Obstfeld (2012) forcefully emphasizes the importance of the current account for the scrutiny of policy makers (see Fratzscher et al., 2010; Sá and Wieladek, 2015). 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.

first order macroeconomic importance. The increased risk taking behavior of banks, probably induced by tough competition, effectively raised LTV ratios in terms of market values and softened lending standards before the crisis (see Figure 2). As a result, Spanish mortgage rates were 21% below European average. Other indicators of bad practice in the banking industry illustrate that credit conditions were often not fine-tuned to creditors. In particular, the employment status and other individual risk characteristics were not reflected by mortgage rates.[5] As mortgage growth 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.

A second prominent pull hypothesis are house price expectations shocks (see, e.g., in't Veld et al., 2014; Gete, 2015). 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), in't Veld et al. (2014), and Gete (2015), house price expectations shocks, moreover, cause current account deficits and thus capital inflows.[6] 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 house price expectations 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 non-durable goods. Accordingly, the domestic demand expansion induces imports of tradable goods, which causes current account deficits. Besides, house price expectations shocks can explain the coincidence of increasing house prices and strong residential investment, whereas LTV shocks fail along this dimension (see Gete, 2015). 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 house price expectations shocks to the broader economy is similar to a financial easing shock. However, policy implications of both pull disturbances are different.

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). 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 have even made investors belief in possible bail outs, decreasing the perception of political risks. Besides, as pointed out in Hale and Obstfeld (2014), the

[5]As explained by Fernández-Villaverde et al. (2013), savings banks (Cajas) were a driving force in financing the housing boom and temporarily accounted for roughly half of all mortgages. To some extent, these banks were governed by "low-human-capital managers" (Fernández-Villaverde et al., 2013), who added political concerns to the ingredients of the credit boost.

[6]Beyond, Cheng et al. (2014) and Ling et al. (2014) stress the importance of house price expectations shocks for a housing boom.

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 literally be 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.[7] Figure 4 depicts net saving rates for Spain and the Euro Area from 1999 to 2013. Since 2003, Spanish saving rates dropped from 7% to 0%, before sharply reverting at the onset of the Great Recession, while the Euro Area counterpart series fluctuated modestly around 8%. This setting is reminiscent to the idea of a savings glut as excessive savings in the core Eurozone were seeking for profitable investment opportunities in the periphery. Slack in core economies depressed Spanish exports, while the booming Spanish economy attracted imports and triggered current account deficits.

3 Robust 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).[8] 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.

3.1 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

[7]See also Adam et al. (2012) and Bofinger et al. (2013) for the interaction of real interest rate dynamics and belief in fueling house price booms.

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

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; Bernanke et al., 1999). Following Smets and Wouters (2003) and Christiano et al. (2005), the model considers several real and nominal frictions.

We discuss the model's endogenous response to exogenous shocks. We select shocks either to derive sign restrictions for the shocks in the empirical analysis or to ensure orthogonality to these disturbances. We restrict the presentation to the optimization problems of home country agents as there exists symmetry across the home country and the rest of the single currency area.

3.1.1 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

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

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, F, country. 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

$$P_{C,t}\tilde{C}_t(b) + P_{D,t}\tilde{X}_t(b) + R_{t-1}\tilde{S}_{t-1}(b) = \sum_{j}^{C,D} \frac{W_{j,t}}{\mathcal{M}_{j,t}} \tilde{L}_{j,t}(b) + \tilde{S}_t(b) + \Pi'_t(b). \quad (2)$$

$P_{j,t}, W_{j,t}$, 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, $\Pi'_t(b)$.

Indebted households borrow against the expected present value of their dwellings, which serve as collateral (see Iacoviello, 2005; Aspachs-Bracons and Rabanal, 2011). The nominal collateral constraint holds in every period and reads

$$R_t \tilde{S}_t(b) \leq \zeta_{LTV,t}(1-\chi)(1-\delta)\mathbb{E}_t \left\{ P_{D,t+1} \tilde{D}_t(b) \right\}, \tag{3}$$

where χ is the rate of down-payment, i.e., $1-\chi$ the LTV ratio, respectively (see, e.g., Calza et al., 2013). $\zeta_{LTV,t}$ 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)

$$\frac{R_t}{R_t^*} = \exp\left[-\kappa\left(b'_t - b'\right) + \zeta_{RP,t}\right]. \tag{4}$$

b'_t is the net foreign asset to nominal GDP ratio and b' the respective steady state. $\kappa \geq 0$ measures how sensitive the risk premium, R_t/R_t^*, reacts to fluctuations in b'_t, 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 (1) subject to (2), which gives

$$\tilde{U}_{C,t} = P_{C,t}\tilde{\lambda}_t \quad \text{and} \quad R_t^{-1} = \tilde{\beta}\mathbb{E}_t \left\{ \frac{P_{C,t}}{P_{C,t+1}} \frac{\tilde{U}_{C,t+1}}{\tilde{U}_{C,t}} \right\} + \tilde{\psi}_t. \tag{5}$$

$\tilde{U}_{C,t}$ 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

$$\frac{\tilde{U}_{D,t}}{\tilde{U}_{C,t}} = \frac{P_{D,t}}{P_{C,t}} - (1-\delta)\zeta_{PD,t}\left(\tilde{\psi}_t(1-\chi)\mathbb{E}_t\left\{\frac{P_{D,t+1}}{P_{C,t}}\right\} - \tilde{\beta}\mathbb{E}_t\left\{\frac{\tilde{U}_{C,t+1}}{\tilde{U}_{C,t}}\frac{P_{D,t+1}}{P_{C,t+1}}\right\}\right), \tag{6}$$

where $\tilde{U}_{D,t}$ denotes the marginal increase in utility from an extra unit of dwellings. $\zeta_{PD,t}$ is a stationary AR(1) shock to the expected future house price and affects the willingness to pay higher prices. 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., $P_{C,t} = (\tau P_{H,t}^{1-\iota} + (1-\tau){P_{F,t}}^{1-\iota})^{\frac{1}{1-\iota}}$.

3.1.2 Savers' Program

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

$$\mathbb{E}_0 \left\{ \sum_{t=0}^{\infty} \zeta_{\beta,t}\beta^t \left(\alpha \log(C_t(s) - hC_{t-1}) + (1-\alpha)\log D_t(s) - \frac{L_t(s)^{1+\eta}}{1+\eta} \right) \right\}, \tag{7}$$

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

$$\begin{aligned} & P_{C,t}C_t(s) + P_{D,t}X_t(s) + P_{I,t}\sum_{j}^{C,D} I_{j,t}(s) + S_t(s) + B_t(s) = \sum_{j}^{C,D} \frac{W_{j,t}}{\mathcal{M}_{j,t}} L_{j,t}(s) \\ & + \sum_{j}^{C,D} R_{j,t}Z_{j,t}(s)K_{j,t-1}(s) - P_{I,t}\sum_{j}^{C,D} a\left(Z_{j,t}(s)\right)K_{j,t-1}(s) + R_{l,t}l(s) + R_{t-1}S_{t-1}(s) \\ & + R_{t-1}B_{t-1}(s) + \Pi_t''(s) + \Pi_t'''(s). \end{aligned} \tag{8}$$

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.[9] 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

$$K_{j,t}(s) = (1-\delta_j)K_{j,t-1}(s) + \left[1 - S\left(\frac{I_{j,t}(s)}{I_{j,t-1}(s)}\right)\right] I_{j,t}(s), \tag{9}$$

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 $S''(1) = \rho > 0$.

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

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

$$U_{C,t} = P_{C,t}\lambda_t \quad \text{and} \quad R_t^{-1} = \beta \mathbb{E}_t \left\{ \frac{\lambda_{t+1}}{\lambda_t} \right\}. \tag{10}$$

Optimal consumption of the housing good implies

$$\frac{U_{D,t}}{U_{C,t}} = \frac{P_{D,t}}{P_{C,t}} - \beta(1-\delta)\mathbb{E}_t \left\{ \frac{U_{C,t+1}}{U_{C,t}} \frac{P_{D,t+1}}{P_{C,t+1}} \right\} \zeta_{PD,t}. \tag{11}$$

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

$$Q_{j,t} = \beta \mathbb{E}_t \left\{ \frac{U_{C,t+1}}{U_{C,t}} \left[(1-\delta_j) Q_{j,t+1} + \left(\frac{R_{j,t+1}}{P_{C,t+1}} Z_{j,t+1} - a(Z_{j,t+1}) \right) \right] \right\}, \tag{12}$$

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

$$1 - \beta \mathbb{E}_t \left\{ Q_{j,t+1} \frac{U_{C,t+1}}{U_{C,t}} S'\left(\frac{I_{j,t+1}}{I_{j,t}} \right) \left(\frac{I_{j,t+1}}{I_{j,t}} \right)^2 \right\}, \text{ and} \tag{13}$$

$$\frac{R_{j,t}}{P_{C,t}} = a'(Z_{j,t}), \tag{14}$$

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

3.1.3 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

$$\log\left(\frac{\omega_{j,t}}{\Pi_{C,t-1}^{\gamma_{W,j}}} \right) = \beta \mathbb{E}_t \left\{ \log\left(\frac{\omega_{j,t+1}}{\Pi_{C,t}^{\gamma_{W,j}}} \right) \right\} - \frac{(1-\theta_{W,j})(1-\beta\theta_{W,j})}{\theta_{W,j}} \log\left(\frac{\mathcal{M}_{j,t}}{\mathcal{M}_j} \right). \tag{15}$$

$\Pi_{C,t} = P_{C,t}/P_{C,t-1}$ 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

$$\mathcal{M}_{C,t} = \frac{W_{C,t}}{P_{C,t}} \frac{U_{C,t}}{(1-\varrho)^{-\iota_L} L_t^{\eta-\iota_L} L_{C,t}^{\iota_L}} \quad \text{and} \quad \mathcal{M}_{D,t} = \frac{W_{D,t}}{P_{C,t}} \frac{U_{C,t}}{\varrho^{-\iota_L} L_t^{\eta-\iota_L} L_{D,t}^{\iota_L}}. \tag{16}$$

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

$$\mathcal{M}_{C,t} = \frac{W_{C,t}}{P_{C,t}} \frac{\tilde{U}_{C,t}}{(1-\varrho)^{-\iota_L} \tilde{L}_t^{\eta-\iota_L} \tilde{L}_{C,t}^{\iota_L}} \quad \text{and} \quad \mathcal{M}_{D,t} = \frac{W_{D,t}}{P_{C,t}} \frac{\tilde{U}_{C,t}}{\varrho^{-\iota_L} \tilde{L}_t^{\eta-\iota_L} \tilde{L}_{D,t}^{\iota_L}}. \tag{17}$$

3.1.4 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

$$Y_{j,t} = \left(\frac{1}{n}\right)^{\lambda} \left(\int_0^n Y_{j,t}(i)^{\frac{1}{1+\lambda}} di\right)^{1+\lambda}. \tag{18}$$

$Y_{j,t}(i)$ 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

$$Y_{C,t}(i) = \frac{1}{n}\left(\frac{P_{H,t}}{P_{H,t}(i)}\right)^{\frac{1+\lambda}{\lambda}} Y_{C,t} \quad \text{and} \quad Y_{D,t}(i) = \frac{1}{n}\left(\frac{P_{D,t}}{P_{D,t}(i)}\right)^{\frac{1+\lambda}{\lambda}} Y_{D,t}. \tag{19}$$

