Can misguided monetary policy explain the European housing bubble?

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Abstract

Using a standard SVAR framework, this dissertation analyses the role of house prices within the monetary transmission mechanism in Europe over the past decades. Real estate values responded negatively to contractionary monetary policy and thereby amplified effects of consumer spending. However, the interpretation of these results is complicated by non-linear house price dynamics during the early 2000s. A statistical test developed by Homm and Breitung (2012) is therefore used to identify bubble periods in the various countries analysed. Once the test results are fed into the SVAR, the measured effect of monetary policy on house prices remains negative, but to a lesser extent. Overall, evidence found here suggests that monetary policy alone was not responsible for the European housing bubble. The boom is better explained by joint effects of loose money, financial liberalisation and associated mortgage market innovations. This paper is therefore in favour of current efforts to raise capital standards and thereby improve the robustness of the financial system.

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1 Introduction

During the Great Moderation central banks in the US and Europe primarily focused on keeping consumer prices and output stable. It was soon recognized, however, that too narrow an agenda may lead to increased asset price volatility (Iacoviello 2000), which in turn compromises financial stability (Bernanke and Gertler 2000 & Bjørnland and Jacobsen (2013)). A theoretical rationale for explosive bubbles in stock prices was established already in the early 1980s in a series of papers including Blanchard and Watson (1982) and Diba and Grossman (1988b), but this debate was mostly left to academics and went largely unnoticed by policy makers. The common perception among the latter was, and still is, that asset prices are determined by the markets (Rajan 2011); factors unobserved by the social planner make it hard to rightly determine asset values, so price setting is better left to competitive agents within the market (International Monetary Fund 2000). Unfortunately, theoretical predictions did not take long to materialize in the bubbles and subsequent financial crises of the 1980s and 1990s in Scandinavia and Japan that brought devastating and prolonged disruptions to these economies. They were soon followed by the burst of the dotcom bubble in 2001 - another setback, this time sending the United States and Europe into temporary recessions where particularly the former experienced a sharp increase in the level unemployment (Rajan 2011). One should think that these were all warning signs of a general trend that ultimately culminated in the worst financial crisis since the Great Depression. Yet, even in 2005 the incumbent Chairman of the US Federal Reserve, Alan Greenspan, refused to call the national housing bubble by its name (Del Negro and Otrok 2007). His European counterpart Jean-Claude Trichet argued that "it is clearly not opportune to introduce asset prices into a monetary policy rule" in 2002, then still acting as the Governor of the Bank of France (Trichet 2003). In both regions policy rates were kept at record lows and deviated negatively from rule prescriptions at that time (Dokko et al. 2011 & Taylor (2007)). In principal this was done in an effort to stimulate demand and hence accommodate the early 2000s recessions (Rajan 2011); yet, loose money also decreased bond yields and thereby pushed investors to adjust their portfolios in favour of riskier, more lucrative assets such as mortgage securities (2011). With respect to Europe Fernandez-Villaverde, Garicano, and Santos (2013) argue that under the cover of reduced exchange rate risk the introduction of the euro led to information asymmetries and thereby set the stage for an influx of cheap credit where it simply was not due. Besides delaying urgent institutional reforms in many peripheral countries, this overdose of finance inflated construction industries in most regions apart from Germany (2013). Coupled with a sudden and rapid expansion of securitisation (Altunbas, Gambacorta, and Marques-Ibanez 2009) real estate prices entered the same explosive, upward path as they did in the US. Today, few would therefore deny the fact that loose money amplified forces that drove the credit and housing bubbles, which would paralyse global economies when they finally burst in 2007 and 2008 (Aoki, Proudman, and Vlieghe 2004).

This begs painful questions: why were interest rates kept so low for so long? Had economists and central bankers underestimated the effects of loose monetary policy on asset price volatility? Motivated by these questions researchers have put monetary economics back on the agenda after it had long been ignored in the macroeconomic literature (Goodhart and Hofmann 2008). A growing body of empirical literature has recently shed light on the role of house prices within the monetary transmission mechanism (MTM) using mostly vector autoregressive frameworks. A negative response of house prices to a contractionary interest rate shock has generally been found, although magnitudes have varied depending on sample periods and model specifications. One of the earliest papers uses a vector error correction model to analyse a set of European countries between 1975 and 2000 (Iacoviello 2000). The author finds particularly pronounced responses for the UK and Italy, where house prices decrease by roughly three percent in response to a one percentage point increase in the interest rate. For France the measured response of house prices is fairly low, decreasing by around one percent, whereas in Germany no significant response is found at all. Giuliodori (2005) uses a recursive VAR for roughly the same sample period and finds that house prices decrease by around two percent in France, the UK and Finland. A response of around 1.5 percent is found for Italy, Ireland and the Netherlands and a slightly smaller response for Spain. Goodhart and Hofmann (2008) set up a panel VAR for 17 industrialized countries and find an overall response of more than three percent covering observations up until 2006. One of the most recent works on Europe includes data on the actual crisis years and finds an average response of 1.6 percent for all countries excluding Germany (Musso, Neri, and Stracca 2010). Finally, Bjørnland and Jacobsen (2013) investigate the US and identify their VAR using a combination of short- and long-run restrictions and find a very pronounced response of four percent for a sample period ending in 2010.

On the back of these findings this dissertation tries to answer whether the responsibility for the crisis lies primarily with monetary policy. A recursive SVAR as in Giuliodori (2005) is run for a sample of eight European countries including observations from the most recent ten years following the crisis. Higher magnitudes than in previous works are found found. Since house prices over this period were subject to non-linear dynamics in most regions, the estimated effect of the policy instrument on fundamental house prices may be upwards biased as clarified in a later section. A statistical bubble test developed by Homm and Breitung (2012) is therefore used to identify bubble periods. The results are used to control for non-linear house price dynamics within the SVAR in order to test if the measured effects are robust or to some extent biased by an unobserved non-fundamental component in house prices. While trend smoothing filters have been used in a similar manner in Goodhart and Hofmann (2008), explicit information about explosive bubbles has not been integrated in the previous SVAR literature to the best of the author's knowledge. The negative relationship between house prices and tight money holds, although it is generally less significant. However, the effects found here are still too small to conclude that the policy instrument on its own was the driver of the house price boom. In order to analyse to what extent increased levels of securitisation can help to explain the crisis, mortgage market variables are added to the analysis. Forecast error variance decompositions demonstrate that, jointly, mortgage market variables and monetary policy have had a significant impact on house price volatility. It is therefore concluded that focusing research and policy measures primarily on interest rate setting is likely to be unsatisfactory. The impact of financial liberalization on stability calls for stricter and more focused intervention. This dissertation is therefore in favour of central banks' recent efforts to improve macroprudential regulation and strengthen banking supervision.

The following section begins by introducing the theoretical underpinnings of the role of house prices for consumption within the MTM. The second section presents the underlying data. Section 4 introduces the econometric methods used. The results section begins with the the baseline estimation, which only considers variables for output, inflation, house prices and monetary policy. The model is then extended to include consumption in the following section. The focus then shifts to non-linear house price dynamics in section 5.4. Finally, the role of securitisation within the MTM is analysed in section 5.5 before concluding in the final section.

2 The effect of house prices on consumers within the MTM

The ultimate goal of policy evaluation is to analyse its final impact on consumers. Housing represents an important if not the largest component of private wealth, so one should expect that price developments have an impact on consumers' expected lifetime income and hence on consumption (International Monetary Fund 2000). A a nominal appreciation of house price values also leads to an increase in collateral values thereby improving agents' borrowing positions almost immediately (2000). In 2006 the ECB estimated that on average housing made up roughly 60 percent of total household wealth in the euro area (European Central Bank 2006). At that point European households had come out of short period of explosive growth of their housing wealth as a share of total wealth from about 300 percent in 1995 to 450 percent only ten years later (2006). Once the house price boom came to an abrupt halt around 2007, household wealth and consumption took a sharp hit. This draws a rather extreme picture of how monetary policy can affect spending through its effect on house prices and can hardly serve as a theoretical foundation.

Even in a less dramatic environment one can explain the underlying relationships to help understand the empirical analysis that is to follow. The effect of housing wealth on lifetime income is referred to as *life-cycle effect*. Additionally, the literature broadly distinguishes between two other forces: an amplifying *credit channel effect* and *confidence effects*. Each effect may cause a boost in spending in response to an increase in house prices. The life-cycle effect is the most straightforward to understand, although some issues are worth mentioning. As noted by Giuliodori (2005) within the life-cycle framework one should not necessarily expect and increase in aggregate consumption since positive and negative spending effects for 'winners' and 'losers' of a house price increase should cancel each other out. Broadly speaking homeowners and tenants fall into the former and latter category, respectively. Additionally, Mishkin (2007) notes that owner-occupiers are faced with increased opportunity costs of housing as property values and rents rise and might therefore choose to cut down spending on non-housing goods. In theory, the implications of the life-cycle effect are therefore ambiguous. However, as pointed out by Goodhart and Hofmann (2008), it is plausible to assume that the ones who gain from an increase in house prices are more likely to consume out of housing wealth and therefore more sensitive to changes in property prices than the cohort identified as 'losers'. Thus, it is often argued that gains in consumption outweigh the losses and aggregate spending can therefore be expected to increase (Goodhart and Hofmann 2008).

The credit channel is perhaps the most interesting to analyse within the MTM framework since it is understood to strengthen conventional interest rate effects on consumption (Bernanke, Gertler, and Gilchrist 1999). It works twofold: firstly, on what is referred to as the *balance sheet* side, an increase of the policy rate is assumed to widen the spread between the cost of borrowing external funds and earnings on internal funds (Iacoviello and Minetti 2008). This reflects a raise in what is sometimes called the *external finance premium*. With respect to housing one can think of higher down payments and mortgage rates to have a negative effect on housing demand and overall spending. Secondly, through the *bank lending channel* tight money decreases banks' liquidity incentivising them to adjust their supply of non-liquid credit such as mortgage loans downwards. Both effects therefore work in the same direction as monetary policy and consequently amplify its impact on spending (Bernanke, Gertler, and Gilchrist 1999).

Finally, much like stock prices house prices may act as an indicator with respect to the overall state of the economy and may in that sense boost investor and consumer confidence (International Monetary Fund 2000 & Giuliodori (2005)). This is plausible to the extend that house price developments are strongly correlated with economic growth (International Monetary Fund 2000). As Giuliodori (2005) argues, positive perceptions about future developments of aggregate consumption and hence income usually stimulate consumption in the short term. However, Mishkin (2007) points to a feature that distinguishes property prices from stock prices in this respect: the former may grow simply because the short-term supply of housing is fixed and therefore do not necessarily reflect future expectations about productivity. Possibly, more sophisticated agents can distinguish between the two sources of house price growth, which would decrease potential confidence effects.

Empirical findings have largely supported the idea that monetary indirectly affects consumers through the house price channel and its overall effect on spending is therefore strengthened. Catte et al. (2004) estimate positive short- and long-run marginal propensities to consume out of housing wealth for a number of industrialized countries. In the UK, for example, an increase of real estate wealth by one pound is found to increase consumption by eight pence in the short term (Catte et al. 2004). Similarly, Slacalek and others (2009) find positive values for all European countries considered in this dissertation with the exception of Finland. Within the MTM literature Giuliodori (2005), for instance, also finds positive responses of consumption to innovations in house prices. As mentioned before, this issue will be investigated empirically in one of the following sections.

3 Data

For the empirical investigation quarterly indicators for inflation, output, consumption, real estate prices and mortgage variables were used. Sample periods generally span the last 37 years implying a number of about 150 observations per variable and country. To account for the heterogeneity of European housing markets the model has been run for a set of eight countries: Germany, Spain, Finland, France, United Kingdom, Ireland, Italy and the Netherlands. This selection broadly overlaps with the countries analysed Giuliodori (2005) who limited his analysis to the pre-euro era ending in late 1998. This paper therefore provides some insight about how the role of house prices within the MTM has potentially changed over the most recent two decades. Data availability varied by region and variable, but for the majority of countries data has been collected from the OECD and Oxford Economics through the Datastream database. For the inflation variable, logged differences of the consumer price index are used in the model. Consumption and output are both measured at constant prices and expressed in total expenditure terms. The house price variable is a property price index expressed in real terms and sourced from the OECD. Inevitably, there exists some country-specific heterogeneity with respect to the construction of these indices. For the policy instrument the 3-month interbank rate was used in line with previous studies. It was collected from the OECD for most countries with the exception of Finland, where only IMF data was available for the entire sample period. Mortgage market data was collected through Datastream and primarily provided by Oxford Economics. Total outstanding mortgage liabilities were used for the mortgage stock. The mortgage rate is defined as the interest rate on building society mortgages. Data on mortgage debt was only available in nominal terms and therefore deflated using each country's CPI series before entering the model. For Germany only monthly data was available for the mortgage rate, so averages were taken. Details about sample periods and data sources for each country and variable can be found in tables (1), (2) and (3) in the appendix.

