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The Heterogeneous Effects of Monetary Policy on Swedish House Prices:

A regional study in times of differing house price growth

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ABSTRACT:

This paper analyzes the impacts of monetary policy on house prices on a national and regional level in Sweden employing a structural vector autoregressive (structural VAR) model. We find that after an initial increase, real house prices turn significantly negative to bottom out at -0.5% after three quarters, and stay negative for another five quarters. On a disaggregated level we find that house prices respond heterogeneously to a monetary policy shock across regions in Sweden. While on a national level monetary policy does not seem to have a meaningful impact on house prices when distinguishing for times of booms, the regional examination uncovers a significant drop for most regions during local boom periods.

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List of abbreviations

BIS	Bank for International Settlement
CPI	Consumer Price Index
ECB	European Central Bank
HP	Hodrick-Prescott
IMF	International Monetary Fund
IRF	Impulse response function
LTV	Loan-to-value
MWW test	Mann-Whitney-Wilcoxon test
OLS regression	Ordinary least squares regression
OECD	Organisation for Economic Co-operation and Development
RGDP	Regional gross domestic product
REER	Real Effective Exchange Rate
SCB	Statistiska centralbyrån
SEK	Swedish Krona
STIBOR	Stockholm Interbank Offered Rate
VAR	Vector autoregressive

1. Introduction

The recent global financial crisis, sparked by the expansion of subprime lending and securitization, was in many countries associated with a rapid decline in house prices. This has, especially among central banks, renewed the interest in the role of asset prices as a contributor to the build-up of financial imbalances and how these might be limited (Björnland & Jacobsen, 2010; Galí & Gambetti, 2014; Shi;Jou;& Tripe, 2014). It is generally recognized that asset prices can have a considerable influence on both inflation and real economic activity – in other words variations in asset prices are an important factor of financial and macroeconomic stability (see e.g. Cechetti, Genberg, Lipsky, & Wadhwani, 2000; Bernanke & Gertler, 2000). However, the current literature has no clear consensus on the quantitative effect of monetary policy on asset prices (Robstad, 2014; Shi, Jou, & Tripe, 2014). In the prevailing monetary policy framework, central banks focus primarily on price stability and the overall economic development, mainly using the instrument of short-term interest rates. It is a still ongoing debate whether monetary policy should respond explicitly to asset price variability or if rather additional, separate macroprudential tools should be employed to guard against financial instabilities (Galí & Gambetti, 2014; Smets, 2013; Björnland & Jacobsen, 2010; Shi, Jou, & Tripe, 2014).

Especially asset price booms in the housing market are seen as a source for macroeconomic and financial instability (Crowe, Dell'Ariccia, Igan, & Rabanal, 2013). Real estate is the most important asset for households in industrialized countries and thereby acts both as a storage of wealth and as a consumption good. Households can use housing as collateral when borrowing for further consumption. This implies that if house prices rise during a housing boom, households can borrow on these increased values and can thus consume more. This in turn could exacerbate general economic booms and thus induce macroeconomic instability. Furthermore, mortgages usually constitute a substantial fraction of household debt. During a house price boom households often take on higher mortgages to afford the increased housing prices. As a consequence these overly levered households might have difficulties in meeting their obligations, which in turn can directly impact banks' solvency and their ability to supply credit, thereby further impacting the economy in total. (Crowe, Dell'Ariccia, Igan, & Rabanal, 2013; Elbourne, 2008)

The housing market is also one of the sectors most sensitive to monetary policy (Dokko, Doyle, Kiley, & Kim, 2011). Since interest payments represent a substantial part of the cost of buying a house, the demand and hence also prices for housing are conventionally assumed to be negatively related to interest rates. The amount someone is willing and able to pay for a house depends on the mortgage interest rates and borrowing opportunities (Englund, 2011; Elbourne, 2008). Even though an increase of the repo interest rate is thus associated with decreasing house prices, in the short run an increase of the interest rate might as well lead to the opposite effect. This can be attributed to two factors. First, expectations of further interest rate increases in the future might drive up the demand for housing in the short-run. Second, loss aversion or loan-to-value constraints of potential house sellers might decrease the supply of housing in a downturn (McCarthy & Peach, 2002).