$P_{j',t}(i)$ 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

$$P_{j',t} = \left(\frac{1}{n}\right)^{-\lambda} \left(\int_0^n P_{j',t}(i)^{-\frac{1}{\lambda}} di\right)^{-\lambda}. \tag{20}$$

3.1.5 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

$$Y_{C,t}(i) = K'_{C,t}(i)^{\mu_C} L_{C,t}(i)^{1-\mu_C},\ Y_{D,t}(i) = \zeta_{AD,t} l(i)^{\mu_l} K'_{D,t}(i)^{\mu_D} L_{D,t}(i)^{1-\mu_l-\mu_D}. \tag{21}$$

$K'_{j,t}(i) = Z_{j,t}(i) K_{j,t-1}(i)$ 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 the standard cost minimization problem, which results in the following sectoral marginal cost Equations

$$MC_{C,t}(i) = \frac{{R_{C,t}}^{\mu_C} {W_{C,t}}^{1-\mu_C}}{\mu_C^{\mu_C}(1-\mu_C)^{1-\mu_C}},\ MC_{D,t}(i) = \frac{{R_{l,t}}^{\mu_l} {R_{D,t}}^{\mu_D} {W_{D,t}}^{1-\mu_l-\mu_D}}{\mu_l^{\mu_l}\mu_D^{\mu_D}(1-\mu_l-\mu_D)^{1-\mu_l-\mu_D}\zeta_{AD,t}}. \quad (22)$$

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

$$R_{l,t} = \frac{\mu_l}{1-\mu_l-\mu_D}\frac{W_{D,t}L_{D,t}(i)}{l}, \quad (23)$$

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

$$\Pi_{C,t}(i) = (P_{H,t}(i) - MC_{C,t}(i))\left(\frac{1}{n}\right)\left(\frac{P_{H,t}(i)}{P_{H,t}}\right)^{-\frac{1+\lambda}{\lambda}} Y_{C,t} \text{ and} \quad (24)$$

$$\Pi_{D,t}(i) = (P_{D,t}(i) - MC_{D,t}(i))\left(\frac{1}{n}\right)\left(\frac{P_{D,t}(i)}{P_{D,t}}\right)^{-\frac{1+\lambda}{\lambda}} Y_{D,t}, \quad (25)$$

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

$$\mathbb{E}_t\left\{\sum_{k=0}^{\infty}\Lambda_{t,t+k}\theta_{P,C}Y_{C,t+k}(i)\left(\frac{\dot{P}_{H,t}(i)}{P_{H,t}}\frac{P_{H,t-1+k}^{\gamma_{P,C}}}{P_{H,t-1}^{\gamma_{P,C}}}\frac{P_{H,t}}{P_{H,t+k}} - (1+\lambda)\frac{MC_{C,t+k}(i)}{P_{H,t+k}}\right)\right\} = 0, \quad (26)$$

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

$$\mathbb{E}_t\left\{\sum_{k=0}^{\infty}\Lambda_{t,t+k}\theta_{P,D}Y_{D,t+k}(i)\left(\frac{\dot{P}_{D,t}(i)}{P_{D,t}}\frac{P_{D,t-1+k}^{\gamma_{P,D}}}{P_{D,t-1}^{\gamma_{P,D}}}\frac{P_{D,t}}{P_{D,t+k}} - (1+\lambda)\frac{MC_{D,t+k}(i)}{P_{D,t+k}}\right)\right\} = 0. \quad (27)$$

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

$${P_{H,t}}^{-\frac{1}{\lambda}} = \theta_{P,C}\left[P_{H,t-1}\left(\frac{P_{H,t-1}}{P_{H,t-2}}\right)^{\gamma_{P,C}}\right]^{-\frac{1}{\lambda}} + (1-\theta_{P,C})\dot{P}_{H,t}(i)^{-\frac{1}{\lambda}}, \quad (28)$$

and the housing sector

$$P_{D,t}{}^{-\frac{1}{\lambda}} = \theta_{P,D}\left[P_{D,t-1}\left(\frac{P_{D,t-1}}{P_{D,t-2}}\right)^{\gamma_{P,D}}\right]^{-\frac{1}{\lambda}} + (1-\theta_{P,D})\dot{P}_{D,t}(i)^{-\frac{1}{\lambda}}. \tag{29}$$

3.1.6 Market Equilibrium

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

$$\begin{aligned} Y_{C,t} = n &\left(\omega \tilde{C}_{H,t} + (1-\omega)\left(C_{H,t} + I^{C}_{H,t} + I^{D}_{H,t}\right)\right) \\ &+ (1-n)\left(\omega^{*}\tilde{C}^{*}_{H,t} + (1-\omega^{*})\left(C^{*}_{H,t} + I^{C*}_{H,t} + I^{D*}_{H,t}\right)\right) + \Omega_t, \end{aligned} \tag{30}$$

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

$$Y_{D,t} = n\left(\omega \tilde{X}_t + (1-\omega)X_t\right). \tag{31}$$

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 $j = C, D$. The equilibrium conditions of domestic and international debt markets are

$$\omega \tilde{S}_t = (1-\omega)S_t \quad \text{and} \quad n(1-\omega)B_t + (1-n)(1-\omega^{*})B^{*}_t = 0. \tag{32}$$

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

$$\begin{aligned} n(1-\omega)B_t = n(1-\omega)&R_{t-1}B_{t-1} \\ &+ (1-n)P_{H,t}\left[\omega^{*}\tilde{C}^{*}_{H,t} + (1-\omega^{*})\left(C^{*}_{H,t} + I^{C*}_{H,t} + I^{D*}_{H,t}\right)\right] \\ &- nP_{F,t}\left[\omega \tilde{C}_{F,t} + (1-\omega)\left(C_{F,t} + I^{C}_{F,t} + I^{D}_{F,t}\right)\right]. \end{aligned} \tag{33}$$

3.1.7 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

$$\frac{R^{*}_t}{R^{*}} = \left(\frac{R^{*}_{t-1}}{R^{*}}\right)^{\mu_R}\left(\frac{\Pi^{*}_t}{\Pi^{*}}\right)^{\mu_\Pi(1-\mu_R)}\left(\frac{Y^{*}_t}{Y^{*}_{t-1}}\right)^{\mu_{\Delta Y}}\left(\frac{\Pi^{*}_t}{\Pi^{*}_{t-1}}\right)^{\mu_{\Delta\Pi}} \exp\left(\epsilon^{*}_{R,t}\right). \tag{34}$$

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.

3.2 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 1).[10] Then we present median impulse responses together with 10% and 90% 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.

3.2.1 Exogenous Processes

We implement the four shocks from Section 2 in the DSGE model as follows.

- *Savings glut shock in the rest of the Eurozone.* This shock makes rest of union households 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 (1) and (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 (4).

- *Financial easing shock in Spain.* This shock enhances credit availability against housing collateral of domestic borrowers and equals a positive LTV shock, $\zeta_{LTV,t}$, in the collateral constraint Equation (3).

- *House price expectations shock in Spain.* As in Gete (2015), this is a shock to expected real house prices, $P_{D,t+1}/P_{C,t+1}$, i.e., an increase of $\zeta_{PD,t}$ in domestic borrowers' and savers' housing Euler Equations (6) and (11).

3.2.2 Calibration Strategy

For parameters governing nominal rigidities in goods and labor markets, we draw on the 90% posterior intervals of Smets and Wouters (2003). Calvo parameters, $\theta_{W,C}$ and $\theta_{P,C}$, range from 0.6 to 0.9.[11] 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 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

[10] We draw on empirical DSGE models like, e.g., Smets and Wouters (2003), Aspachs-Bracons and Rabanal (2011), in't Veld et al. (2014), Coenen et al. (2008), and Sá and Wieladek (2015) to specify parameter ranges.

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

Wouters, 2003; in't Veld et al., 2014). For the inverse Frisch elasticity, η, we allow for variations from 1.5 to 2.5 (Sá and Wieladek, 2015; 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%, and 10% in the non-durables sector. The capital share is 30% in the non-durables and 20% in the housing sector, while the land share is 10% 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% 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% 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 1), with standard deviations as in Aspachs-Bracons and Rabanal (2011).

3.2.3 Shock Propagation

Figure 5 displays a financial easing shock, where we calculate home country bond rates as a geometric average of short-term interest rates over a 10-year horizon (see Sá and Wieladek, 2015). 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. Thus imports 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 following the shock, savers' residential investment drops, which overcompensates borrowers' investment.

In contrast, a house price expectations shock can account for a co-movement of residential investment and real house prices (see Figure 6). While the ratio of consumption and residential investment increases following a financial easing shock, it decreases in the face of a house price expectations shock as in Gete (2015). We use this feature to disentangle the two shocks (see Figure 7).[12] 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 8 depicts the dynamics following a risk premium shock. Rest of union

[12] House price expectations shocks also encompass the idea of housing preference shocks. Among others, Aspachs-Bracons and Rabanal (2011) use the latter to simulate housing bubbles by directly shocking utility derived from housing services.

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 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. This calls upon the central bank to decrease the policy instrument. Hence, due to asymmetric business cycles, domestic interest rates are too low triggering a boom in this economy. In addition, lower interest rates decrease borrowers' cost of financial services and relax borrowing constraints. This financial accelerator effect supports domestic absorption and reinforces a deterioration of the home country's current account.

In summary, Table 2 displays the uniquely identifying set of robust sign restrictions, which we employ in the empirical analysis. Moreover, 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 (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 10). Thus a monetary policy shock is inconsistent with the restrictions in Table 2. Second, we study an increase in home country's housing sector-specific technology, $\zeta_{AD,t}$, in Equation (21). Again, all considered structural models robustly predict an increase in domestic and foreign consumption. Furthermore, and in contrast to the restrictions for the house price expectations shock, the housing technology shock predicts a negative correlation for the reponses of the current account and foreign consumption (see Figure 11).

4 Empirical Methodology

Now, we empirically analyze the effects of savings glut, risk premium, financial easing, and house price expectations 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.

4.1 Estimation, Data, and Inference

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

$$\boldsymbol{y}_t = \boldsymbol{c} + \sum_{l=1}^{p} \boldsymbol{\Phi}_l \boldsymbol{y}_{t-l} + \boldsymbol{\varepsilon}_t, \text{ where } \mathbb{E}[\boldsymbol{\varepsilon}_t] = 0 \text{ and } \mathbb{E}[\boldsymbol{\varepsilon}_t \boldsymbol{\varepsilon}_t'] = \boldsymbol{\Sigma}_\varepsilon. \tag{35}$$

$\boldsymbol{c}$ 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

$$\boldsymbol{y}_t = [\ \mathtt{C}_t\ \mathtt{C}_t^*\ \mathtt{CPI}_t\ \mathtt{CPI}_t^*\ \mathtt{BOND}_t\ \mathtt{BOND}_t^*\ \mathtt{EONIA}_t\ \mathtt{CA}_t\ \mathtt{RINV}_t\ \mathtt{CPIH}_t\ ]'. \tag{36}$$

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] $\mathtt{CPI}_t$ is the log level of the Harmonized Index of Consumer Prices (HICP). $\mathtt{C}_t$ denotes the $\mathtt{CPI}_t$ deflated log level of private consumption expenditures, and $\mathtt{BOND}_t$ 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 Euro Area expenditures.[14] $\mathtt{EONIA}_t$ represents interest rates in percent for unsecured, overnight lending in Euro Area interbank markets. As in Ciccarelli et al. (2015), we use $\mathtt{EONIA}_t$ 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 $\mathtt{EONIA}_t$ toward the ECB's deposit facility interest rate (see Lenza et al., 2010). Therefore, $\mathtt{EONIA}_t$, 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). $\mathtt{CA}_t$ stands for the Spanish current account to GDP ratio in percent. $\mathtt{RINV}_t$ and $\mathtt{CPIH}_t$ are log levels of real residential investment and a real house price index measuring residential property prices of all Spanish dwellings, respectively. Except for $\mathtt{CPIH}_t$, 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 and housing variables to study the effects of capital inflows on the Spanish housing market (see Sá and Wieladek, 2015, for a similar set of variables). 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.