4 VAR Methodology

The model specification is in line with existing studies and follows most closely the frameworks in Giuliodori (2005) and Goodhart and Hofmann (2008). The same general set up and identification strategy is applied to all eight countries to be able to compare the results across regions. The reduced form VAR with p lags is defined in equation (1)

$$\mathbf{Y}_t = \mathbf{A}_1\mathbf{Y}_{t-1} + \mathbf{A}_2\mathbf{Y}_{t-2} + ... + \mathbf{A}_p\mathbf{Y}_{t-p} + u_t$$

$$\mathbf{Y}_t = \mathbf{A}(L)\mathbf{Y}_t + u_t \tag{1}$$

where $\mathbf{Y}_t$ represents the vector of endogenous variables , $\mathbf{A}(L)$ is the coefficient lag polynomial[1] and u are reduced form errors.

If the VAR is stable and hence invertible, impulse response functions (IRF) can be derived by restating the reduced form VAR in equation (1) in its Wold moving average representation (Lütkepohl and Krätzig 2004):

$$\mathbf{Y}_t(\mathbf{I} - \mathbf{A}(L)) = u_t$$

$$\mathbf{Y}_t = (\mathbf{I} - \mathbf{A}(L))^{-1}u_t$$

$$\mathbf{Y}_t = \mathbf{C}(L)u_t \tag{2}$$

The coefficients contained in $\mathbf{C}(L)$ are best understood as the effects of reduced form shocks u_t in any given period $t - s$ on variables $\mathbf{Y}_t$, s periods ahead. However, there is no reason to assume that reduced form shocks to the different variables are independent of each other, which inhibits a clear interpretation of equation (2). Specifically, there are likely to exist unobserved contemporaneous effects, $\mathbf{B}_0$, determining the relationships between the endogenous variables. In order to derive the structural VAR in equation (3), both sides of equation (1) are pre-multiplied by this contemporaneous effects matrix

$$\mathbf{B}_0\mathbf{Y}_t = \mathbf{B}_0\mathbf{A}_1\mathbf{Y}_{t-1} + \mathbf{B}_0\mathbf{A}_2\mathbf{Y}_{t-2} + ... + \mathbf{B}_0\mathbf{A}_p\mathbf{Y}_{t-p} + \mathbf{B}_0u_t$$

$$\mathbf{B}_0\mathbf{Y}_t = \mathbf{B}_0\mathbf{A}(L)\mathbf{Y}_t + \mathbf{B}_0u_t$$

$$\mathbf{B}_0\mathbf{Y}_t = \mathbf{B}(L)\mathbf{Y}_{t-1} + \varepsilon_t \tag{3}$$

where $\mathbf{B}(L)$ and ε_t represent structural coefficients and innovations, respectively. The latter are normalized such that the structural error covariance matrix, Σ_ε, is equal to the identity matrix, $\mathbf{I}$, as defined in equation (4):

[1] Defined as $A(L) = A_1L + A_2L^2 + ...$

$$\Sigma_u = \mathbf{B}_0^{-1}\Sigma_\varepsilon \mathbf{B}_0^{-1\prime} = \mathbf{B}_0^{-1}\mathbf{B}_0^{-1\prime}$$

$$\mathbf{B}_0 \Sigma_u \mathbf{B}_0{}' = \Sigma_\varepsilon = \mathbf{I} \tag{4}$$

To estimate and identify the SVAR with k endogenous variables one needs to solve equation (4) where the k^2 parameters of $\mathbf{B}_0^{-1}$ are unknown. Note that only $\frac{(k+1)k}{2}$ elements in Σ_u can be uniquely estimated in the reduced form VAR (Gottschalk 2001), since off-diagonal elements in the lower and upper triangle are duplicates of each other. Thus, one needs to impose restrictions on the $\frac{(k-1)k}{2}$ remaining parameters of $\mathbf{B}_0^{-1}$ for the model to be just identified. Equation (2) can then restated in terms of its structural innovations to yield

$$\mathbf{Y}_t = \mathbf{C}(L)\mathbf{B}_0^{-1}\mathbf{B}_0 u_t$$

$$\mathbf{Y}_t = \mathbf{C}^*(L)\varepsilon_t \tag{5}$$

where $\mathbf{C}^* = \mathbf{C}(L)\mathbf{B}_0^{-1}$. By virtue of the orthogonality restriction in equation (4) the effects contained in $\mathbf{C}^*(L)$ have a meaningful interpretation, since off-diagonal elements in Σ_ε are zero and structural shocks therefore happen in isolation (Gottschalk 2001). Thus, equation (5) allows us to estimate how the different endogenous variables respond to an impulse to any one of them, which will be the main focus of the following empirical analysis.

Following the majority of existing literature a Choleski decomposition of $\mathbf{B}_0^{-1}$ is used to restrict the contemporaneous parameters recursively. The baseline estimation follows Giuliodori (2005) and includes only the following endogenous variables in this order: the change in log consumer prices, Δcpi, log output, Δy, log real house prices, Δhp, as well as the short-term interest rate, i, in levels. The former three enter the model in differences since for most countries they were found to be integrated of order one. Since the change in log prices reflects inflation, Δcpi will be notated as π in the following. Upon inspection of individual time series inflation and interest rates were found to exhibit deterministic trends and were consequently detrended before entering the VAR[2].

With $k = 4$ adding the required six zero restrictions yields the baseline model

$$\begin{bmatrix} \pi_t \\ \Delta y_t \\ \Delta hp_t \\ i_t \end{bmatrix} = \mathbf{C}(L) \begin{bmatrix} B_{\pi,\pi} & 0 & 0 & 0 \\ B_{y,\pi} & B_{y,y} & 0 & 0 \\ B_{hp,\pi} & B_{hp,y} & B_{hp,hp} & 0 \\ B_{i,\pi} & B_{i,y} & B_{i,hp} & B_{i,hp} \end{bmatrix} \begin{bmatrix} \varepsilon_{\pi,t} \\ \varepsilon_{y,t} \\ \varepsilon_{hp,t} \\ \varepsilon_{i,t} \end{bmatrix} \tag{6}$$

Note that the ordering of the variables matters when using these short-run restrictions. Under the current specification the monetary policy instrument i is assumed to be the only variable that reacts contemporaneously to all other variables in the system. The remaining variables react with a lag to at least one of the other variables. Thus, house prices are assumed to react to changes in the interest rate only after the first quarter as in previous studies (Iacoviello (2000) and Giuliodori (2005)). Both output and inflation also react sluggishly to the interest rate in line with the MTM literature and standard macroeconomic theory (Bjørnland and Jacobsen 2013).

[2]Time series plots for all variables and countries available upon request.

Alternative variable orderings and model specifications are thinkable and have been used in the past. Iacoviello (2005) assumes that monetary policy makers do not act immediately to changes in house prices. Bjørnland and Jacobsen (2013) incorporate stock prices in their SVAR and argue that asset prices and interest rates react simultaneously making it necessary to use a combination of short-run and long-run restrictions. Although they find a comparably strong response of house prices to changes in the interest rate looking at US data, the qualitative results are consistent with other papers. Finally, as in Goodhart and Hofmann (2008) one might change the ordering of inflation and growth and assume that Δy reacts with a quarterly lag to all other variables. Different orderings were tested here and did not change the main qualitative results.

Previous studies also use various approaches to deal with integration and cointegration of variables. In principal, if some variables are integrated of order one, as in the underlying case, non-stationarity compromises the model's stability and hence its invertibility (Lütkepohl and Krätzig 2004). The VAR can still be estimated consistently though, if only some variables are non-stationary (Sims, Stock, and Watson 1990). Many authors therefore choose to still let all variables enter in levels mainly to avoid overdifferencing and losing information contained in levels (Giuliodori 2005 & Musso, Neri, and Stracca (2010)). The levels specification is also adequate if cointegration exists between some of the I(1) variables. Ideally, error correction terms are then included in the model making the VAR stationary without eliminating information about equilibrium relationships as in Iacoviello (2000). Differencing the relevant variables also resolves the non-stationarity issue, but potentially miss-specifies long-term relationships (Kilian and Lütkepohl 2016). For the underlying case, the levels and VECM specifications were considered, but finally rejected in favour of differencing non-stationary variables. Tables (4) and (5) in the appendix present unit root test outcomes for each country and variable in levels and differences, respectively. Non-stationarity was found for many variables and largely avoided through differencing. For the sake of consistency and comparability the same set of variables was differenced for each country, although this led to overdifferencing in some cases as evident from Table (5).

There was only limited evidence for cointegration. For the baseline model, the Johansen procedure revealed one cointegrating rank for only one country (Juselius and others 1992). Although higher ranks were generally found once consumption and mortgage variables were included, the model specification was kept consistent.[3]

Despite its caveats, differencing fared better than levels with respect to serial correlation of residuals, most visibly shown in Tables (8) and (9) in the appendix. Serial correlation may bias the estimated coefficients and makes statistical inference based on standard errors invalid (Jeffrey and others 2009). Note that while almost no serial correlation was found when taking differences and running the model over a reduced sample (1980 - 1998), some serial correlation remained once all observations were included as evident from Tables (6) and (7). This issue is likely due to non-linear dynamics of some variables during the boom period in the early 2000s. Musso, Neri, and Stracca (2010) similarly find that VAR residuals increase towards the far end of their sample. While this issue is tackled in section 5.4, significance intervals for the baseline estimation should be considered with some caution.

[3]Results not reported. Available upon request.

5 Results

5.1 Baseline estimation

To test for a negative relationship between interest rates and house prices in Europe the baseline model is run for the aforementioned countries covering a timespan from 1980 Q1 to 2016 Q4. Lag lengths were chosen based on the Akaike Information Criterion and varied by country with the maximum lag length set to five, which produced stable VAR equations in most cases.[4]

The resulting impulse response functions of the different variables to a one-unit interest rate shock are depicted in Figure 1 and Figure 2.[5]