Real house prices in Sweden have on average been increasing by about 5% annually between 1996 and 2014 (SCB, 2015). The rise in house prices has mainly been financed by increasing household leverage, leading to a record high of the average household debt-to-income ratio in 2014 from a historical perspective (Sveriges Riksbank, 2014). Sweden has been one of the first countries to explicitly use monetary policy as a tool to counteract rapid asset price increases in the housing market. Rising household indebtedness together with increasing house prices were in 2012 explicitly cited as reasons to keep interest rates higher than would have been justified by solely considering inflation (Crowe, Dell'Ariccia, Igan, & Rabanal, 2013; Svensson, 2013; Ekholm, 2014). With this approach the Swedish Riksbank chose a “leaning against the wind” track of monetary policy, meaning that financial stability is included as a secondary objective for the central bank. According to this track, price stability and financial stability are thus closely intertwined and financial stability concerns are directly incorporated in the monetary policy framework when deciding for the optimal adjustment paths for inflation (Smets, 2013). Interestingly, while both house prices and household indebtedness in Sweden continue to grow, it seems as if the latest interest rate decisions do not reflect this stance of monetary policy anymore – the repo rate has been decreased considerably since the second half of 2014 (Sveriges Riksbank, press release 2015; Sveriges Riksbank, 2015b). This might also be related to the shift in responsibility for financial stability in Sweden, which in 2014 has been moved from the Riksbank to the financial supervisory authority, the Finansinspektion (SEB, 2014; Finansdepartementet, 2014). The developments in the Swedish housing market in combination with the Riksbank’s policy stance of having explicitly addressed these in monetary policy decisions, make it particularly interesting and relevant to study the actual effects of monetary policy on house prices in Sweden.

This paper will first examine the impact of monetary policy on house prices in Sweden using a structural vector autoregressive model (structural VAR), an approach which is commonly employed by previous literature in this context (Björnland & Jacobsen, 2010). The analyses include data from 1996 Q1 to 2014 Q2. We will not only consider the effects on an aggregated, national level, as was done for Sweden by Björnland & Jacobsen (2010) by using data from 1983 Q1 to 2006 Q4, but also examine the differences across regions. This is motivated by the fact that housing markets usually are determined locally. Previous research on the US has found that a monetary policy shock can have differing impacts on house prices across regions depending on the respective underlying economic conditions (Fratantoni & Schuh, 2003; Vargas-Silva, 2008). Consequently, as Goodman (1998) shows, results from geographically aggregated data might lead to puzzling findings or even to biased estimates. In our study we will solely focus on the effects of monetary policy in Sweden, since country specific factors, such as taxes, housing construction flexibility or mortgage market regulations, are likely to lead to considerable differences across countries (Assenmacher-Wesche & Gerlach, 2008).

Second, the paper will examine if house prices react differently to monetary policy shocks during times of house price booms versus periods not characterized by a boom. Previous research, for example Goodhart & Hofmann (2008), find a more pronounced impact of monetary policy on house prices during times of house price booms as compared to times not characterized by booms. This poses the question if the same effect can also be found in Sweden, i.e. whether monetary policy decisions during house price booms become even more relevant for the Riksbank. For this analysis boom periods in the Swedish housing market on a national level and for the different regions will be identified and the impact of monetary policy will be measured using a structural VAR respectively. To the best of our knowledge the effects of monetary policy in Sweden have not been considered on a disaggregated, regional level, and also not regarding differences in boom versus non-boom periods.