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

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

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

information by relying on the coarsest data frequency for all variables. Thus we use 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

$$\boldsymbol{x}_t = [\ \texttt{CPI}_t\ \texttt{CPI}^*_t\ \texttt{BOND}_t\ \texttt{BOND}^*_t\ \texttt{EONIA}_t\ ]' \text{ and } \boldsymbol{z}_t = [\ \texttt{C}_t\ \texttt{C}^*_t\ \texttt{CA}_t\ \texttt{RINV}_t\ \texttt{CPIH}_t\ ]'. \quad (37)$$

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 $\widehat{z}^i_1, \widehat{z}^i_2, \widehat{z}^i_4 \dots \widehat{z}^i_{T-1}$. Furthermore, let $\widehat{\boldsymbol{z}}^i_{-t}$ represent the complete vector $\widehat{\boldsymbol{z}}^i$ except for element $\widehat{z}^i_t$. 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 $\widehat{z}^i_t$ from a multivariate normal density, while conditioning on $\boldsymbol{x}_t$, $\widehat{\boldsymbol{z}}^{i-1}_{-t}$, $\boldsymbol{c}^{i-1}$, $\boldsymbol{\Phi}^{i-1}_l$, and $\boldsymbol{\Sigma}^{i-1}_\varepsilon$. Second, we draw $\boldsymbol{c}^i$ and $\boldsymbol{\Phi}^i_l$ for given $\boldsymbol{x}_t$, $\widehat{\boldsymbol{z}}^i$, and $\boldsymbol{\Sigma}^{i-1}_\varepsilon$, and third, we obtain $\boldsymbol{\Sigma}^i_\varepsilon$ by conditioning on $\boldsymbol{x}_t$, $\widehat{\boldsymbol{z}}^i$, $\boldsymbol{c}^i$, and $\boldsymbol{\Phi}^i_l$. 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).[16]

4.2 Identification

From the VAR model in Equation (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 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 $\mathbb{E}[\boldsymbol{\varepsilon}_t\boldsymbol{\varepsilon}'_t] = \boldsymbol{\Sigma}_\varepsilon = \boldsymbol{PQQ}'\boldsymbol{P}'$, 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). Since impulse responses of these point-wise posterior statistics are not generated by the same structural model, we draw inference from the median target solution (Fry and Pagan, 2011). The median target refers to a single model producing impulse responses, which minimize the weighted distance to the median. Thus this model renders a structural interpretation.

[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 estimate the VAR with $p = 6$ lags, i.e., 2 quarters after linearly de-trending all series. However, results are robust to a more parsimonious specification with $p = 3$ lags, for instance.

We simultaneously identify four types of macroeconomic shocks by imposing sign restrictions as summarized in Table 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). 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.

4.3 Results

Figures 12 to 15 display the propagation of the identified shocks through the variables in $\boldsymbol{y}_t$. The shaded area denotes the 68% credible set from the Bayesian estimation, the solid line represents the median impulse response function, and the broken line depicts the median target solution. We report the dynamics of the system for 60 months, i.e., for five years as in the DSGE analysis. We define all monthly shocks to reduce consumption in the rest of the Euro Area, i.e., $\texttt{C}_t^*$ falls, as well as to incur a current account deficit in Spain, i.e., $\texttt{CA}_t$ declines.

After a savings glut shock, the current account remains significantly negative for twelve months, although we impose restrictions only for a horizon of nine months (see Figure 12). The unrestricted housing variables follow a hump-shaped increase, which for real house prices is significant (lasting for three years) and for residential investment slightly more persistent, however, only significant at the margin. Figure 13 displays adjustment patterns after a Spanish risk premium shock. This macroeconomic disturbance produces housing market and current account dynamics quantitatively similar to the savings glut shock. Though, it reveals more inertia with respect to house prices and the current account, which both stay significantly different from zero after five years. The financial easing shock from Figure 14 sluggishly forces the current account to a hump-shaped decline, which peaks after nine months before slowly decaying. Compared to the other shocks, the current account response is about half as pronounced. While real house prices feature a similar increase compared to the capital push shocks, residential investment, on the contrary, declines, albeit insignificantly. Nevertheless, our DSGE model predicts this negative reaction (see Figure 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 after a financial easing shock are counter-factual to the Spanish housing boom. This finding complements in't Veld et al. (2014), who provide empirical evidence for a short-term decline in Spanish residential investment in response to a firm-level LTV shock, whereas a household-level LTV shock predicts a rise in investment. As opposed to the financial easing shock, the house price expectations shock generates a negative correlation between the current account and all housing market variables in the VAR (see Figure 15). Most notably, residential investment slowly builds up for three years and remains statistically significant over a horizon of five years.

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 reveal the fractions of the forecast error variance, which can be attributed to the respective shocks over various forecast horizons in percent. To warrant a feasible structural interpretation, we present all k-step ahead forecast revisions for the median target model (Fry and Pagan, 2011). Clearly, the winner of the horse race regarding both housing market variables and the current account is the savings glut shock accounting for 25% to 33% of the variation in these variables. This reinforces the findings of Sá and Wieladek (2015) for the US, where savings glut shocks are quantitatively more suited to explain real house prices and residential investment than monetary policy shocks or housing demand factors. At short forecast horizons, the risk premium shock explains up to 17% of fluctuations in residential investment and up to 8% of current account variations, over the medium run. The financial easing shock, however, has little explanatory power for all key variables, which in't Veld et al. (2014) also report for the LTV shocks in an estimated DSGE model. Beyond, in a similar order of magnitude as the risk premium shock, we find a prominent role for the house price expectations shock in causing fluctuations of residential investment. This finding is in accordance with in't Veld et al. (2014) and Gete (2015). Yet the house price expectations shock in the former analysis has relatively more weight in explaining the current account, and the housing risk premium shock in the latter analysis has more power to explain real house prices, compared to our analysis. In general, the analysis leaves substantial fractions of forecast revisions undeclared, i.e., explained by structural shocks that we do not identify. However, we explain a total share of fluctuations, which exceeds explained variances in, e.g., Sá and Wieladek (2015), who employ a similar identification scheme.

5 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 build-up 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 quantitatively study this joint co-movement 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 shocks contributed substantially to the imbalances of Spain vis-à-vis the rest of the Eurozone as well as to the housing boom. To some extent, the latter phenomenon can also be traced back to risk premium and house price expectations shocks, whereas financial easing shocks play a minor role for housing markets and the current account. The analysis, however, is orthogonal to macroeconomic disturbances emerging from, e.g., asymmetric housing technology dynamics or the common monetary policy conduct. As a result, considerable variations in the data remain unexplained.

References

Acharya, V. V., Schnabl, P., 2010. Do global banks spread global imbalances? Asset-backed commercial paper during the financial crisis of 2007-09. IMF Economic Review 58 (1), 37–73.

Adam, K., Kuang, P., Marcet, A., 2012. House price booms and the current account. NBER Macroeconomics Annual 26 (1), 77 – 122.

Akin, O., Montalvo, J., Villar, J. G., Peydró, J.-L., Raya, J., 2014. The real estate and credit bubble: Evidence from Spain. SERIEs 5 (2), 223–243.

Aspachs-Bracons, O., Rabanal, P., 2011. The effects of housing prices and monetary policy in a currency union. International Journal of Central Banking 7 (1), 225–274.

Bernanke, B. S., 2005. The global saving glut and the U.S. current account deficit. Speech 77, Board of Governors of the Federal Reserve System (U.S.).

Bernanke, B. S., Gertler, M., Gilchrist, S., 1999. The financial accelerator in a quantitative business cycle framework. In: Taylor, J. B., Woodford, M. (Eds.), Handbook of Macroeconomics. Vol. 1 of Handbook of Macroeconomics. Elsevier, Ch. 21, pp. 1341–1393.

Bofinger, P., Debes, S., Gareis, J., Mayer, E., 2013. Monetary policy transmission in a model with animal spirits and house price booms and busts. Journal of Economic Dynamics and Control 37 (12), 2862–2881.

Bonhomme, S., Hospido, L., 2012. The cycle of earnings inequality: Evidence from Spanish social security data. IZA Discussion Papers 6669, Institute for the Study of Labor (IZA).

Calvo, G., Leiderman, L., Reinhart, C., 1993. Capital inflows and real exchange rate appreciation in latin america: The role of external factors. IMF Staff Papers Vol. 40, No. 1, 108-151.

Calvo, G. A., 1983. Staggered prices in a utility-maximizing framework. Journal of Monetary Economics 12 (3), 383–398.

Calza, A., Monacelli, T., Stracca, L., 2013. Housing finance and monetary policy. Journal of the European Economic Association 11, 101–122.

Canova, F., de Nicolo, G., 2002. Monetary disturbances matter for business fluctuations in the G-7. Journal of Monetary Economics 49 (6), 1131–1159.

Canova, F., Paustian, M., 2011. Business cycle measurement with some theory. Journal of Monetary Economics 58 (4), 345–361.

Catão, L. A., Milesi-Ferretti, G. M., 2014. External liabilities and crises. Journal of International Economics 94 (1), 18–32.

Cheng, I.-H., Raina, S., Xiong, W., 2014. Wall street and the housing bubble. American Economic Review 104 (9), 2797–2829.

Christiano, L. J., Eichenbaum, M., Evans, C. L., 2005. Nominal rigidities and the dynamic effects of a shock to monetary policy. Journal of Political Economy 113 (1), 1–45.

Christoffel, K., Coenen, G., Warne, A., 2008. The New Area-Wide Model of the Euro Area: A micro-founded open-economy model for forecasting and policy analysis. Working Paper Series 0944, European Central Bank.

Ciccarelli, M., Maddaloni, A., Peydro, J. L., 2015. Trusting the bankers: A new look at the credit channel of monetary policy. Review of Economic Dynamics, in press.

Coenen, G., McAdam, P., Straub, R., August 2008. Tax reform and labour-market performance in the Euro Area: A simulation-based analysis using the New Area-Wide Model. Journal of Economic Dynamics and Control 32 (8), 2543–2583.

Crespo-Cuaresma, J., Fernández-Amador, O., 2013. Business cycle convergence in EMU: A first look at the second moment. Journal of Macroeconomics 37 (C), 265–284.

Davis, M. A., Heathcote, J., 2005. Housing and the business cycle. International Economic Review 46 (3), 751–784.

Dees, S., di Mauro, F., Smith, L. V., Pesaran, M. H., 2007. Exploring the international linkages of the Euro Area: A global VAR analysis. Journal of Applied Econometrics 22 (1), 1–38.