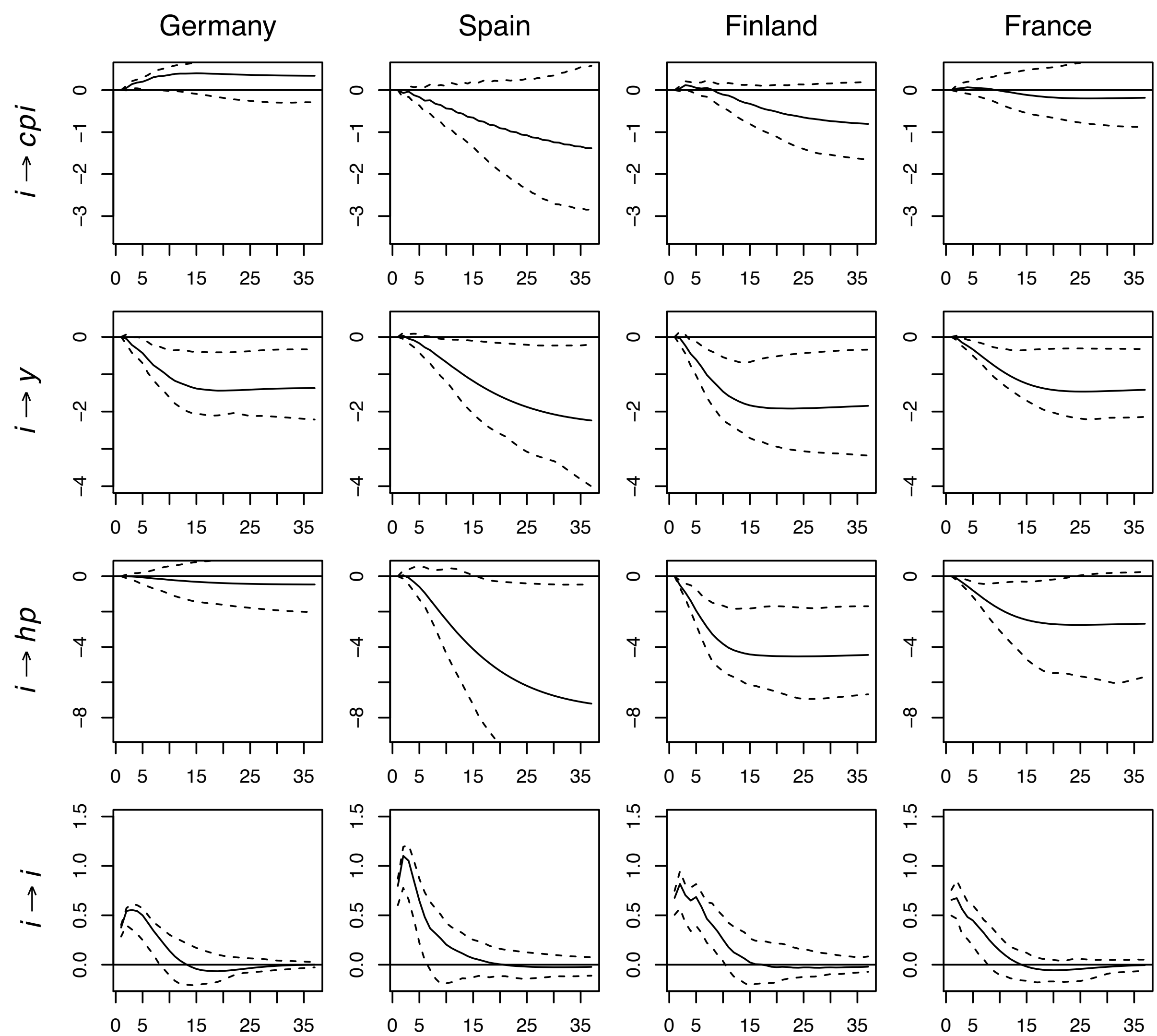

Figure 1: Impulse responses of all variables to an interest rate shock for Germany, Spain, Finland and France. Dashed lines are 95 percent bootstrapped error bands. Axes are standardised to ease visual comparison.

[4]Stability was tested using OLS-CUSUM tests as suggested in Pfaff (2008).
[5]All computations were implemented in R.

The responses in terms of percentage changes of the respective variable are shown for each country by column. Only the IRF for the interest rates is measured on a percentage-point scale. For variables that have entered the model in differences, the impulse response functions are cumulative. 95 percent confidence intervals are illustrated as the area between the dashed lines. As expected, substantial heterogeneity with respect to different countries is found. The qualitative findings generally reflect the standard relationships found in the MTM literature and predicted by Keynesian monetary theory (Sims 1992).

Following the unit shock, the interest rate starts at varying values generally slightly below one depending on the estimated contemporaneous effect of i on itself contained in $\mathbf{B}_0$. It subsequently gradually reduces to zero after around three to four years.

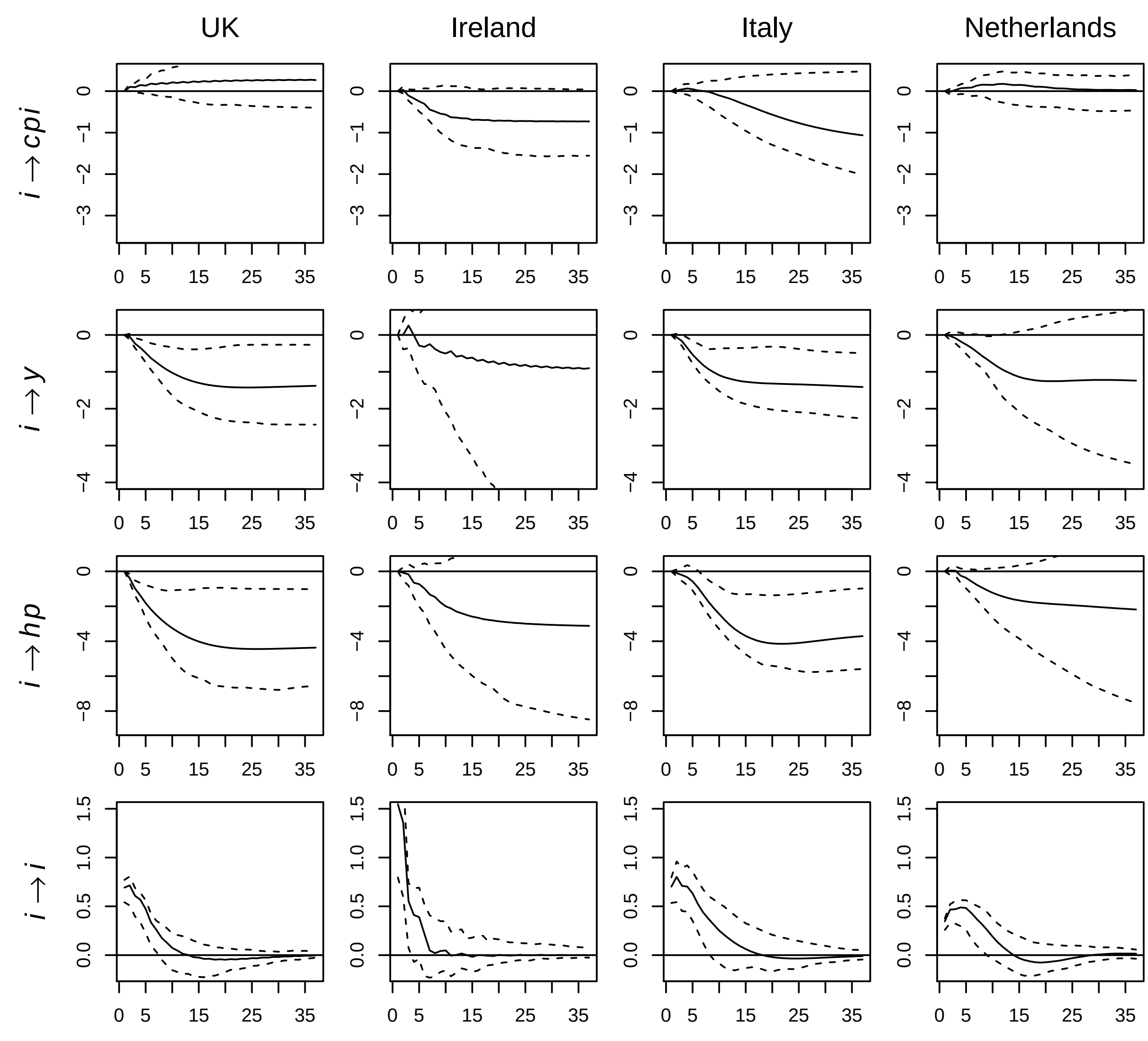

Figure 2: Impulse responses of all variables to an interest rate shock for the UK, Ireland, Italy and the Netherlands. Dashed lines are 95 percent bootstrapped error bands. Axes are standardised to ease visual comparison.

Prices are initially sluggish to respond in some countries, but subsequently fall in most places, which is the response predicted by theory. Note that while the consumer price variable enters the model as an inflation rate, the cumulative IRFs show the percentage change of the CPI. In Spain, for example, the price level is found to decrease by roughly two percent after five years. The literature has often found that prices are not only rigid, but actually rise following an interest rate rise. Sims (1992) explains that this "price puzzle" is due to the fact that policy makers react to anticipated inflationary pressures, hence an interest rise is often followed by a short period of inflation. In Germany and the UK the response actually remains positive which is counter-intuitive. A similar result was also obtained in Goodhart and Hofmann (2008).

An increase in the interest rate stiffens demand and therefore leads to a significant decrease in output. In most countries GDP decreases by slightly less than two percent only three to four years after the shock. Interestingly, the measured effect is slightly higher than in Giuliodori (2005) and more in line with Bjørnland and Jacobsen (2013) who also include observations from the post-crisis era.

A negative effect of tight money on house prices is evident for all countries although with varying degrees of significance. As in Giuliodori (2005) the effect is particularly pronounced in Finland and the UK where house prices here decrease by roughly five percent after four years. In most other countries responses lie between two and three percent with the exception of Germany where the effect is weak and statistically insignificant. Contrary to Giuliodori (2005) a very strong response is found for Spain. Magnitudes here are generally higher by about two to three percentage points and more in line with Goodhart and Hofmann (2008) who also uses the variables in differences. To control for whether the higher magnitude is due to differencing the model has also been rerun in levels and still slightly stronger responses were found for most countries. Since Giuliodori (2005) explicitly focuses on the pre-EMU era from 1979 to 1998, a straightforward conclusion might be that monetary policy had a stronger influence on house prices during and around the global financial crisis (Goodhart and Hofmann 2008). This is analysed further in the following sections.

Finally, it should be noted that results for Germany are generally somewhat peculiar, something also observed in Musso, Neri, and Stracca (2010). Fernandez-Villaverde, Garicano, and Santos (2013) point out that in contrast to most other European countries financial conditions for Germany did not change with the introduction of the Euro, since the eurozone interest rate was essentially aligned to the German one. Secondly, Germany was actually considered as the "sick man" of Europe at that point with a rapidly ageing population and an urgent need for reform. Together these factors impeded a strong increase in the demand for housing and other goods and may explain why the impulse responses found for Germany follow uncommon paths.

5.2 Relative contributions to house price fluctuations

Since by construction the VAR explains all variation in its endogenous variables, one might expect that the estimated absolute impact of *all* variables in the system on house prices was stronger during the boom period. Impulse response functions focusing exclusively on the nominal effect of a shock to monetary policy may therefore be too narrow and misleading. Instead "innovation accounting" can help to understand how strongly interest rates have influenced house prices relatively compared to the remaining variables (Lütkepohl 2005). Figure 3 shows forecast error variance decompositions of house prices for each country. The different areas represent the average proportion of house price volatility accounted for by innovations in each of the four variables over five years[6] (Lütkepohl 2005).

[6] A complete derivation demonstrating how this is done can be found in the appendix.

Interestingly, the relative contribution of the interest rate appears to be marginally lower than in Giuliodori (2005). Monetary policy shocks explain slightly less than 20 percent of house price fluctuations in Finland, Spain and Italy and less than 10 percent in all other countries. Moreover, for all countries but the UK, shocks to the house price equation itself drive house price volatility more persistently than in Giuliodori (2005). Even after five years, around 70% of house price fluctuations are explained by movements in the price itself in many cases. It will be shown below that from a theoretical standpoint this is not a surprising observation when bubbles are present and determine the growth path of asset prices. Together these findings imply that in relative terms, the effect of monetary policy on house prices was not actually stronger during recent decades. Rather, it seems that to this point unobserved factors have driven house prices. Section 5.5 will turn back to this issue.

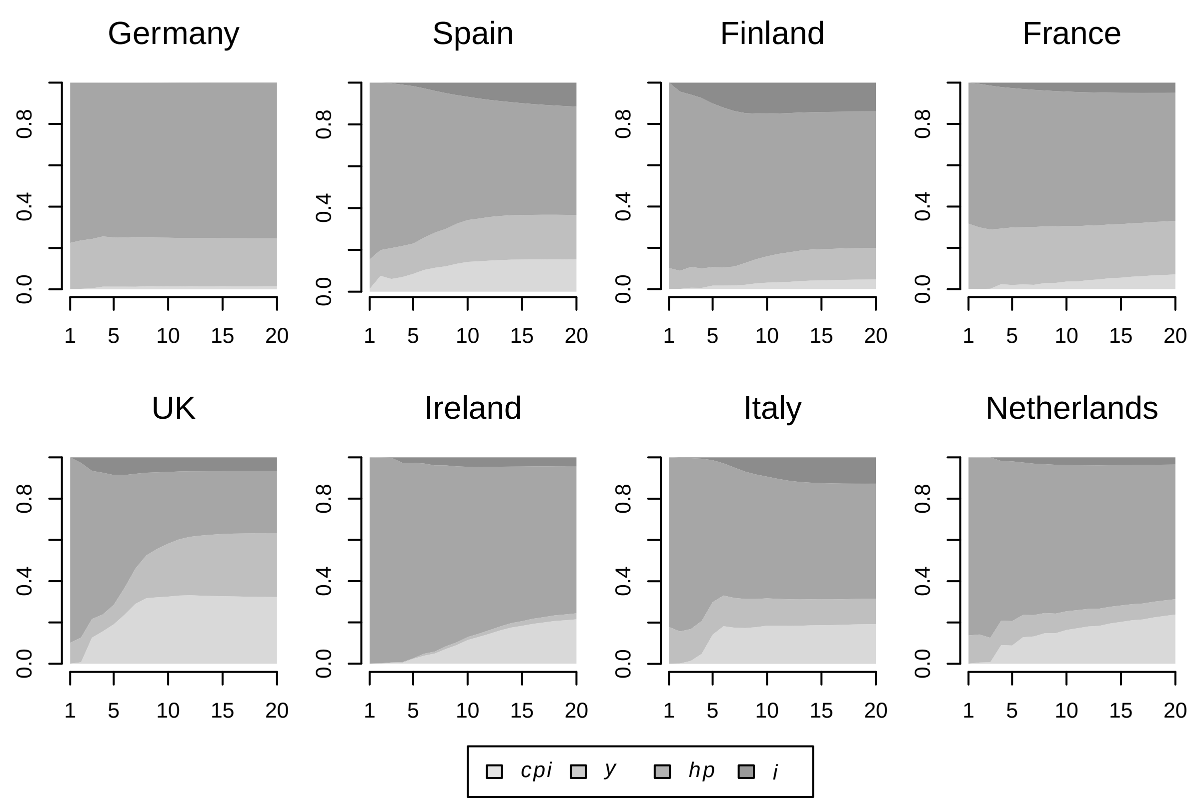

Figure 3: Forecast error variance decomposition over twenty quarters. Percentage contribution to volatility on vertical axis.