Overall, our results suggest that monetary policy shocks affect house prices in Sweden negatively; on an aggregated, national level, house prices bottom out at 0.5% below the baseline after an initial increase as a reaction to a 1% contractionary monetary policy shock. On a regional level we can confirm the heterogeneous responses of house prices to a monetary policy shock as it has been found in the US. Looking at the effects of monetary policy on house prices on a national level in boom versus non-boom periods gives no clear picture. Responses seem to be insignificant during boom periods, thereby questioning the effectiveness of monetary policy as a tool for mitigating financial instability. On a regional level, on the other hand, we find strong and significantly negative effects of monetary policy on house prices during house price booms and only little reactions during non-boom periods. The differing effects might be explained by the way households interpret the increase in the short-term interest rate. In non-boom times it might be seen as a signal that the central bank expects further economic expansion, which would open up for further interest rate increases; in boom periods the interest rate increase might rather be interpreted a warning signal of overheating.

The remainder of the paper is structured as follows: Section 2 provides an overview of the existing literature. Section 3 illustrates the background and recent developments of Swedish monetary policy and the Swedish housing market. Section 4 outlines the methodology and the model used in this paper, and Section 5 presents the data. The results are presented in Section 6. Lastly, Section 7 concludes.

2. Literature review, monetary policy and house prices

Numerous studies have examined the impact of monetary policy on house prices, often across several countries but also across the different regional housing markets within one country. Most of the research is conducted using data from the Euro area or the US; only few studies focus on the dynamics of single countries, especially when looking at small, open economies. Since country-specific factors such as policies regarding taxation, mortgage market regulation, transaction costs, and housing market flexibility considerably influence the interaction between monetary policy and house prices, it is common that studies exhibit large variations in their results. However, it is mostly found that a contractionary policy shock impacts house prices negatively. Even though there are different approaches to assessing the impacts of monetary policy shocks on house prices, the effects are most commonly analyzed using a vector autoregressive model (VAR).

The literature on the transmission mechanisms of monetary policy on house prices focusing on Sweden in particular or on other small, open economies in general is very limited. The only paper we have found specifically analyzing the link between monetary policy and house prices in Sweden using a VAR methodology is the paper by Björnland & Jacobsen (2010). The authors analyze the role of house prices in the monetary policy transmission mechanism on a national level in Norway, Sweden, and the UK using structural VARs and quarterly data from 1983 to 2006. The response of house prices to a 1% contractionary monetary policy shock is found to be approximately -3% for Norway, -4% for Sweden and -6% for the UK. By allowing the interest rates and house prices to react simultaneously on news, the authors find that the role of house prices in the monetary policy transmission mechanism increases considerably, compared to the case without simultaneous interaction. Björnland & Jacobsen (2010) conclude that monetary policy has a strong and prolonged impact on house prices. Laseén & Strid (2013) also conduct a structural VAR to examine the effects of monetary policy in Sweden, using quarterly data from 1995 to 2013. However, Laseén and Strid's focus is on the relationship between monetary policy and debt levels in Sweden. The estimated impact of monetary policy on house prices is a side-product of their analysis, and indicates that a 1% interest rate shock corresponds approximately to a 1% decrease in house prices, which is considerably lower than what Björnland & Jacobsen (2010) found. Robstad (2014) examines the impact of a monetary policy shock on house prices and household credit in Norway between 1994 and 2013 and finds similar results as Björnland & Jacobsen (2010) do for the impact of monetary policy on house prices in Norway. Hence, also Robstad (2014) concludes that monetary policy has a fairly pronounced impact on house prices while the effect on household credit is muted.