Eraker, B., Chiu, C. W. J., Foerster, A. T., Kim, T. B., Seoane, H., 2014. Bayesian mixed frequency VARs. Journal of Financial Econometrics (forthcoming).

Erceg, C. J., Henderson, D. W., Levin, A. T., 2000. Optimal monetary policy with staggered wage and price contracts. Journal of Monetary Economics 46 (2), 281–313.

Faust, J., 1998. The robustness of identified VAR conclusions about money. Carnegie-Rochester Conference Series on Public Policy 49, 207–244.

Fernández-Villaverde, J., Garicano, L., Santos, T., 2013. Political credit cycles: The case of the Eurozone. Journal of Economic Perspectives 27 (3), 145–66.

Foroni, C., Marcellino, M., 2013. A survey of econometric methods for mixed-frequency data. Economics Working Papers ECO2013/02, European University Institute.

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

Fratzscher, M., Juvenal, L., Sarno, L., 2010. Asset prices, exchange rates and the current account. European Economic Review 54 (5), 643–658.

Fry, R., Pagan, A., 2011. Sign restrictions in structural vector autoregressions: A critical review. Journal of Economic Literature 49 (4), 938–60.

Gete, P., 2015. Housing markets and current account dynamics. Globalization and Monetary Policy Institute Working Paper 221, Federal Reserve Bank of Dallas.

Ghysels, E., 2015. Macroeconomics and the reality of mixed frequency data. Journal of Econometrics, forthcoming.

Giavazzi, F., Spaventa, L., 2011. Why the current account may matter in a monetary union: Lessons from the financial crisis in the Euro Area. The Euro Area and the Financial Crisis , in D. Cobham (ed.), Cambridge University Press.

Goddard, J., Molyneux, P., Wilson, J. O., Tavakoli, M., 2007. European banking: An overview. Journal of Banking & Finance 31 (7), 1911–1935.

Gonzalez, L., Ortega, F., 2013. Immigration and housing booms: Evidence from Spain. Journal of Regional Science 53 (1), 37–59.

Hale, G. B., Obstfeld, M., 2014. The Euro and the geography of international debt flows. CEPR Discussion Papers 9937, C.E.P.R. Discussion Papers.

Helbling, T., Huidrom, R., Kose, M. A., Otrok, C., 2011. Do credit shocks matter? A global perspective. European Economic Review 55 (3), 340–353.

Hristov, N., Hülsewig, O., Wollmershäuser, T., 2012. Loan supply shocks during the financial crisis: Evidence for the Euro Area. Journal of International Money and Finance 31 (3), 569–592.

Hristov, N., Hülsewig, O., Wollmershäuser, T., 2014. Financial frictions and inflation differentials in a monetary union. Manchester School 82 (5), 549–595.

Iacoviello, M., 2005. House prices, borrowing constraints, and monetary policy in the business cycle. American Economic Review 95 (3), 739–764.

Iacoviello, M., Neri, S., 2010. Housing market spillovers: Evidence from an estimated DSGE model. American Economic Journal: Macroeconomics 2 (2), 125–64.

in't Veld, J., Kollmann, R., Pataracchia, B., Ratto, M., Roeger, W., 2014. International capital flows and the boom-bust cycle in Spain. Journal of International Money and Finance 48 (PB), 314–335.

Justiniano, A., Primiceri, G., Tambalotti, A., 2014. The effects of the saving and banking glut on the U.S. economy. Journal of International Economics 92, S52–S67.

Justiniano, A., Primiceri, G., Tambalotti, A., 2015. Household leveraging and deleveraging. Review of Economic Dynamics 18 (1), 3–20.

Kiyotaki, N., Moore, J., 1997. Credit cycles. Journal of Political Economy 105 (2), 211–48.

Laibson, D., Mollerstrom, J., 2010. Capital flows, consumption booms and asset bubbles: A behavioural alternative to the savings glut hypothesis. Economic Journal 120 (544), 354–374.

Lenza, M., Pill, H., Reichlin, L., 2010. Monetary policy in exceptional times. Economic Policy 25, 295–339.

Ling, D., C., Ooi, J., Le, T., 2014. Explaining house price dynamics: Isolating the role of non-fundamentals. Journal of Money, Credit, and Banking, forthcoming.

Mayer, E., Gareis, J., 2013. What drives Ireland's housing market? A Bayesian DSGE approach. Open Economies Review 24 (5), 919–961.

Mendoza, E. G., Quadrini, V., Ríos-Rull, J.-V., 2009. Financial integration, financial development, and global imbalances. Journal of Political Economy 117 (3), 371–416.

Monacelli, T., 2009. New Keynesian models, durable goods, and collateral constraints. Journal of Monetary Economics 56 (2), 242–254.

Obstfeld, M., 2012. Financial flows, financial crises, and global imbalances. Journal of International Money and Finance 31 (3), 469–480.

Pariès, M. D., Notarpietro, A., 2008. Monetary policy and housing prices in an estimated DSGE model for the US and the Euro Area. Working Paper Series 972, European Central Bank.

Paustian, M., 2007. Assessing sign restrictions. The B.E. Journal of Macroeconomics 7 (1), 1–33.

Peersman, G., 2005. What caused the early millennium slowdown? Evidence based on vector autoregressions. Journal of Applied Econometrics 20 (2), 185–207.

Peersman, G., Straub, R., 2009. Technology shocks and robust sign restrictions in a Euro Area SVAR. International Economic Review 50 (3), 727–750.

Qian, H., 2013. Vector autoregression with mixed frequency data. Munich personal repec archive, No. 57856.

Quint, D., Rabanal, P., 2014. Monetary and macroprudential policy in an estimated DSGE model of the Euro Area. International Journal of Central Banking 10 (2), 169–236.

Rabanal, P., 2009. Inflation differentials between Spain and the EMU: A DSGE perspective. Journal of Money, Credit and Banking 41 (6), 1141–1166.

Sá, F., Wieladek, T., 2015. Capital inflows and the US housing boom. Journal of Money, Credit and Banking 47, 221–256.

Schmitt-Grohé, S., Uribe, M., 2003. Closing small open economy models. Journal of International Economics 61 (1), 163–185.

Shiller, R. J., 2005. Irrational exuberance. Priceton University Press.

Shiller, R. J., 2007. Understanding recent trends in house prices and home ownership. NBER Working Papers 13553, National Bureau of Economic Research, Inc.

Shin, H. S., 2012. Global banking glut and loan risk premium. IMF Economic Review 60 (2), 155–192.

Smets, F., Wouters, R., 2003. An estimated dynamic stochastic general equilibrium model of the Euro Area. Journal of the European Economic Association 1 (5), 1123–1175.

Smets, F., Wouters, R., 2007. Shocks and frictions in US business cycles: A Bayesian DSGE approach. American Economic Review 97 (3), 586–606.

Taylor, J. B., 1993. Discretion versus policy rules in practice. Carnegie-Rochester Conference Series on Public Policy 39 (1), 195–214.

Uhlig, H., 2005. What are the effects of monetary policy on output? Results from an agnostic identification procedure. Journal of Monetary Economics 52 (2), 381–419.

Tables

Table 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, 0.25]
$\mu_{\Delta\Pi}$	Reaction coefficient: change in inflation	[0.00, 0.25]
ρ_{PD}	Persistence: house price expectations shock	[0.95, 0.99]
ρ_{LTV}	Persistence: LTV 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.

Table 2: Sign Restrictions

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

Notes: We impose restrictions for three quarters, i.e., 9 months as ≤ 0 or ≥ 0 (see, e.g., Sá and Wieladek, 2015).

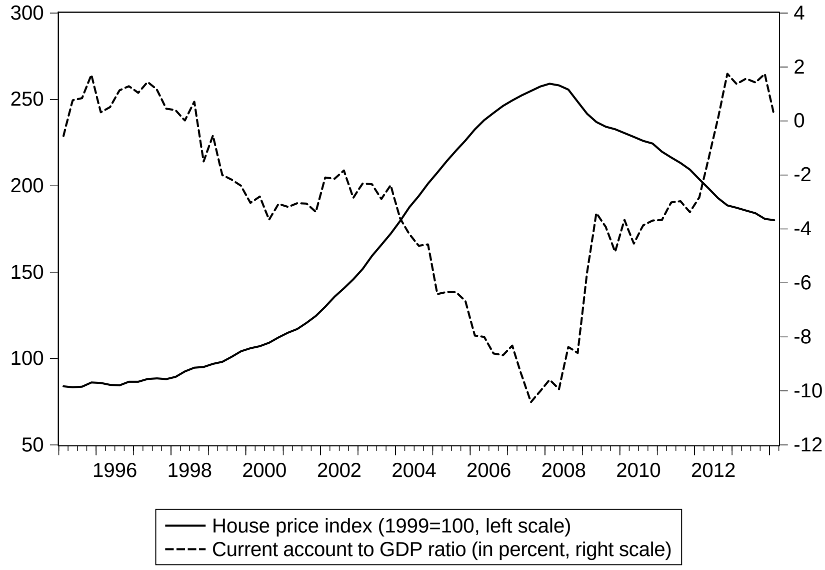
Table 3: Forecast Error Variance Decomposition

	Horizon	Current Account	House Prices	Res. Investment
Savings Glut Shock	6 Months	24.99	26.57	31.93
	12 Months	26.06	26.64	32.76
	18 Months	27.26	26.80	33.10
	24 Months	28.28	27.01	33.21
	60 Months	31.08	28.13	33.38
Risk Premium Shock	6 Months	4.17	1.71	16.80
	12 Months	4.80	2.26	15.59
	18 Months	5.88	2.71	14.71
	24 Months	6.54	3.08	14.04
	60 Months	7.75	4.17	12.69
Financial Easing Shock	6 Months	3.09	1.60	1.28
	12 Months	0.87	1.27	1.17
	18 Months	0.50	1.13	1.05
	24 Months	0.39	1.04	0.95
	60 Months	0.20	0.83	0.76
House Price Expectations Shock	6 Months	1.15	2.51	16.72
	12 Months	1.40	2.61	16.53
	18 Months	1.95	2.70	15.87
	24 Months	2.35	2.77	15.09
	60 Months	3.32	3.01	12.92

Notes: Results are in percent for the median target model (see Fry and Pagan, 2011).