5.3 The effect of house prices on private consumption within the MTM

To draw conclusions for policy makers, it is important to analyse how interest rates actually influence consumers through house prices via the various channels outlined earlier. For this purpose the baseline model is extended by an equation for private consumption. The new vector of endogenous variables is ordered as in Giuliodori (2005)

$$\mathbf{Y}_t = \begin{bmatrix} \pi_t & \Delta y_t & \Delta c_t & \Delta hp_t & i_t \end{bmatrix}'$$

where Δc_t represents differenced logged values of private consumption and everything else is defined in the same way as above. Responses of consumer spending to a one-percentage point increase in house price growth provide insight about the direction of the real estate channel (Giuliodori 2005). They are plotted for each country in Figure 4 in the same way as earlier. In most countries spending is stimulated growing by up to two percent in Spain and Ireland after just three to four years. The same qualitative responses were also found in Giuliodori (2005) for the majority of countries and reflect the idea that an increase in housing wealth has a positive affect on consumption. The opposite is the case in Italy, where consumers cut spending. For France and Germany the relationship is not very clear and in Italy consumers appear to actually reduce spending by one percent in response to a house price shock.

To obtain a more explicit idea of the magnitude of the credit channel, a simple counterfactual simulation can be used to produce an estimate of the propagating contribution of house prices within the MTM (Giuliodori 2005 & Sims and others (1998)). In order to do so, the response of consumer spending to an interest rate increase are firstly calculated using the coefficients from the unrestricted SVAR as under the baseline specification. For the counterfactual estimation, the same coefficients are used, but house prices are restricted from interacting with any of the other variables. Following Giuliodori (2005) this is done by setting to zero the relevant coefficients in the structural equations of inflation, output, consumption and the interest rate, which yields

$$\mathbf{B}_t = \begin{bmatrix} B_{\pi,\pi} & B_{\pi,y} & B_{\pi,c} & 0 & B_{\pi,i} \\ B_{y,\pi} & B_{y,y} & B_{y,c} & 0 & B_{y,i} \\ B_{c,\pi} & B_{c,y} & B_{c,c} & 0 & B_{c,i} \\ B_{hp,\pi} & B_{hp,y} & B_{hp,c} & B_{hp,hp} & B_{hp,i} \\ B_{i,\pi} & B_{i,y} & B_{i,c} & 0 & B_{i,i} \end{bmatrix}_t$$

where $\mathbf{B}_t$ represent the coefficient matrices within the lag polynomial $\mathbf{B}(L)$ in equation (3). Specifically, zeros are imposed on contemporaneous effects as well as on lags.

The resulting impulse responses for both the unrestricted model and the counterfactual simulation are plotted as solid and dashed lines, respectively, for each country in Figure 5. Note that as expected a negative response of spending to tight money is found in each case where magnitudes are generally on the range of one to two percent. More importantly, in countries where previously consumers were found to react positively to an increase in house prices, spending is indeed less sensitive to monetary shocks once the house price channel is closed. In those countries the dashed line lies well above the solid one. In many cases like Spain and the UK the negative effect of contractionary monetary policy on spending is about twice as strong if amplified by the credit channel. As expected, for Italy the propagating contribution of house prices to monetary policy is negative. These mixed results found here reflect the concerns raised by Mishkin (2007), that the theoretical implications of housing wealth effects remain somewhat unclear. Still, in numerous countries, there is concrete evidence that house price deviations strengthen rather than attenuate the effects of monetary policy on consumers.

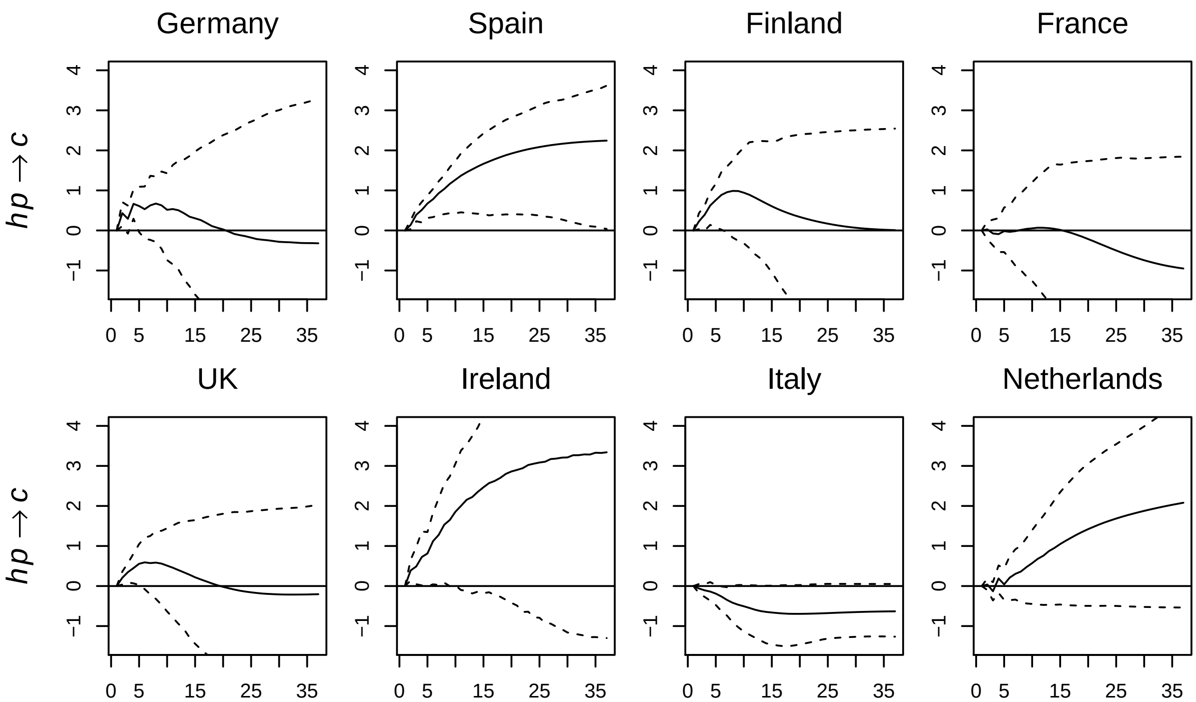

Figure 4: Response in consumtion to a one percentage point increase in house price growth.

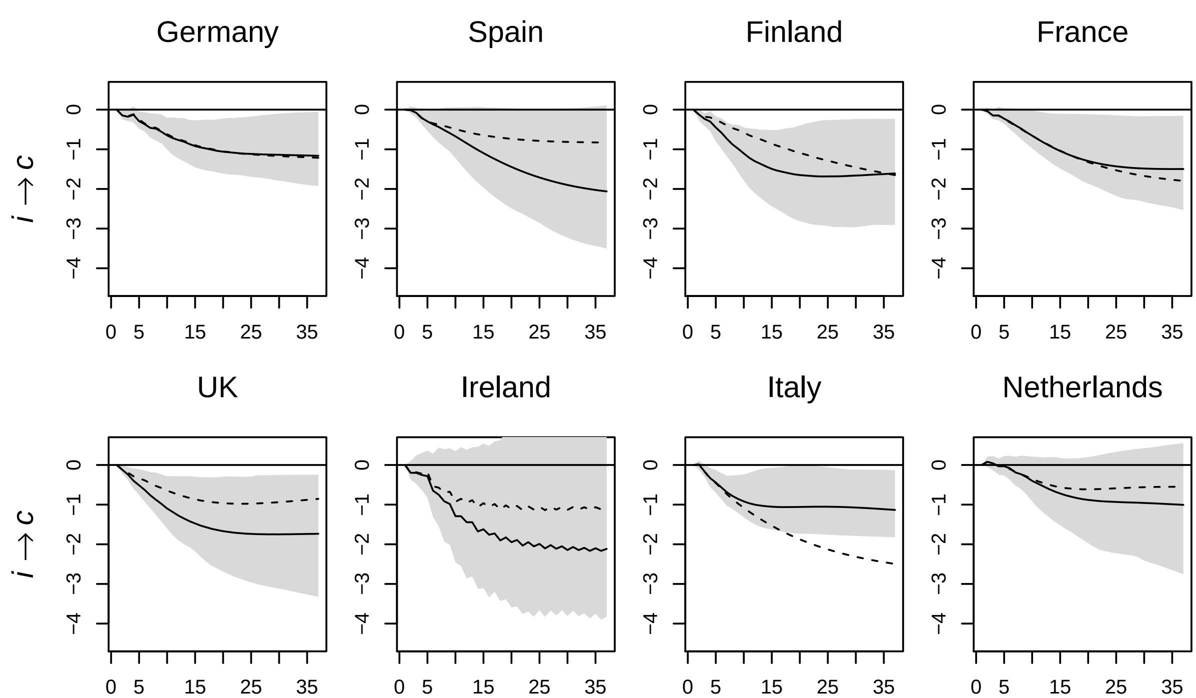

Figure 5: Solid lines reflect responses of consumption to a one percentage point increase in the monetary interest rate with 95 percent confidence intervals illustrated as shaded areas. Dashed lines represent responses under the counterfactual simulation.

5.4 Non-linear house price dynamics

VARs generally model relationships linearly and therefore fail to fully capture any non-linear dynamics (Musso, Neri, and Stracca (2010)). Many existing studies have pointed to the issue that if asset prices are included in a VAR and speculative bubbles were present at any point within the sample period, then the effect of monetary policy on the fundamental component of house prices may be overestimated (Iacoviello 2000). Several approaches have been used to overcome or at least moderate this problem. Bjørnland and Jacobsen (2013) augment their model by various impulse and step dummy variables to control for the boom period and important events such as the Lehman bankruptcy. As expected, they consequently find a decrease in the measured effect of the interest rate on house prices. Goodhart and Hofmann (2008) augment their model using an indicator variable for house price booms. They similarly find higher estimates for the effect of monetary policy on house prices during boom periods. This section analyses the implications of these findings and investigates the issue empirically using a bubble testing procedure designed by Homm and Breitung (2012), briefly presented in the following.