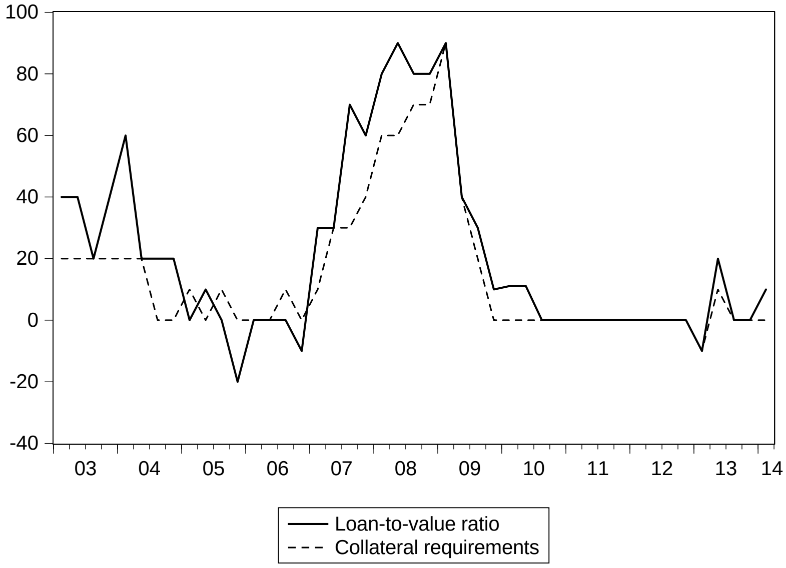
Figures

Figure 1: Current Account and House Price Dynamics



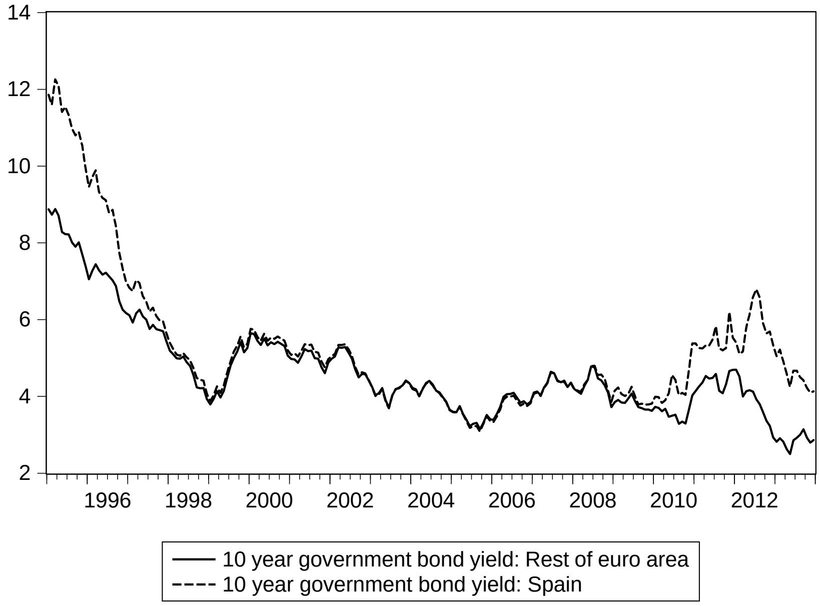
Notes: The figure presents the development of the current account to GDP ratio and house prices for Spain. We obtain the data from Eurostat and BIS.

Figure 2: Changes in Spanish Banks' Lending Standards



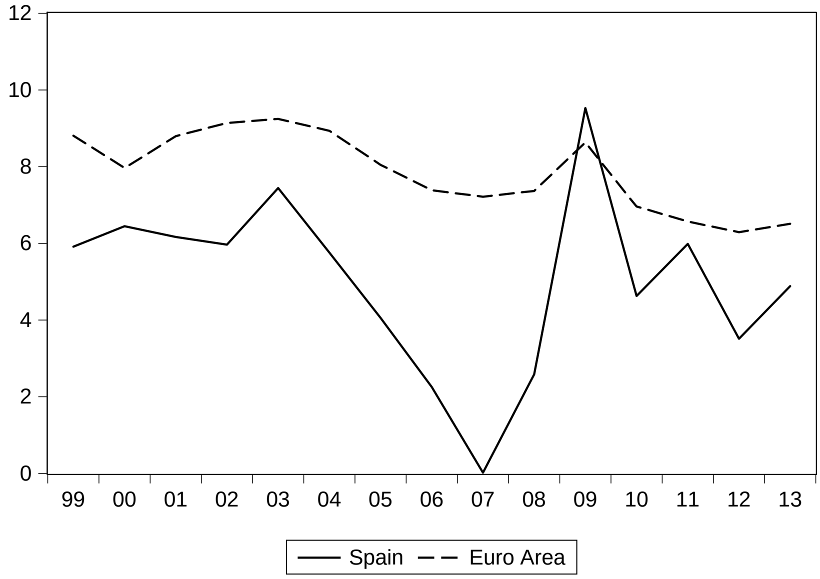
Notes: The figure shows the change in banks' conditions for housing loans to households over the past three months (frequency of tightened minus eased lending standards). We obtain the data from the ECB's bank lending survey, which is available since 2003.

Figure 3: 10-Year Government Bond Yields



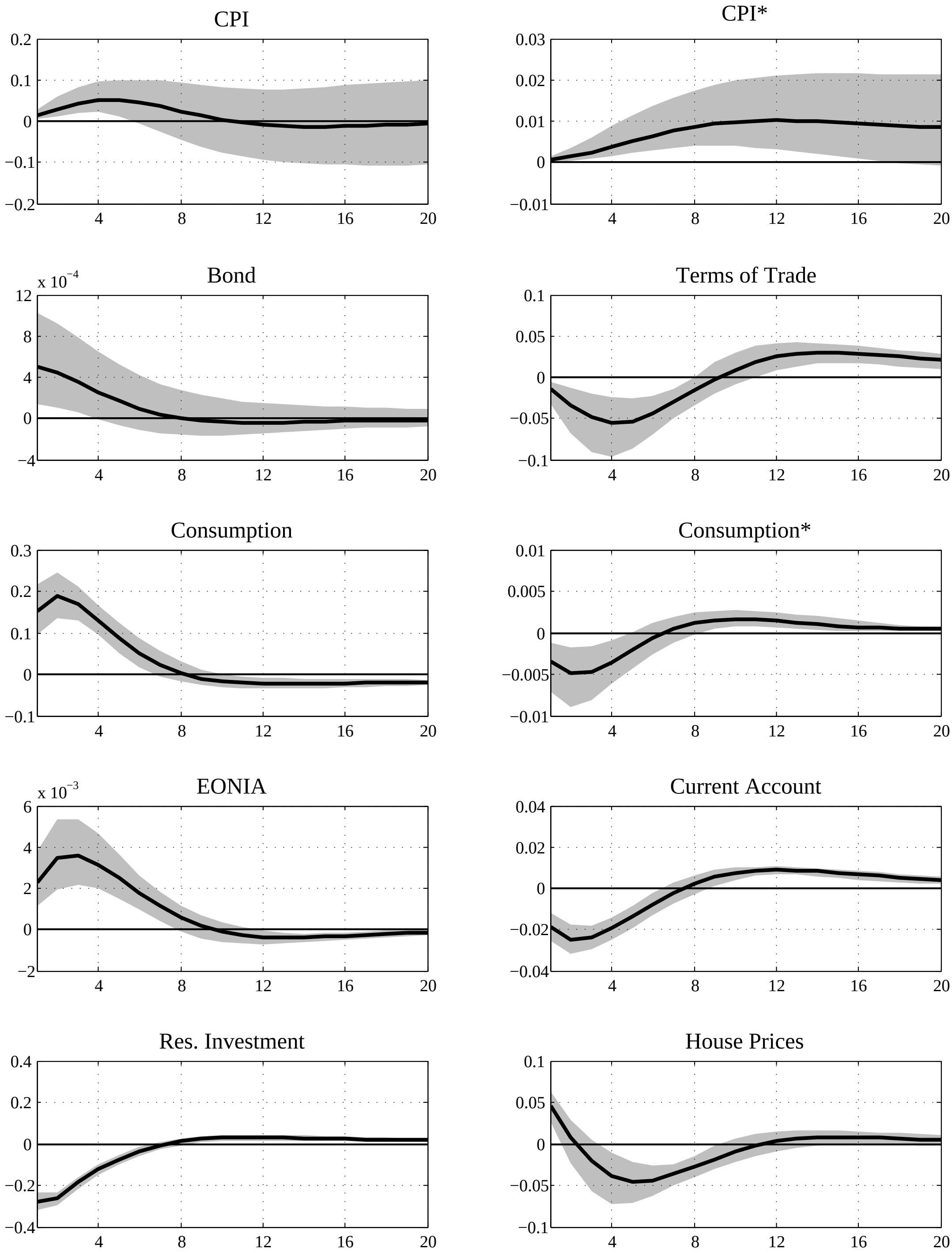
Notes: The figure depicts the development of 10-year government bond yields for Spain and the rest of the Euro Area. We obtain the data from Eurostat.

Figure 4: Net Household Saving as Percentage of Net Disposable Income



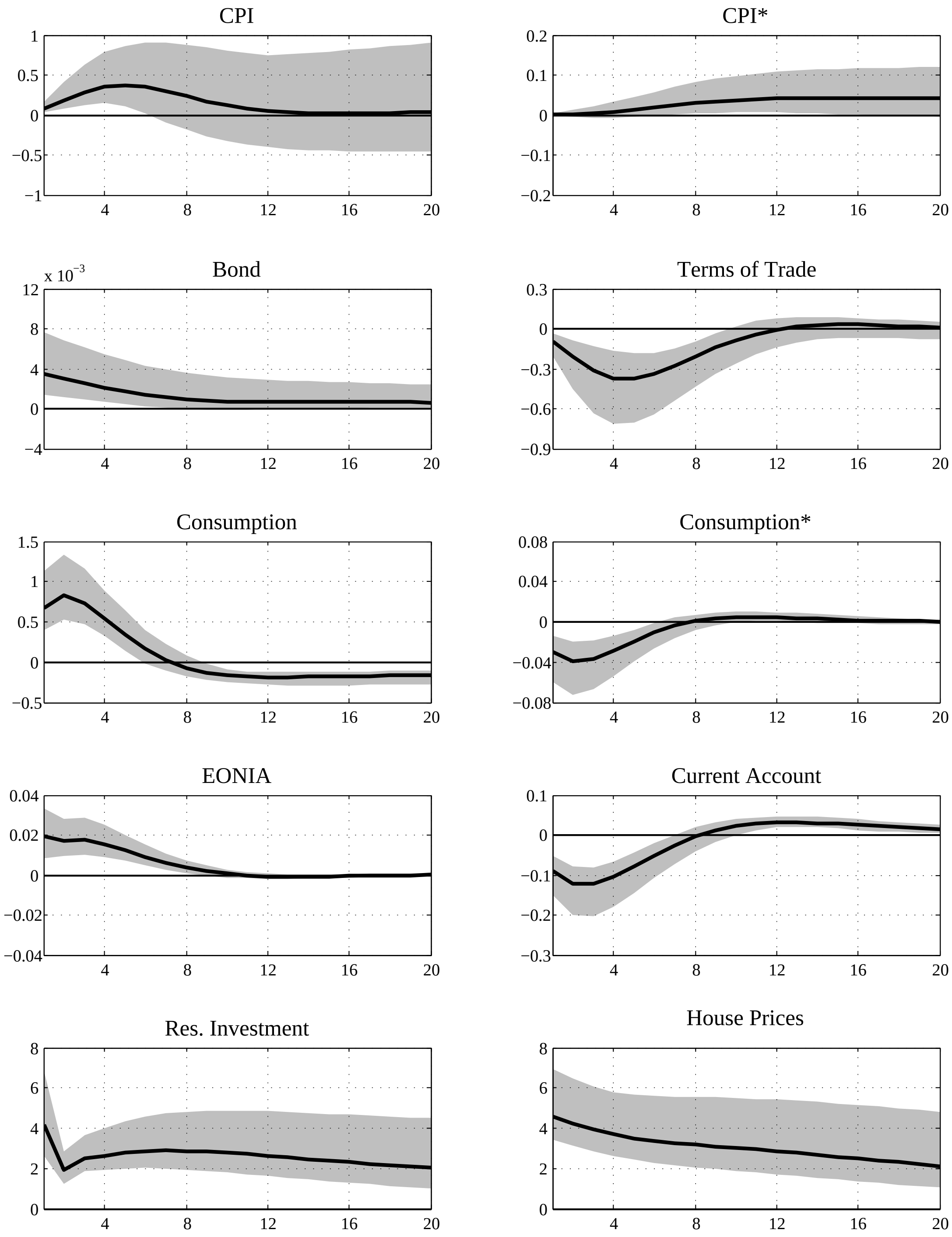
Notes: The figure portrays net household saving as a percentage of net disposable income for Spain and the rest of the Euro Area. We obtain the data from the OECD.

Figure 5: Financial Easing Shock



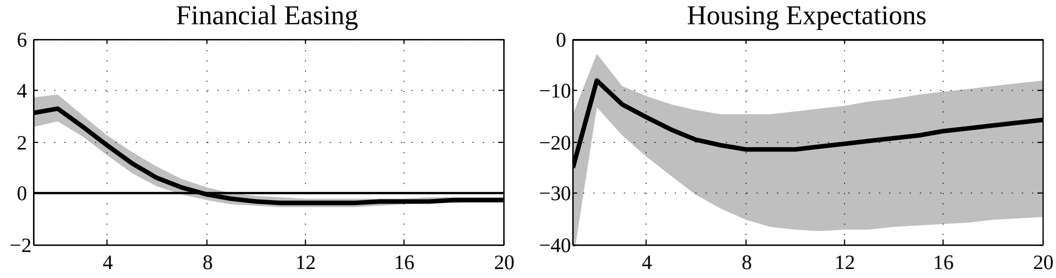
Notes: The x-axis is in quarters. The y-axis measures percent deviations from steady state. The solid line represents the median impulse response. Shaded areas display 10% and 90% percentiles of the simulated impulse responses.

Figure 6: Housing Expectations Shock



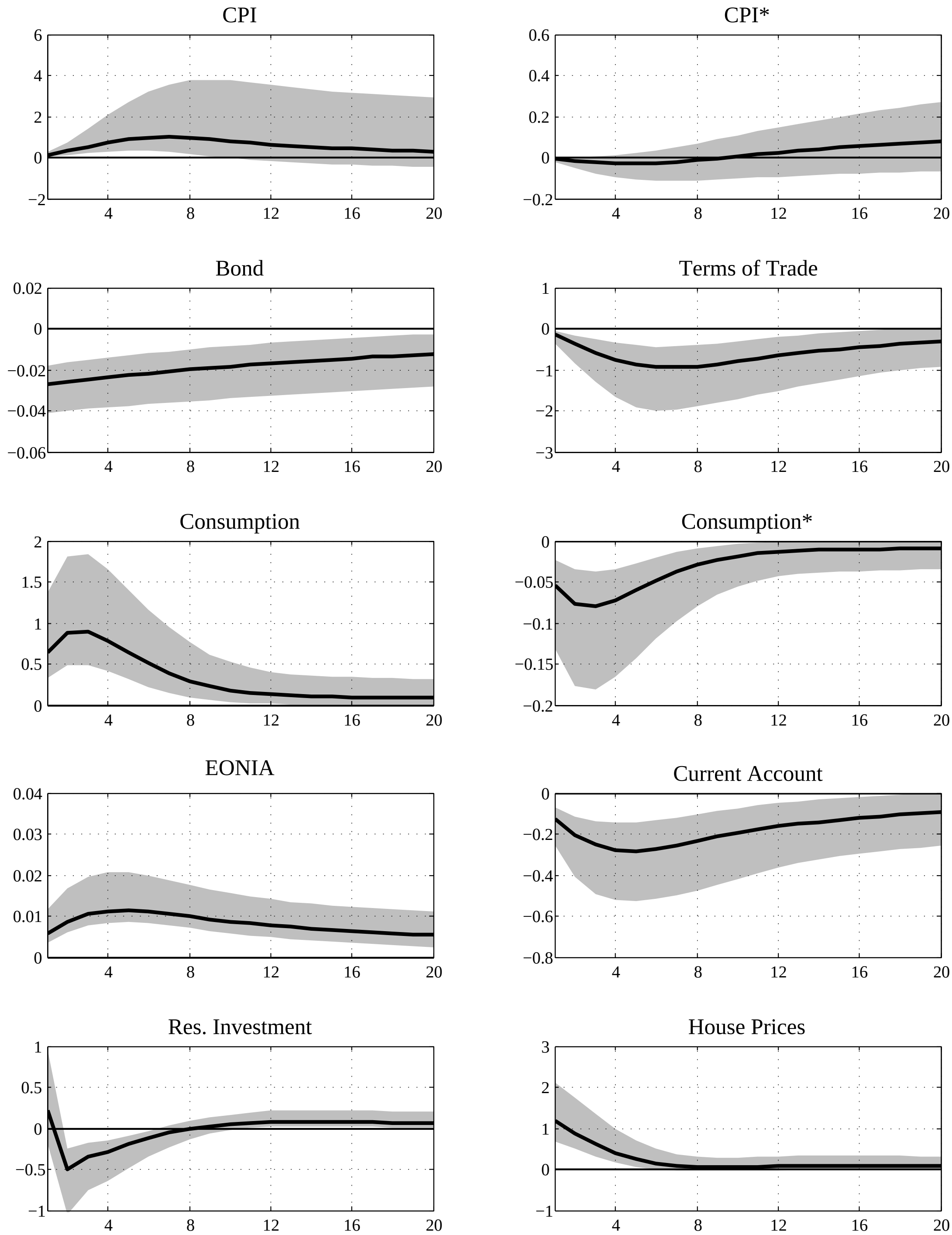
Notes: The x-axis is in quarters. The y-axis measures percent deviations from steady state. The solid line represents the median impulse response. Shaded areas display 10% and 90% percentiles of the simulated impulse responses.

Figure 7: Consumption to Residential Investment Ratio



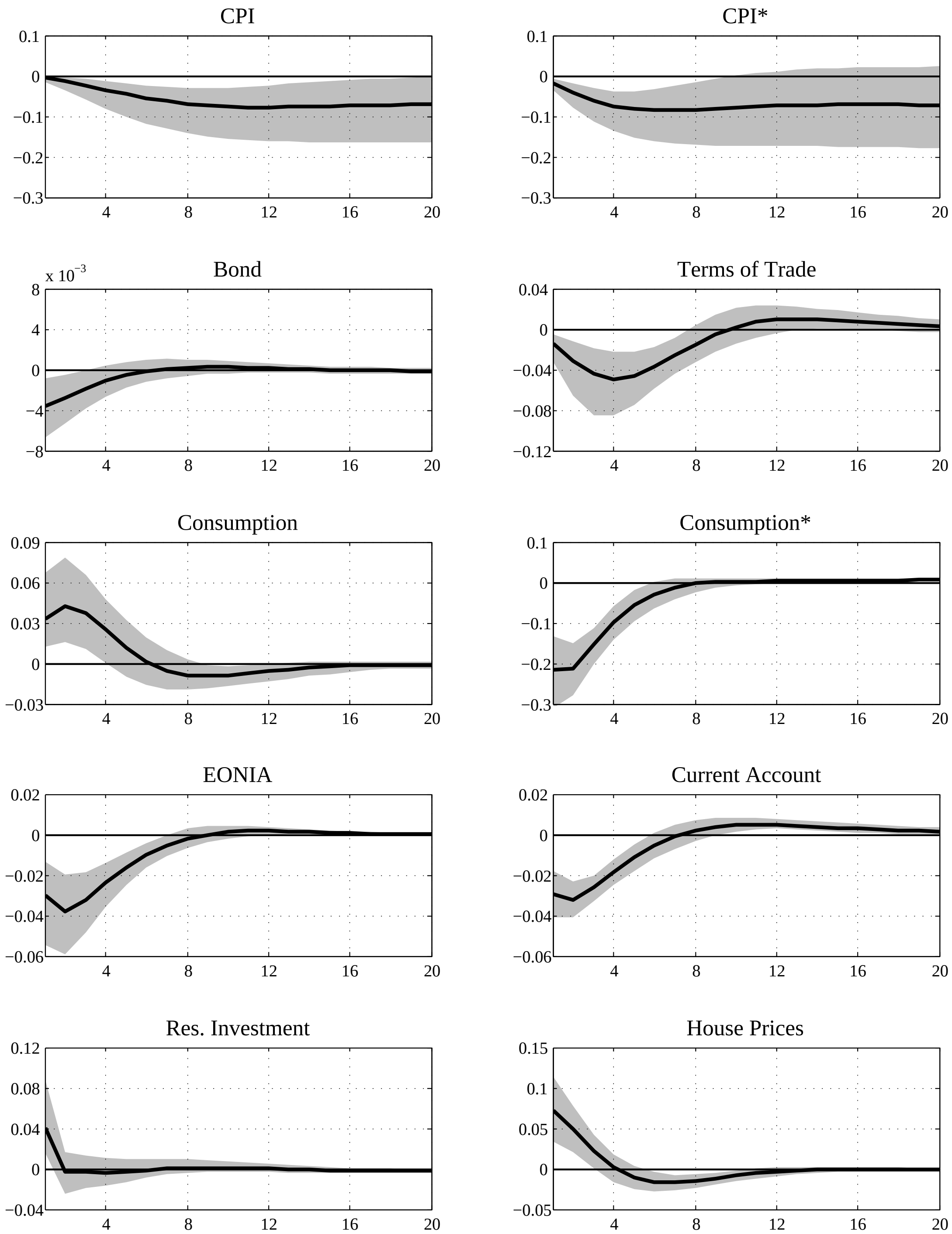
Notes: The x-axis is in quarters. The y-axis measures percent deviations from steady state. The solid line represents the median impulse response. Shaded areas display 10% and 90% percentiles of the simulated impulse responses.