5.4.1 Bubble theory and statistical testing

The theory on asset bubbles is derived from and based around the net present value model. Prices are assumed to reflect the value of discounted future earnings with perfectly informed agents (Cochrane 2009). In empirical works dividends usually enter the model for stock earnings, whereas rents are used in the real estate context (Homm and Breitung 2012) yielding the following equation

$$F_t = \sum_{i=1}^{\infty} \delta^i E_t(R_{t+i}) \tag{7}$$

where F_t is the fundamental value of real estate, $\delta = \frac{1}{1+r} < 1$, is a constant discount factor depending on the discount rate r, and R_t represent rents. Diba and Grossman (1988a) derive a theoretical framework for asset bubbles from equation (7). They demonstrate that if at time zero agents assess the actual property value to be higher than its fundamental value by some positive value b_0, this belief will confirm itself in future periods and hence rationalize the bubble. It can easily be shown that in equilibrium the price $\hat{P}_t$ will then no longer reflect the fundamental asset value, but rather incorporate information that is not related to any market-fundamentals yielding

$$\hat{P}_t = b_t + F_t \tag{8}$$

where b_t follows an explosive process:

$$E_t[b_{t+i}] = (\frac{1}{\delta})^i b_t \tag{9}$$

This has two important implications. Firstly, note that for the fundamental value of property in equation (7) to be finite, the long-term growth rate of rents needs to be smaller than r (Shiller 1980). Since b_t, however, grows at rate r this implies that the overall growth rate of house prices will be primarily driven by the bubble. The high levels of persistency found earlier when inspecting the variance decomposition plots therefore seem to reflect this theoretical prediction. Secondly, rents are generally assumed to be generated by a random walk process (Homm and Breitung 2012). This implies that the observed real estate price should also follow a random walk unless it contains an explosive bubble component b_t, which lays the foundation for the majority of bubble detection tests including the one applied in this work. The *supDFC* test from Homm and Breitung (2012) recursively checks for a regime switch of the underlying time series from I(1) to explosive. Comparing various bubble tests in empirical applications they find their own test to be particularly powerful at identifying bubble break dates (2012).[7]

For the underlying analysis, the test has been applied to the real house price indices of the different countries. Results indicate that house prices grew explosively over varying time spans in all countries but Germany. Bubble periods are indicated in Figure 6 below as shaded areas where the solid lines map the observed house price indices. Note that across countries the test indicates break dates in the 1990s - as early as 1993 Q2 for Finland - preceding the crash of the dotcom stock bubble. The fact that price anomalies of different assets coincided shows that asset price volatility was generally high during this time as discussed earlier.

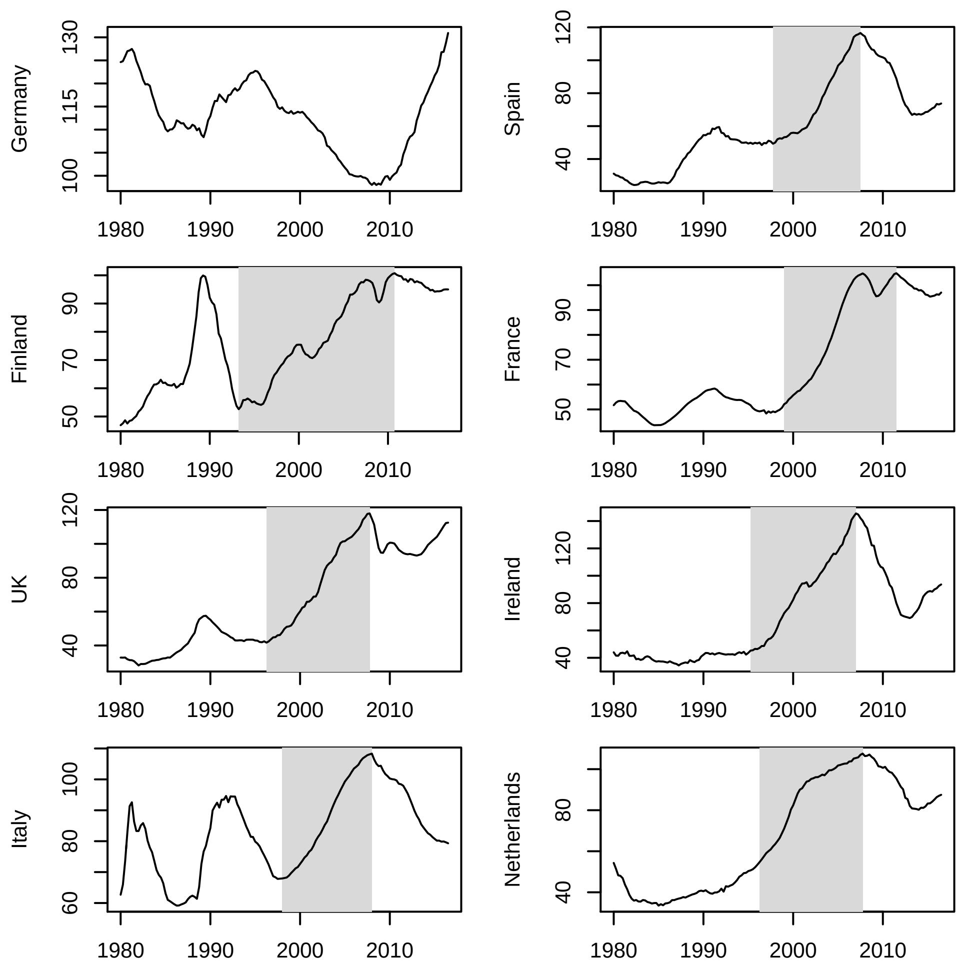

Figure 6: House price trends with shaded areas indicating bubble periods

[7]For a detailed explanation of how the test works please refer to the appendix.

5.4.2 Controlling for bubbles in the SVAR

Having identified bubble periods in the different countries this information is now incorporated in the SVAR to gain a better understanding of the fundamental relationship between monetary policy and house prices. Following Goodhart and Hofmann (2008), the first and most straightforward approach is to simply rerun the baseline model for reduced sample periods excluding the grey areas in Figure 6 from the estimation. This way deals with non-linear dynamics to the extent that it simply gets rid of them. Based on the findings in previous sections, one would expect a contractionary monetary shock to have a less significant effect on real estate prices. Empirically, the bubble test results are used to construct a dummy variable, $\mathbf{D}_t^{NB}$, that equals one during no-bubble periods and zero whenever a bubble is present. Since Germany did not actually go through a property bubble, tentative bubble start and break dates (1997 Q1 - 2007 Q1) have manually been added to still check for whether the effect of monetary policy differed. Augmenting the baseline model by this dummy yields:

$$\mathbf{B}_0\mathbf{Y}_t = \mathbf{B}(L)\mathbf{Y}_{t-1} \times \mathbf{D}_t^{NB} + \varepsilon_t \tag{10}$$

The resulting impulse response functions are plotted against the baseline model results in the left column of Figure 7 and Figure 8. For each country by row the red lines represent the responses of house prices to a contractionary monetary policy shock under the new specification with corresponding 95% confidence intervals illustrated as red shaded areas. The black lines represent the baseline results. To highlight that Germany did not actually go through a bubble the response under the new model is pictured in grey. As expected and in line with Goodhart and Hofmann (2008), the qualitative reponses do not change but decrease significantly in magnitude for all but one country. Responses get somewhat closer to the levels observed in Giuliodori (2005), where house prices in France and Italy now decrease by one and roughly three percent, respectively, three years after the policy shock. The response in the UK is still high at approximately three percent, but reduced nonetheless. In Spain the response of house prices has been halved although it still negative as in Giuliodori (2005). Finland represents an outlier where a stronger decrease under the new specification is may be explained by the earlier bubble at the end of the 1990s. The test used here fails to deal with repetitive bubbles. For Germany the result remains as insignificant as under the baseline specification.

Using a reduced sample comes at the cost of losing information. A way to avoid this and still control for a bubble effect is to simply add an exogenous impulse or step dummy to the VAR as in Bjørnland and Jacobsen (2013). This is realized by constructing a dummy variable, $\mathbf{D}_t^{B}$, defined in the opposite way of $\mathbf{D}_t^{NB}$: it switches from zero to one as soon as the bubble breaks out and back to zero once it bursts. Augmenting the baseline model yields

$$\mathbf{B}_0\mathbf{Y}_t = \mathbf{B}(L)\mathbf{Y}_{t-1} + \Phi\mathbf{D}_t^{B} + \varepsilon_t \tag{11}$$

where Φ is the coefficient matrix on the step dummy. The dummy is best understood as a regime shift indicator, where the entire system enters a new regime as house prices go from I(1) to explosive. Note that under both specifications the bubble is assumed to be exogenous, which is in line with its theoretical definition of being unrelated with any fundamental factors. While this model transformation remains a modest step aimed at overcoming issues related with non-linear dynamics of the house price variable, Clements and Mizon (1991) argue that dummy augmented linear models can often serve as good approximations. The results very much mirror what was

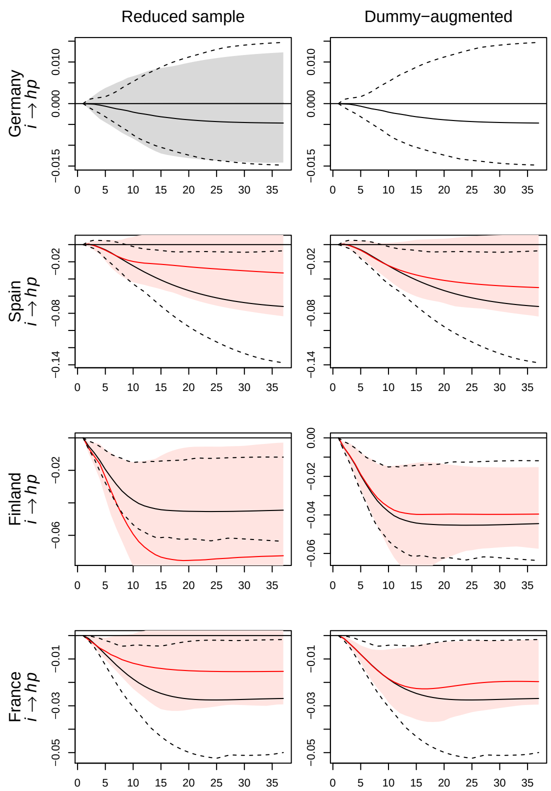

Figure 7: Responses of house price to an interest shocks under the reduced sample specification on the left and the dummy-augmented model on the right. Baseline responses in black. Germany, Spain, Finland and France.

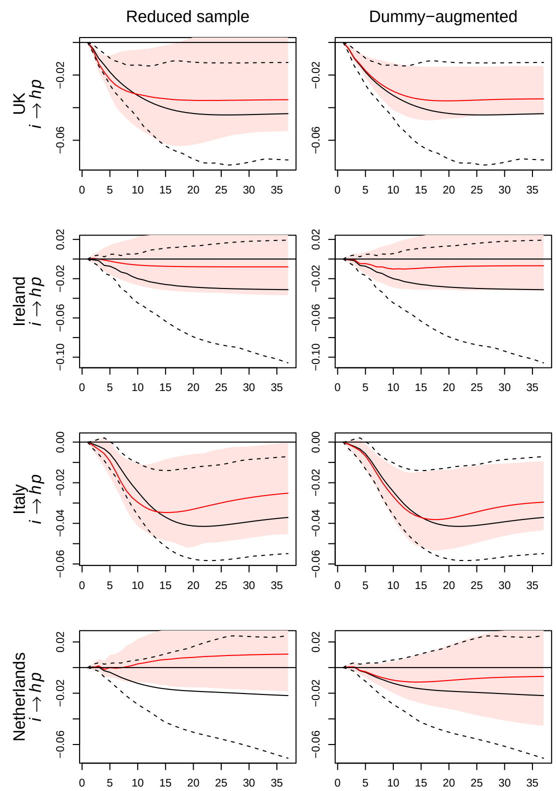

Figure 8: Responses of house price to an interest shocks under the reduced sample specification on the left and the dummy-augmented model on the right. Baseline responses in black. Germany, Spain, Finland and France.

observed using the reduced sample above. IRFs are depicted in the right column of Figure 7 and Figure 8 where once again the black curves reflect the baseline responses and the red elements represent the output of the modified model. The model is not run for Germany in this case. In most countries the negative response of house prices to a policy shock is less pronounced when controlling for bubbles. The decrease in magnitude is somewhat weaker than under the reduced sample specification. Importantly, even with the step dummy included, the qualitative responses remain the same, which is consistent with what was found in Bjørnland and Jacobsen (2013).

Table (10) in the appendix sheds some light on how the two model specifications deal with the issue of serially correlated errors, which was previously assumed to be caused by non-linear dynamics. Evidently, once the identified bubble periods are excluded from the estimation the reduced sample SVAR is free of serial correlation in all but one cases. The dummy-augmented model fails to resolve the issue, which implies that this simple method falls somewhat short of fully incorporating non-linearity.