Figure 8: Risk Premium Shock



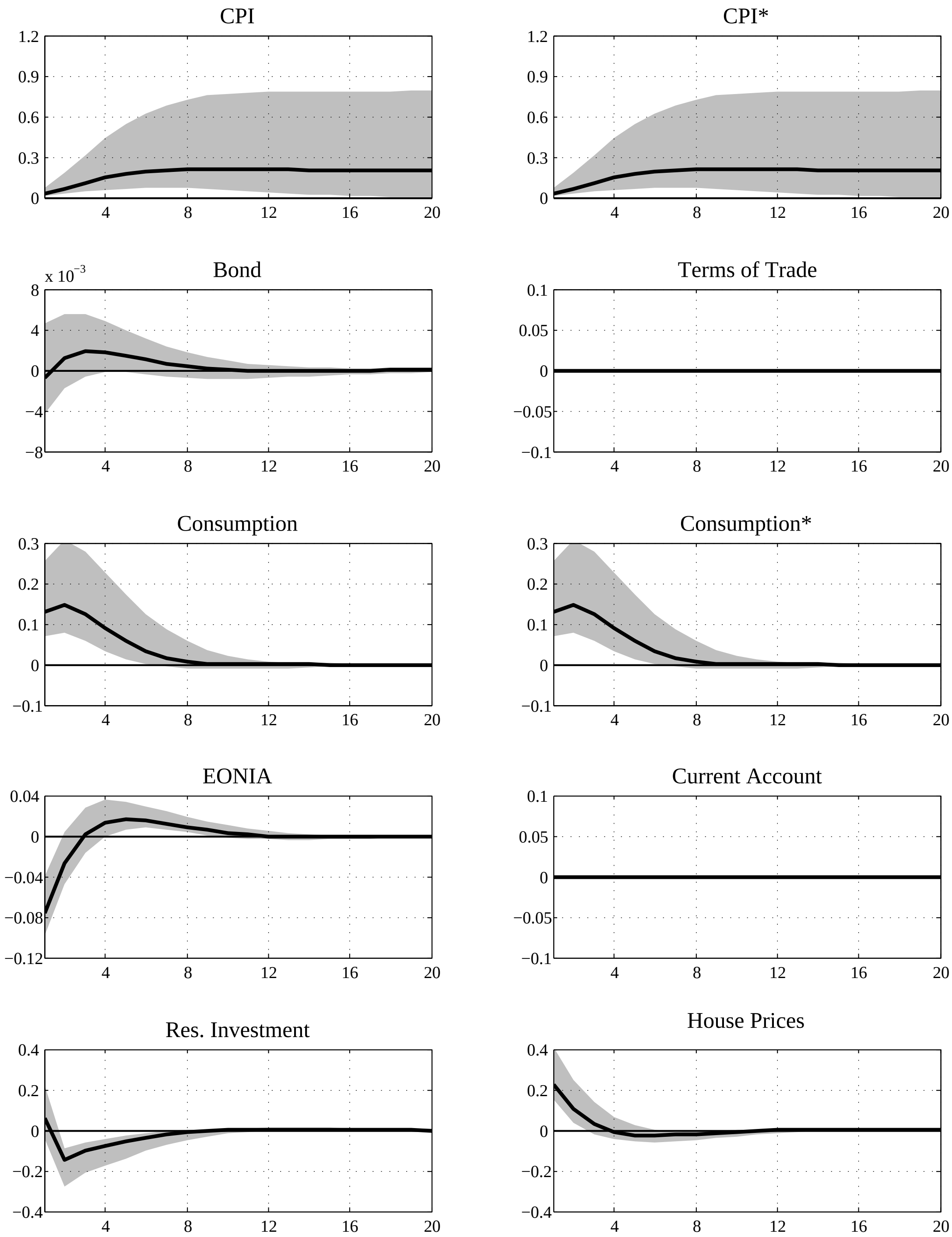
Notes: The x-axis is in quarters. The y-axis measures percent deviations from steady state. The solid line represents the median impulse response. Shaded areas display 10% and 90% percentiles of the simulated impulse responses.

Figure 9: Savings Glut Shock



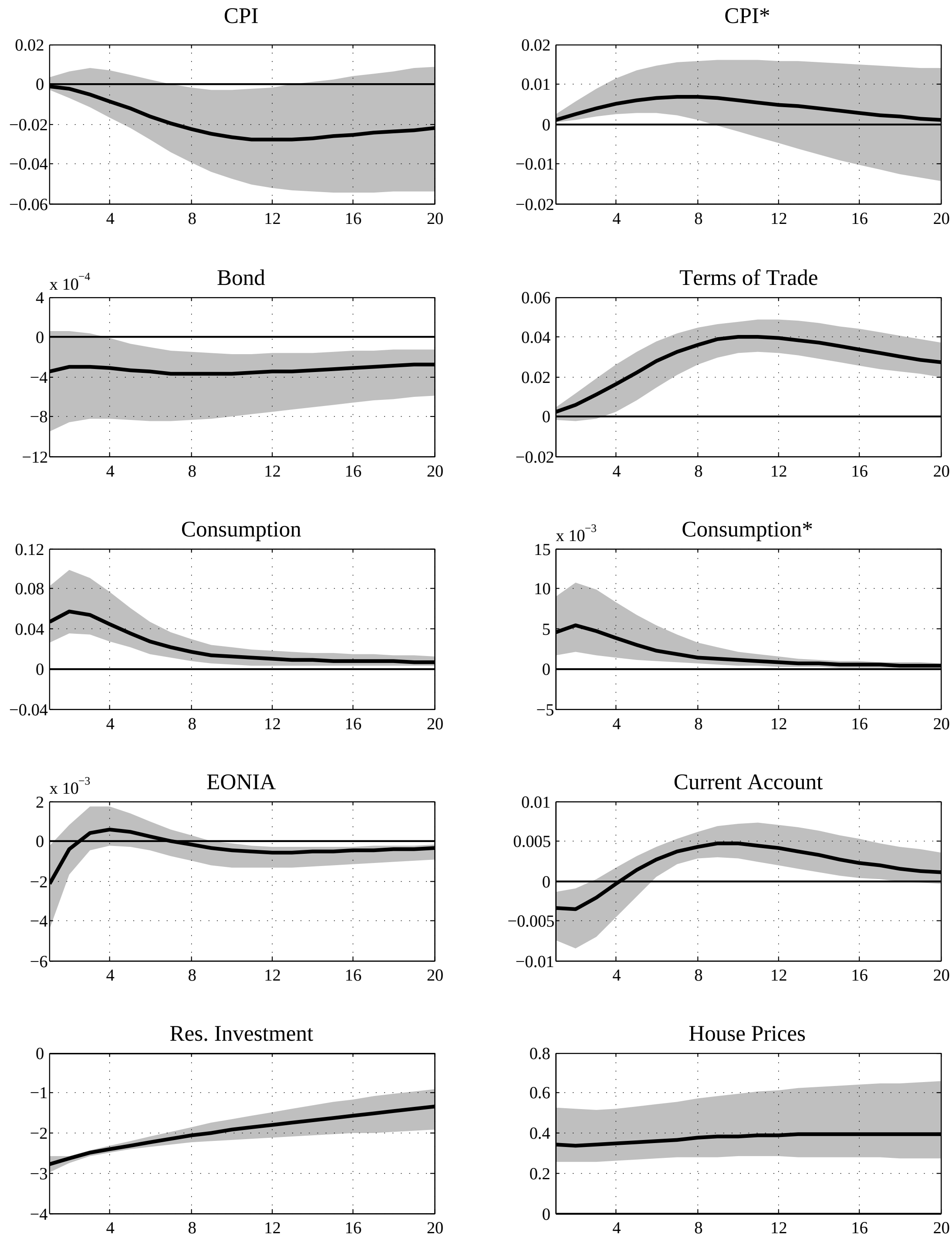
Notes: The x-axis is in quarters. The y-axis measures percent deviations from steady state. The solid line represents the median impulse response. Shaded areas display 10% and 90% percentiles of the simulated impulse responses.

Figure 10: Monetary Policy Shock



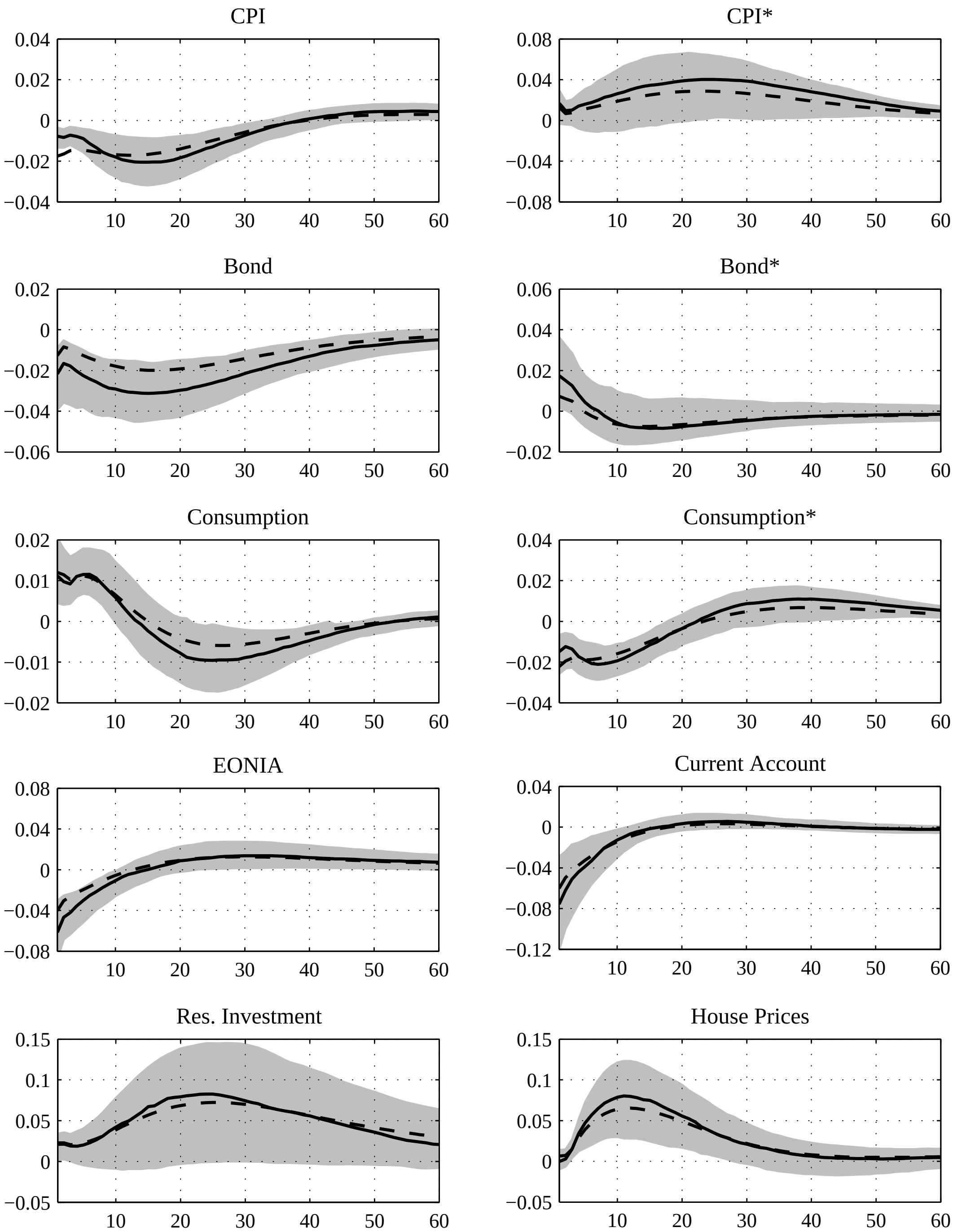
Notes: The x-axis is in quarters. The y-axis measures percent deviations from steady state. The solid line represents the median impulse response. Shaded areas display 10% and 90% percentiles of the simulated impulse responses.