Still, the two approaches yield uniform implications. Identifying and controlling for extraordinary dynamics in the model makes the effect of monetary policy on house prices less dramatic than the baseline model originally suggested. Various conclusions can be derived from this observation depending on how one interprets observed house prices. Goodhart and Hofmann (2008) conclude that monetary policy has a stronger effect on prices during boom periods. An alternative conclusion may be that the VAR overestimates the interest rate's effect on *fundamental* house prices. Suppose that the observed house price is composed of its fundamental part and some positive bubble component, where the latter is unobserved. As previously discussed a large and increasing part of house price dynamics are then subject to this non-fundamental factor. If the bubble component is unobserved the model may overestimate the contribution of the interest rate to house price dynamics.

5.4.3 Scenario based analysis

Even in recognition of advanced statistical tests for bubbles, the question about whether or not econometricians can truly distinguish between fundamental asset price movements and speculative bubbles is an ongoing debate. Gürkaynak (2008), for instance, review existing testing procedures for asset bubbles and conclude that results remain somewhat unsatisfactory. In cases like this, when economic theory is not well established, simply letting the data speak may provide further insight. A number of works have made use of conditional VAR projections to construct *what-if* type scenarios. Trying to answer how house prices would have behaved under different circumstances, the paths of other variables are exogenously implied or relationships between some variables set to zero.

Taylor (2009), for instance, analyses how house prices in the US would have behaved, if the Fed had explicitly followed the Taylor Rule, which acts as guidance for policy makers. The implied interest rate would have been up to three percentage points higher just before the crisis in 2004 and based on the VAR projection, property price growth would consequently have been lower. Dokko et al. (2011) also point out that the interest rate set by the Fed was lower than the rule prescription. They also include a number of European countries in the analysis and find similar results. However, they argue that these deviations from the rule were by no means unprecedented. For example, the rate had been substantially too low during the 1970s and excessively high during the disinflation era under Volcker (2011).

Dokko et al. (2011) also construct conditional VAR forecasts for property prices and interest rates based on actual outcomes of all other variables, but restricting the effects of housing market variables and the policy rate on all other variables. They find that simulated house prices are far below actual realizations, whereas the interest rate projection diverges only slightly from its actual growth path. This finding implies that house prices have played a large role in determining their own growth path and hence provides some evidence for a bubble. As Stiglitz (1990) once described: "[I]f the reason that the price is high today is only because investors believe that the selling price is high tomorrow – when 'fundamental' factors do not seem to justify such a price – then a bubble exists." Similarly, Jarocinski and Smets (2008) generate conditional forecasts from a Bayesian VAR and find that low interest rates in the US have only a limited ability to explain the excessive increase in house prices.

Independent of how one chooses to interpret house prices, the analysis so far has shown that the sustained high supply of money undeniably added to the inflation of real estate prices rather than helping to avoid it (Dokko et al. 2011). Still, even in countries like the UK, where strong links persist between monetary policy and the housing market, an interest rate drop of two to three percentage points over the boom period fails to explain a near tripling of property prices. This suggests that other forces were also at work and hence the model is extended to include further variables in the following section.

5.5 Market for mortgages

Many have argued in the past that financial liberalization and resulting innovations in the mortgage market led to a vast expansion of available credit and thereby contributed to the housing bubble - most significantly so in the US, but also in many European countries (Dokko et al. 2011). In case of the former, packaging of fixed-rate mortgages originated as a financial instrument in the mid 1980s (Green and Wachter 2005). While in principal this provided a way to deal with high inflation and the consequential rise in nominal interest rates typical for that time (2005), securitisation reached unsustainable levels in the early 2000s (Mayer, Pence, and Sherlund 2009). Mortgage backed securities more than doubled in total value from \$3.0 trillion to \$6.9 trillion between 2000 and 2007 (Dokko et al. 2011). Europe was originally much slower to adopt these financial products as pointed out by Jaffee and Renaud (1995). The introduction of the euro, however, opened the monetary union up to new forms of capital (Fernandez-Villaverde, Garicano, and Santos 2013). Asset-backed derivatives became quickly popularized (Altunbas, Gambacorta, and Marques-Ibanez 2009); total securitisation activity measured in terms of the number of euro-denominated asset-backed securities outstanding increased sixfold from 2000 until the credit bubble burst in mid 2007. Thus the sudden impact of financial deregulation was even stronger in some European countries than in the US.

Unsurprisingly, many have drawn a connection between loose monetary policy and securitisation commonly arguing that the latter amplified the conventional credit effects of the former (Dokko et al. 2011). Maddaloni and Peydró (2011) test this hypothesis and indeed find that low short-term interest rates coupled with high degrees of securitisation have had a strong negative impact on lending standards in Europe. As a consequence all countries except Germany have experienced an unprecedented increase in the growth of outstanding mortgage liabilities. This section begins by exploring the relationships between monetary policy and mortgage market variables. The analysis then turns to a comparison of the relative contributions of money and mortgage channels to house price volatility.

Following the previous literature the baseline model has been augmented by mortgage variables such that the new vector of endogenous variables is given as

$$\mathbf{Y}_t = \begin{bmatrix} \pi_t & \Delta y_t & \Delta hp_t & i_t & \Delta ms_t & mr_t \end{bmatrix}'$$

where ms_t represents the mortgage stock in log differences and mr_t is the detrended interest rate on building society mortgages. For the structural model the Choleski decomposition is maintained as before. Thus, monetary policy is assumed to react with a lag to mortgage variables implying that credit reacts immediately to monetary policy rather than vice versa. Though somewhat arbitrary this assumption is plausible and commonly made in the literature (Goodhart and Hofmann 2008). The chosen ordering also follows Giuliodori (2005), so once again this allows for a direct comparison of the results. The augmented model has been rerun for all countries and impulse response functions have been plotted in the same way as before in Figure 9 and Figure 10. Additionally, responses for the sample excluding the bubble periods as specified in equation (10) have been added in red. It should be noted that for some countries data availability with respect to the mortgage stock was limited and sample lengths therefore differ making a cross-country comparison more problematic in this section. Sample start dates for Ireland and France are 2003Q1 and 1994Q4, respectively, which led to short sample issues for these countries.

Still, the overall evidence is conclusive. The relationships between the original endogenous variables are very much in line with the findings under the baseline specification, not only qualitatively but

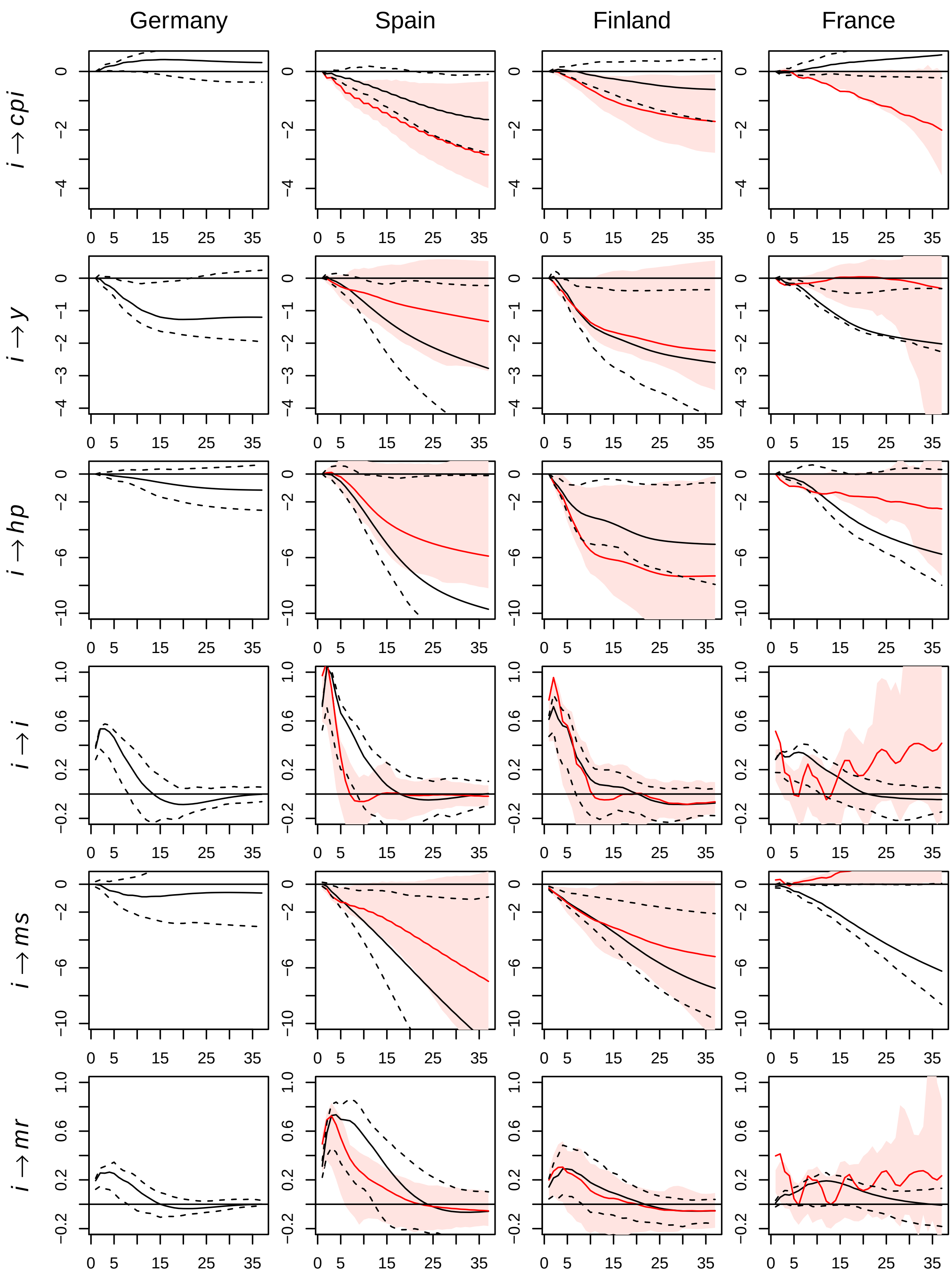

Figure 9: Impulse responses for extended model including mortgage variables in black. Red elements represent results using a reduced sample excluding bubble periods. Germany, Spain, Finland and France.

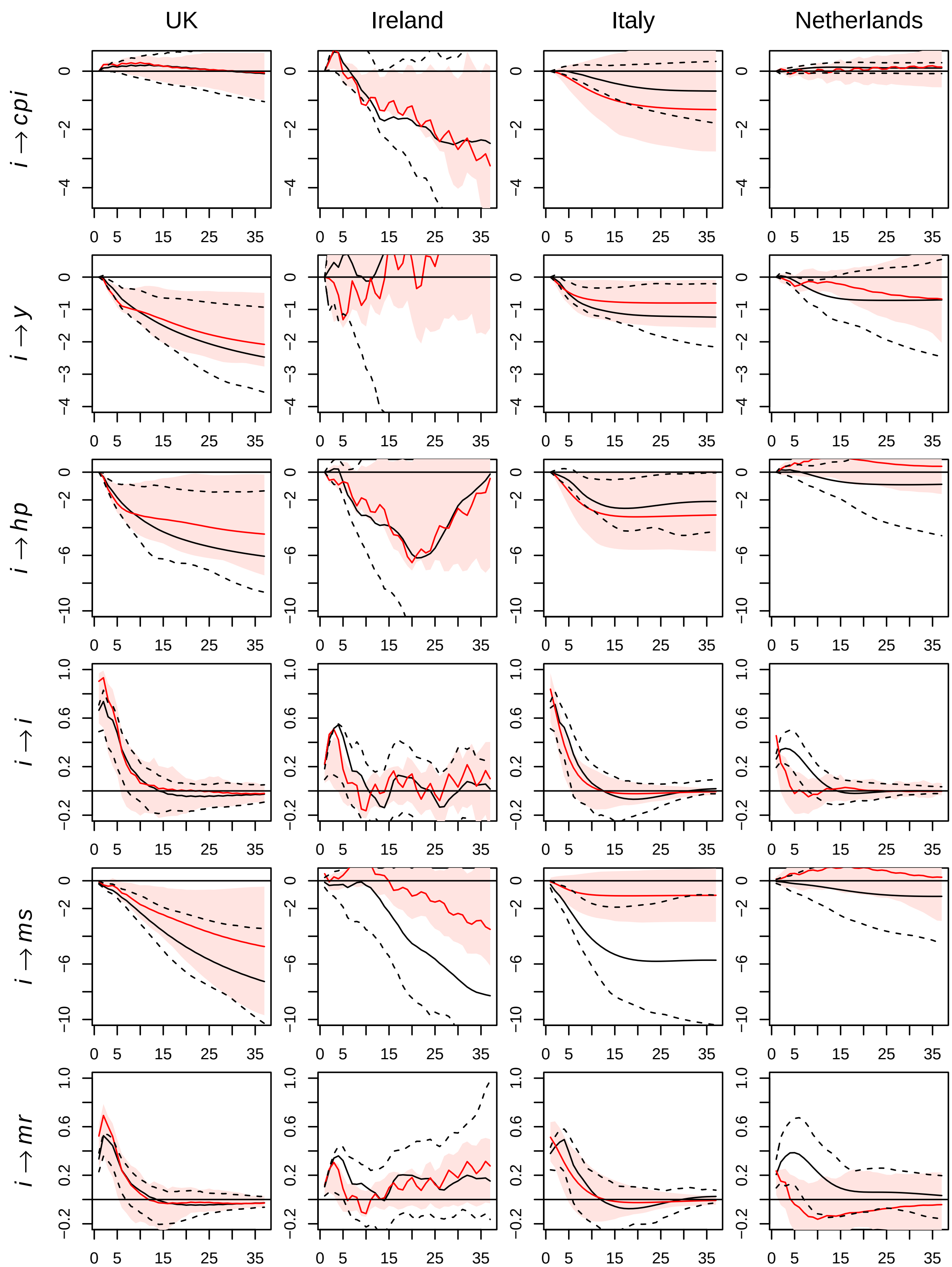

Figure 10: Impulse responses for extended model including mortgage variables in black. Red elements represent results using a reduced sample excluding bubble periods. UK, Ireland, Italy and the Netherlands.

also in terms of magnitudes. Looking at the UK, for example, the responses for inflation, output, house prices and the interest almost exactly mirror the ones obtained earlier. Once again it can be concluded that the model specification is robust.

Contractionary monetary policy has the expected negative pass-through effect on mortgage credit in all countries. The same is observed in Giuliodori (2005), although just as under the baseline specification the responses found here are slightly higher and more in line with Goodhart and Hofmann (2008) and Musso, Neri, and Stracca (2010). Outstanding liabilities fall by as much as 10 percent in countries like Spain and Finland, which appears extremely large. Also as expected, the interest rate on mortgages follows the path of the monetary interest rate closely in all countries, although the response is generally around 10 to 20 basis points smaller in magnitude. This can be interpreted as short-term stickiness of mortgage rates and was also found in (Musso, Neri, and Stracca 2010). Despite some rigidity though, there is clear evidence of a pass through from nominal interest rates to the actual price of taking out a mortgage, which directly affects consumers (Giuliodori 2005).

Under the reduced sample specification that controls for bubble periods, the response to a monetary shock is not only moderated for house prices. Production, too, decreases less significantly, whereas monetary policy has actually been more effective in determining consumer prices in no-bubble years. The former finding may indicate that output was to some extent also inflated by the housing bubble. After all, the construction industry was a main driver of growth in the year leading up to the crisis. In Spain, for instance, it attracted one in every four male workers during its peak (Fernandez-Villaverde, Garicano, and Santos 2013). The latter finding may reflect the fact that conventional monetary policy loses its influence on prices around the zero lower bound (Bernanke and Reinhart 2004). Most notably perhaps, the pass-through effect of interest rates on credit availability was apparently lower and more “normal” during the no-bubble period. This is an important finding since it supports the point made earlier, that the renovation of European mortgage markets through the rapid introduction of securitisation during the early 2000s amplified the impact of loose monetary policy on credit availability (Maddaloni and Peydró 2011). Finally, for France the reduced sample estimation produces odd results, probably due to issues related to the smaller sample size. Overall, the results clearly indicate that as suggested above mortgage variables do indeed play a significant role within the MTM.

The analysis can again be aided by exploring the relative contributions of all variables to house price volatility. If mortgage market innovation rather than monetary policy was the main driver of the real estate bubble as argued by Dokko et al. (2011), then the relative contribution of monetary policy should diminish once mortgage variables are included. Variance decompositions have been calculated and plotted in the same way as earlier for each country in Figure 11. Dark and light red areas represent the proportionate shares of the mortgage stock and rate, respectively. The light blue area shows the contribution of the monetary interest rate to house price fluctuations over time. For some countries such as the UK, Ireland and the Netherlands it seems as if mortgage variables have greater explanatory power, whereas for Spain and Finland the opposite appears to be the case. This evidence is therefore not conclusive. More interestingly though, the joint relative impact of monetary policy and the mortgage market is significantly higher than the impact of monetary policy alone observed earlier in Figure 3. It explains as much as 40 percent of house price fluctuations in countries like Ireland and the Netherlands. Generally, there seems to be less persistency with regard to how house prices determine their own growth path. In most countries less than 50 percent of house price volatility can now be explained by house prices themselves, considerably less than the levels of up to 70 percent found under the baseline specification. It appears that taking into account the *joint* impact of loose money and increased levels of mortgage availability the model gets closer to explaining the house price boom.

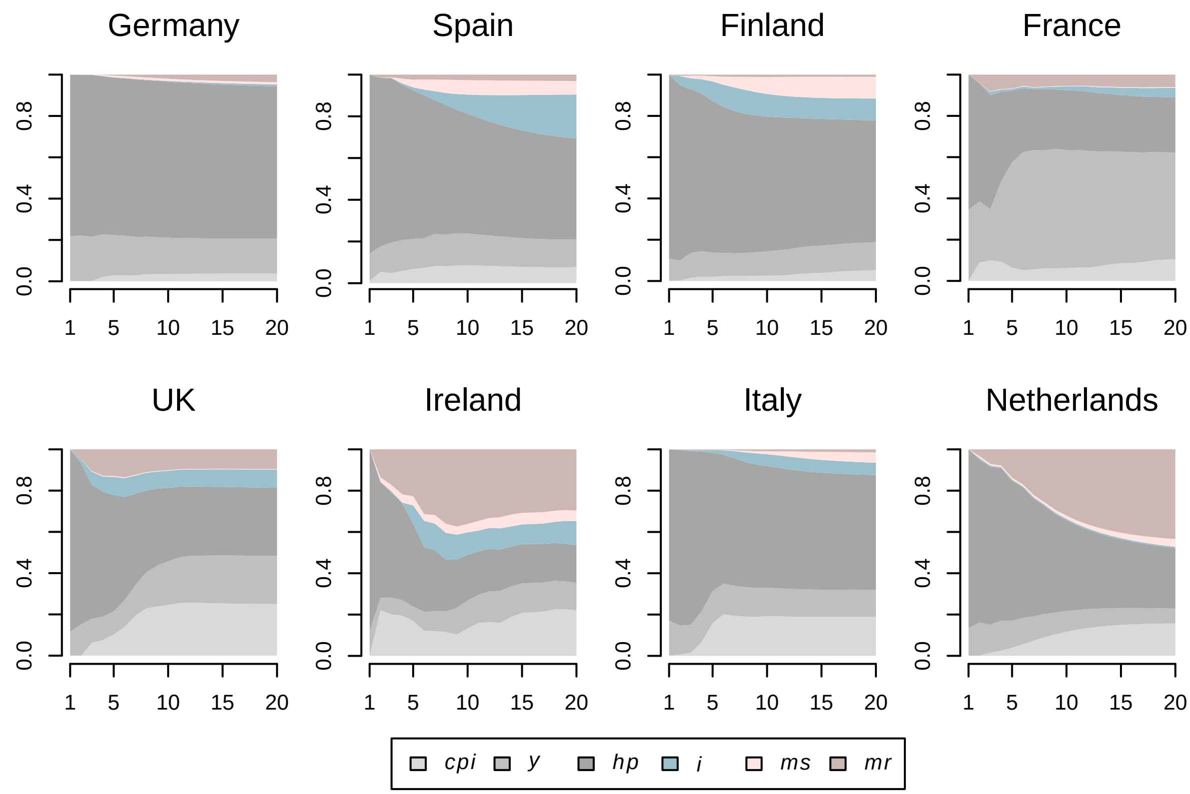

Figure 11: Forecast error variance decomposition for extended model including mortgage variables.

6 Concluding thoughts

Analysing most recent data, the SVAR has produced strong responses of house prices to monetary policy, especially compared to previous papers such as Giuliodori (2005) that focused on the pre-euro era. Even when controlling for non-linear house price dynamics it appears that monetary policy has affected property values more heavily during the boom years. Still, to answer the overriding question of this paper, misguided monetary policy cannot explain the European housing bubble on its own despite the high magnitudes found here. Rather, evidence suggests that the conventional effects of monetary policy on house prices in Europe were amplified by financial liberalisation and associated mortgage market innovations. Jointly these two forces have contributed heavily to house price volatility. The results also suggest that uncovering the roots of the crisis is very complex and holding only certain institutions or individuals accountable for it is complicated. A number of conclusions for future policy can be drawn nonetheless.

While central banks have failed to avoid the crisis, it appears that stricter interest rates more closely aligned with rules could not have entirely avoided the bubbles, hence this approach is not recommended. Putting more focus on asset price stability and thereby departing from the policy rate's traditional role of smoothing consumption and consumer prices is not advisable, either. Trichet

(2003) rightly pointed out that it would be extremely difficult to pin down exactly what asset prices should be targeted and then adjust a single policy instrument to all of them. In the eurozone, matters would be further complicated by regional economic differences as pointed out by Goodhart and Hofmann (2008). Moreover, since stock and house prices tend to move with output anyway, monetary policy already implicitly addresses them (Trichet 2003).

Aside from the interest rate, other more flexible instruments are available to policy makers and promise greater scope. In light of the finding that financial innovations have greatly contributed to the bubbles, policy makers should continue current efforts on imposing stricter regulation through macroprudential measures. Specifically, as argued by Admati et al. (2013) the social benefits associated with stricter capital requirements are likely to outweigh the private costs. While concerns have been raised that stricter requirements would hamper the provision of liquidity and ultimately adversely affect consumers, Admati et al. (2013) question to what extent mortgage debt creation is actually socially desirable. In line with their doubts, recent research in behavioural economics has demonstrated that easy access to housing finance may actually decrease consumer welfare in the presence of self-control issues (Schlafmann 2016). While the rise of subprime mortgages during the boom years certainly improved the financial positions of affected income groups in the short term, the ensuing housing crisis also often had the most detrimental long-term effects on these borrowers.

On the institutional side, an undeniable benefit of lower leverage ratios is that banks will be better equipped to deal with financial distress in the future. Higher capital requirements may therefore help to curb the "too big to fail" issue without explicitly restricting the actual size of banks (Admati et al. 2013). The gradual implementation of the Basel capital standards has already shown some success at improving banks' strength in Europe and the UK as recent stress tests have shown (European Banking Authority 2016 & Bank of England (2016)). With respect to the housing market directly, Eckley et al. (2017) show that since the introduction of Basel II large lenders in the UK have decreased their supply of high-risk mortgages in favour of low loan-to-value loans. They also note, however, that smaller banks have been pushed to specialize in the supply of high-LTV loans. Goodhart and Hofmann (2008) go one-step further suggesting that LTVs should be adjusted countercyclically, directly. Decreasing them in times of high house price growth, while relaxing caps in times of moderate house price inflation could help to decrease risk (2008). Regulation has only recently been tightened and the discussion is slowly beginning to benefit from observing real outcomes. This avenue of research is therefore likely to produce fruitful results in the coming years.