Figure 11: Housing Technology Shock



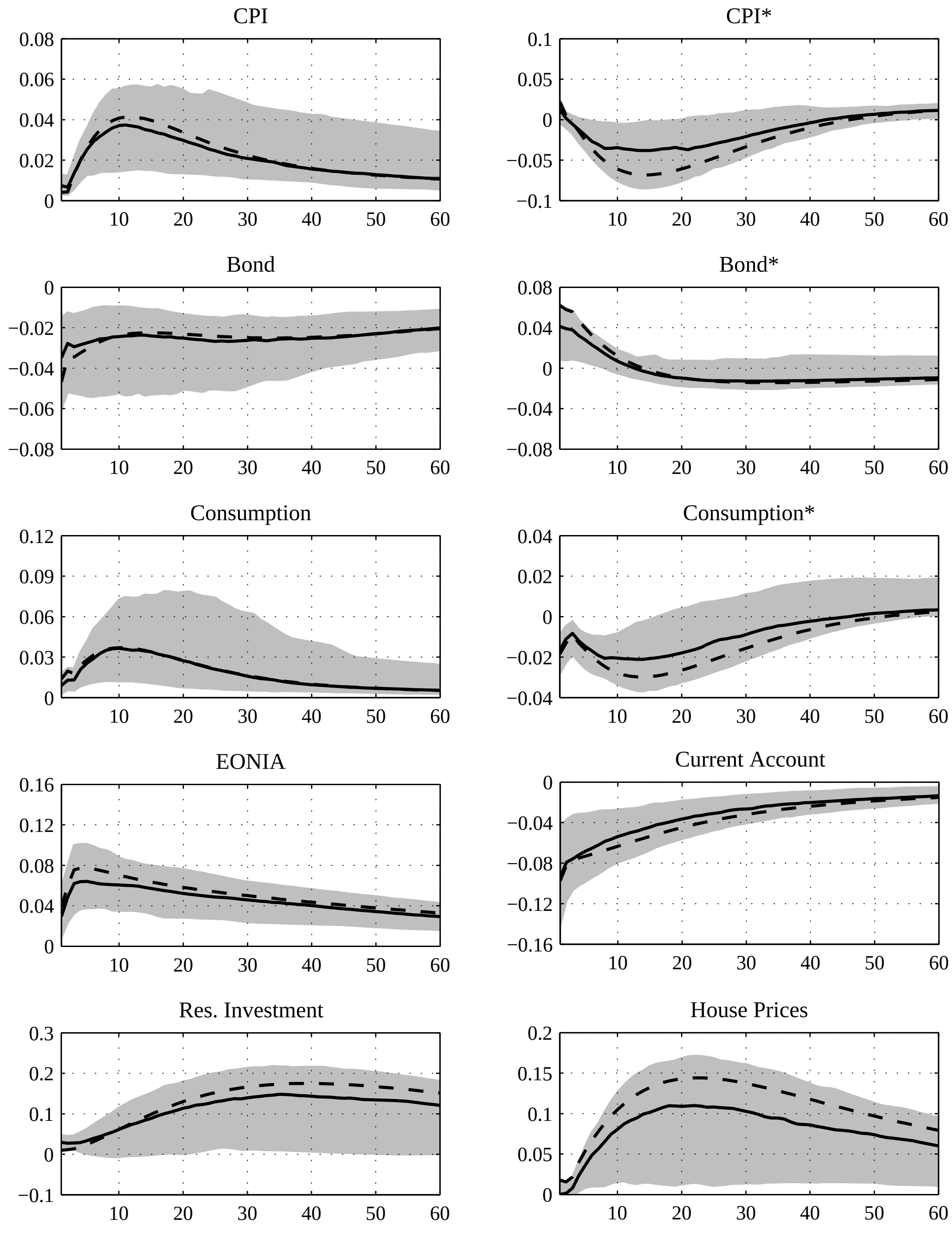
Notes: The x-axis is in quarters. The y-axis measures percent deviations from steady state. The solid line represents the median impulse response. Shaded areas display 10% and 90% percentiles of the simulated impulse responses.

Figure 12: Savings Glut Shock



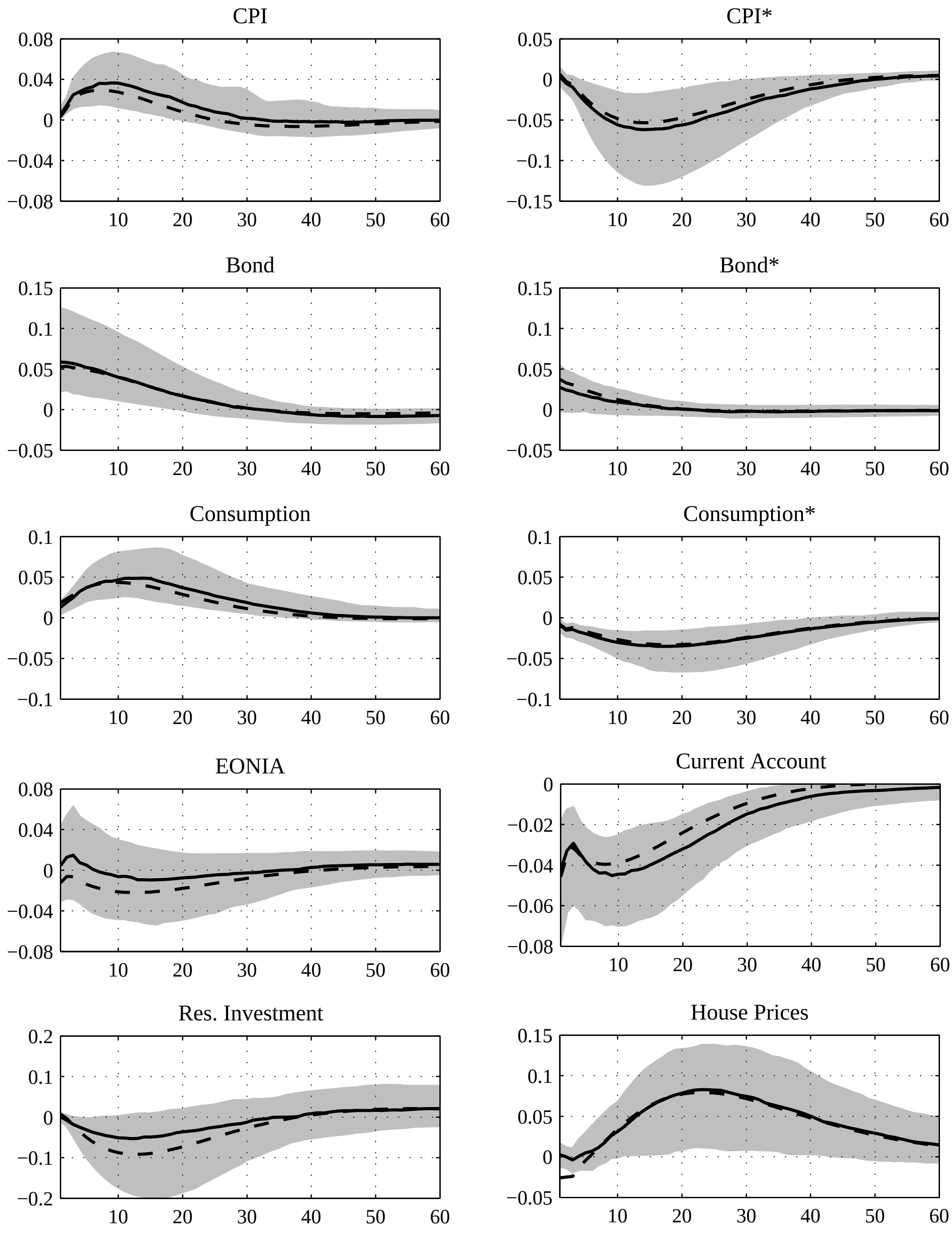
Notes: The x-axis is in months. The solid line represents the median impulse response functions from the BVAR. Shaded areas display 16% and 84% percentiles of the posterior distribution and the dashed line denotes the median target (see Fry and Pagan, 2011).

Figure 13: Risk Premium Shock



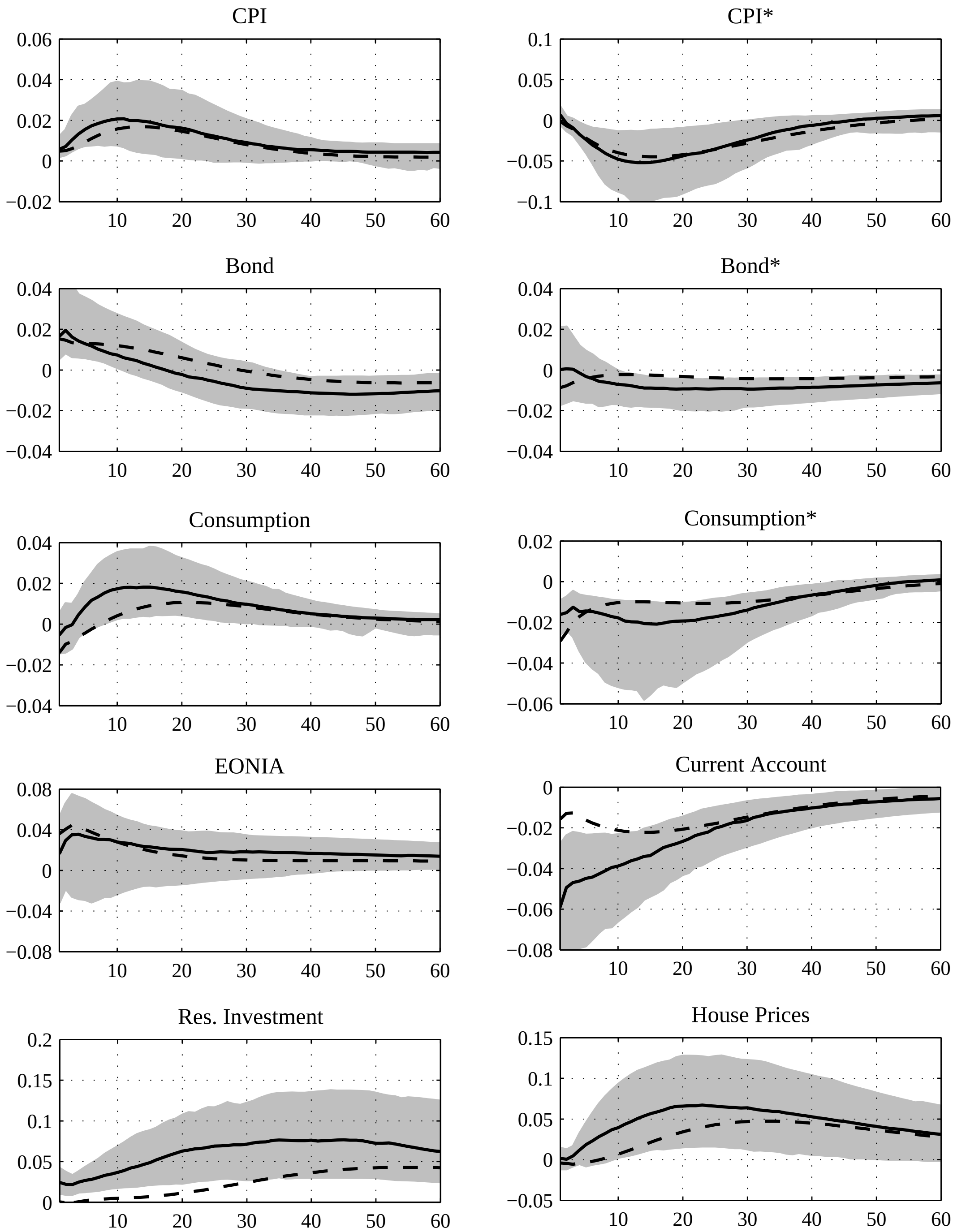
Notes: The x-axis is in months. The solid line represents the median impulse response functions from the BVAR. Shaded areas display 16% and 84% percentiles of the posterior distribution and the dashed line denotes the median target (see Fry and Pagan, 2011).

Figure 14: Financial Easing Shock



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 15: House Price Expectations Shock



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).