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A Appendix

A.1 Forecast error variance decomposition for SVAR

The following derivation follows Kilian and Lütkepohl (2016) and is applied to the underlying case to give the reader an impression of how $FEVD$s were constructed. For simplicity assume that the lag length is set to one. Starting from the reduced form model one can derive an expression for a h-period forecast using forward iteration as in the following.

$$\mathbf{Y}_{T+1} = \mathbf{A}\mathbf{Y}_T + u_{T+1}$$
$$\mathbf{Y}_{T+2} = \mathbf{A}(\mathbf{A}\mathbf{Y}_T + u_{T+1}) + u_{T+2} \Leftrightarrow [...]$$
$$\mathbf{Y}_{T+h} = \mathbf{A}^h\mathbf{Y}_T + \sum_{i=0}^{h-1}\mathbf{A}^i u_{T+h-i}$$

The h-step ahead forecast error is then

$$FE(h) = \mathbf{Y}_{T+h} - \mathbf{Y}_{T+h|T} = \sum_{i=0}^{h-1}\mathbf{A}^i u_{T+h-i}$$

where A^i can be restated in terms of Wold moving average coefficients $C_i = JA^iJ'$ in equation (2) following Kilian and Lütkepohl (2016) such that

$$FE(h) = \sum_{i=0}^{h-1}\mathbf{C}_i u_{T+h-i}$$

As earlier one can derive the structural form using the fact that $\mathbf{B}_0 u_t = \varepsilon_t$ to:

$$FE(h) = \sum_{i=0}^{h-1}\mathbf{C}_i^* \varepsilon_{T+h-i}$$

For the h-step ahead mean squared forecast error it follows (Kilian and Lütkepohl 2016):

$$MSFE(h) = E[(\mathbf{Y}_{T+h} - \mathbf{Y}_{T+h|T})(\mathbf{Y}_{T+h} - \mathbf{Y}_{T+h|T})'] = \sum_{i=0}^{h-1}\mathbf{C}_i\Sigma_u\mathbf{C}_i'$$

$$MSFE(h) = \sum_{i=0}^{h-1}\mathbf{C}_i^*\Sigma_\varepsilon\mathbf{C}_i^{*\prime} = \sum_{i=0}^{h-1}\mathbf{C}_i^*\mathbf{C}_i^{*\prime}$$

where as above $\Sigma_\varepsilon = \mathbf{I}$. Since interest lies in the relative contributions of shocks to each endogenous variable $j \in \mathbf{Y}$ to the forecast error variance in $k = hp$, let $c_{kj,h}$ be the $kj\hat{}\{\text{th}\}$ element in $\mathbf{C}_h^*$ (Kilian and Lütkepohl 2016). Then, for example, the resulting $MSFE$ for the h-step forecast of hp given a unit shock to the interest rate $i \in \mathbf{Y}$ follows as

$$MSFE_i^{hp}(h) = c^2_{(hp,i),0} + c^2_{(hp,i),1} + [...] + c^2_{(hp,i),h}$$

The total $MSFE^{hp}(h)$ would simply be the sum of $MSFE$s resulting from shocks to each variables $j \in \mathbf{Y}$. Finally, relative contributions can be found by dividing each variables $MSFE$ by the total. As noted in Kilian and Lütkepohl (2016) these ratios then give an impression of how much each variable contributes on average to the overall variation in hp as $h \to \infty$.

A.2 Chow-type test for explosive bubbles

In the following the bubble test proposed by Homm and Breitung (2012) is explained in some detail. As mentioned earlier econometric tests usually test for a regime switch from a random-walk process to an explosive trend. Thus under the null we have

$$H_0: \qquad \rho_t = 1 \qquad \forall \quad t$$

where ρ_t is the coefficient on the lag in an AR(1) representation of the underlying time series hp. Under the alternative hypothesis

$$H_1: \qquad \rho_t = \begin{cases} 1 & \forall \quad t \in [1, (\tau^* T)] \\ \rho^* > 1 & \forall \quad t \in ((\tau^* T), T] \end{cases}$$

where T is the total number of time periods and $(\tau^* T)$ is the unknown break date. In the same way as for a standard Dickey-Fuller test the AR(1) is restated in terms of differences, hence

$$hp_t - hp_{t-1} = (\rho_t - 1)(y_{t-1}\mathbb{I}_{\{t>[\tau T]\}}) + u_t$$

where $\mathbb{I}$ is indicator function equal to one if the time period is above the potential break date (τT) and u_t is white noise. The resulting test statistic DFC_τ for the coefficient $\delta = (\rho_t - 1)$ is estimated recursively where the potential break date is increased by one period each time:

$$H_0: \qquad \delta_t = 0$$
$$H_0: \qquad \delta_t > 0$$

The final test statistic can be found in Homm and Breitung (2012) and has been derived here. One can write

$$DFC_\tau = \frac{\hat{\delta} - \delta_0}{\sqrt{\frac{1}{T-2} var(\delta)}} = \frac{\sum_{t=\tau T+1}^{T} \frac{\Delta hp_t}{hp_{t-1}}}{\sqrt{\frac{1}{T-2} \frac{cov(\Delta hp_t, hp_{t-1})}{var(hp_{t-1})}}} = \frac{\sum_{t=\tau T+1}^{T} \frac{\Delta hp_t}{hp_{t-1}}}{\sqrt{\frac{1}{T-2} \frac{\sum_{t=\tau T+1}^{T} (\Delta hp_t - hp_{t-1})^2}{\sum_{t=\tau T+1}^{T} (hp_{t-1})^2}}}$$

where $\hat{\delta}$ is the estimated coefficient, the equation for $var(\delta)$ is just the standard variance formula for an OLS estimator and $cov(\Delta hp_t, hp_{t-1})$ can be restated in terms of the variance of residuals $u_t = \Delta hp_t - hp_{t-1}$. Multiplying both the numerator and the denominator by $\sum_{t=\tau T+1}^{T} (hp_{t-1})^2$ yields the equation for the test statistic found in Homm and Breitung (2012):

$$DFC_\tau = \frac{\sum_{t=\tau T+1}^{T} \Delta hp_t hp_{t-1}}{\sqrt{\frac{1}{T-2} \sum_{t=\tau T+1}^{T} (\Delta hp_t - hp_{t-1})^2 (hp_{t-1})^2}} = \frac{\sum_{t=\tau T+1}^{T} \Delta hp_t hp_{t-1}}{\sigma_\tau \sqrt{hp_{t-1}^2}}$$

The test is essentially a one-sided t-test where H_0 is rejected for high values of DFC_τ. The break date is identified as $(\tau^* T)$ at the highest test value

$$\sup_{\tau^* \in [0, (1-\tau_0)]} DFC_\tau$$

where the starting value for the recursive iterations is set to $\tau_0 = 0.1$. A bubble is identified whenever the supremum is higher then the 95% critical value of 1.9327 provided in Homm and Breitung (2012). For the empirical implication Matlab codes were thankfully received from Joerg Breitung and Konstantin Kholodolin and subsequently translated into R.

A.3 Data characteristics

	Germany	Spain	Finland	France
PI	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q4	1980 Q1 - 2016 Q3
CONS	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q4	1980 Q1 - 2016 Q3
Y	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q4	1980 Q1 - 2016 Q3
HP	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q4	1980 Q1 - 2016 Q3
I	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q4	1980 Q1 - 2016 Q3
MS	1980 Q1 - 2016 Q3	1983 Q3 - 2016 Q3	1980 Q1 - 2016 Q4	1994 Q4 - 2016 Q3
MR	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q4	1980 Q1 - 2016 Q3

Table 1: Sample periods for each variable and country

	UK	Ireland	Italy	Netherlands
PI	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q3
CONS	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q3
Y	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q3
HP	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q3
I	1980 Q1 - 2016 Q3	1984 Q1 - 2016 Q3	1980 Q1 - 2016 Q3	1986 Q1 - 2016 Q3
MS	1980 Q1 - 2016 Q3	2003 Q1 - 2016 Q3	1980 Q1 - 2016 Q3	1990 Q4 - 2016 Q3
MR	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q3	1980 Q1 - 2016 Q3

Table 2: Sample periods for each variable and country

	Description	Source	Exceptions
PI	Inflation: log CPI	OECD Main Economic Indicators	none
CONS	Consumption: Final household consumption	OECD Economic Outlook	Germany and Ireland (Oxford Economics)
Y	Output: log GDP, seasonally adjusted at constant prices	OECD Economic Outlook	Germany (Oxford Economics)
HP	House prices: House price index, real	OECD Analytical House Price Indicators	none
I	Interest rate: Money market interest rate, 3-month interbank rates	OECD Monthly Monetary and Fiscal Statistics	Finland (IMF Interntional Financial Statistics)
MS	Mortgage stock: Total outstanding mortgage liabilities, deflated	Oxford Economics, Datastream	none
MR	Mortgage rate: interest rate on builing society mortgages	Oxford Economics, Datastream	Germany (Federal Statistics Office); Netherlands (consumer loans)

Table 3: Data sources for each variable and country

A.4 Stationarity

	Germany	Spain	Finland	France	UK	Ireland	Italy	Netherlands
PI	no	no	no	stationary	no	no	no	stationary
CONS	no	no	no	no	no	no	no	no
Y	no	no	no	no	no	no	no	no
HP	no	no	no	stationary	no	no	stationary	no
I	stationary	no	no	no	no	stationary	no	stationary
MS	no	no	stationary	no	no	stationary	no	no
MR	stationary	no	no	no	no	no	no	no

Table 4: Integration of variables in levels for each country

	Germany	Spain	Finland	France	UK	Ireland	Italy	Netherlands
PI	stationary	no	stationary	stationary	stationary	stationary	no	stationary
CONS	stationary	stationary	stationary	stationary	stationary	stationary	stationary	stationary
Y	stationary	stationary	stationary	stationary	stationary	stationary	stationary	stationary
HP	no	stationary	stationary	no	stationary	no	no	stationary
I	stationary	stationary	stationary	stationary	stationary	stationary	stationary	stationary
MS	stationary	no	no	stationary	no	no	stationary	no
MR	stationary	stationary	stationary	stationary	stationary	stationary	stationary	stationary

Table 5: Integration of variables in differences for each country

A.5 Serial correlation of residuals

	Baseline	p-value	Consumption	p-value	Mortgage	p-value
DEU	serial correlation	0.025	serial correlation	0.017	no	0.245
ESP	serial correlation	0	serial correlation	0	serial correlation	0
FIN	serial correlation	0	serial correlation	0	serial correlation	0
FRA	serial correlation	0.002	serial correlation	0.017	no	0.182
GBR	serial correlation	0.008	no	0.065	serial correlation	0
IRL	no	0.19	no	0.184	serial correlation	0
ITA	serial correlation	0.001	serial correlation	0	serial correlation	0
NLD	no	0.095	serial correlation	0.013	serial correlation	0.027

Table 6: Serial correlation of errors for levels specification

	Baseline	p-value	Consumption	p-value	Mortgage	p-value
DEU	no	0.376	no	0.079	no	0.975
ESP	serial correlation	0.001	serial correlation	0.019	serial correlation	0
FIN	serial correlation	0.017	serial correlation	0.002	serial correlation	0
FRA	serial correlation	0.001	no	0.055	serial correlation	0.007
GBR	serial correlation	0.015	no	0.125	serial correlation	0
IRL	no	0.709	no	0.738	serial correlation	0
ITA	serial correlation	0.017	serial correlation	0.004	serial correlation	0.027
NLD	no	0.085	no	0.076	serial correlation	0.042

Table 7: Serial correlation of errors for differences specification

	Baseline	p-value	Consumption	p-value	Mortgage	p-value
DEU	serial correlation	0.005	serial correlation	0.008	serial correlation	0.006
ESP	no	0.089	no	0.076	serial correlation	0.018
FIN	serial correlation	0.023	no	0.107	serial correlation	0
FRA	no	0.638	no	0.803	-	-
GBR	no	0.18	no	0.222	serial correlation	0
IRL	serial correlation	0.012	serial correlation	0.01	-	-
ITA	no	0.356	no	0.063	serial correlation	0.024
NLD	serial correlation	0	serial correlation	0.005	-	-

Table 8: Serial correlation of errors for levels specification using the reduced sample (1980-1998)

	Baseline	p-value	Consumption	p-value	Mortgage	p-value
DEU	no	0.177	serial correlation	0.026	serial correlation	0.034
ESP	no	0.392	no	0.521	no	0.055
FIN	no	0.204	no	0.164	no	0.114
FRA	no	0.73	no	0.955	-	-
GBR	no	0.266	no	0.378	no	0.151
IRL	no	0.361	no	0.1	-	-
ITA	no	0.566	no	0.354	no	0.797
NLD	serial correlation	0.027	serial correlation	0.03	-	-

Table 9: Serial correlation of errors for differences specification using the reduced sample (1980-1998)

	No bubble sample	p-value	Dummy augmented	p-value
ESP	no	0.058	serial correlation	0.001
FIN	no	0.453	no	0.051
FRA	no	0.262	serial correlation	0.001
GBR	no	0.187	serial correlation	0.017
IRL	no	0.531	no	0.634
ITA	no	0.391	serial correlation	0.016
NLD	no	0.178	no	0.062

Table 10: Serial correlation of errors for the reduced sample specification excluding bubble periods and the step dummy specification