Further studies focusing on the effects of monetary policy on asset prices in single countries examine usually larger open economies, such as the US or the UK. Elbourne (2008) conducts a structural VAR analysis in the UK for a monthly sample from January 1987 to May 2003. He finds that a 1% positive interest rate shock leads to a house price fall of 0.75%. That a monetary policy shock affects house prices negatively is in line with the above specified studies, however, the estimated effect is considerably lower than the effect found by Björnland & Jacobsen (2010) for the UK, presumably caused by differing model specifications. Elbourne (2008) identifies furthermore that the fall in house prices can explain up to 15% of the fall of consumption following an interest rate shock. Jarocinski and Smets (2008) review the role of monetary policy and the housing market in the US between 1987 and 2007 by employing a Bayesian VAR model and find that house prices bottom out at 0.5% below the baseline as a reaction to a 0.25% tightening in monetary policy. Additionally, the authors see indications that the unusually low level of short and long term interest rates might have contributed to the boom in the US housing markets. A similar structural VAR analysis on the US housing market is also done by McCarthy & Peach (2002) but in contrast to Jarocinski & Smets (2008) the authors find, using quarterly data between 1986 and 2000, that the initial reaction of house prices to an interest rate increase is positive; prices do not turn negative before the third quarter after the shock. McCarthy & Peach (2002) refer to the initial positive response as the "home prize puzzle" and argue that this might reflect two factors. First, potential house sellers might be reluctant to realize losses in a downturn due to loss aversion or due to loan-to-value (LTV) constraints, thereby reducing the supply of housing. Evidence of this has for example been found by Genesove & Mayer (2001) and Genesove & Mayer (1997). Second, potential house buyers might fear further interest rate increases in the future and thus prefer buying a house rather now than waiting, since financing might become even more expensive. In this way the demand in the housing market is sustained at least for some time after a monetary tightening (McCarthy & Peach, 2002).

Country-specific differences regarding, for example, the regulation and taxation of transactions or the mortgage market, imply that cross-country studies examining the relation between monetary policy and house prices can lead to uncertain results exhibiting large variations (Assenmacher-Wesche & Gerlach, 2008). However, the comparison of monetary policy effects on house prices across countries has been a common research design and results usually confirm the negative effect of monetary policy on house prices. Goodhart & Hofmann (2008) for example study the dynamic relations between money, credit, house prices and economic activity in the period from 1970-2006 in 17 industrialized countries using a fixed-effects panel VAR. The authors find a significant link between house prices, monetary variables, and the macro economy. An increase in nominal interest rates leads to a decrease in house prices. The results reveal that the link between house prices and monetary variables is stronger in the more recent sub-samples from 1985-2006. Additionally, when house prices are booming, the effects of shocks to money and credit are stronger than in non-boom periods. Musso, Neri, & Stracca (2010) study whether monetary policy shocks have differing impacts in the US versus the Euro area using data from 1986-2008 in a structural VAR. Similarly to the findings of Goodhart & Hofmann (2008), also Musso, Neri, & Stracca (2010) find that positive monetary policy shocks have negative effects on house prices both in the US and in the Euro area; however, the impact is greater in the US housing market than in the Euro area.

Not all papers agree that monetary policy has a strong impact on house prices. For example Miles (2014) examines the effects of the short term Fed rate on house prices in the US compared to long term interest or mortgage rates using a reduced form VAR. Miles finds that the Fed rate's influence on the long-term interest rates has been decreasing over time and hence concludes that the loose monetary policy in the US might not have been crucial for the recent house price boom and bust. Similarly, also Glaeser, Gottlieb, & Gyourko (2010) question the importance of interest rates on house prices, stating that lower interest rates can only explain one-fifth of the rise in house prices from 1996-2006.

Another recent study by Shi, Jou, and Trip (2014) investigates how monetary policy impacts house prices in New Zealand both on a national and on a regional level between 1999 and 2009. Instead of a VAR model the study uses the present value model to determine how the real rental rate and the real interest rate affect real house prices. In contrast to other papers, Shi, Jou, and Trip (2014) find that real interest rates are significantly related to real housing prices in a positive way, suggesting that increasing interest rates might not be an effective tool to reduce house prices.

Heterogeneous effects of monetary policy shocks

As monetary policy is not able to target or control region specific conditions, the conventional view is that it should focus solely on nationwide economic conditions. At the same time it is acknowledged that regional economic conditions influence how the regions respond to monetary policy actions, both due to regional sensitivities and the initial regional economic conditions at the time of the monetary policy shock (see e.g. Carlino & DeFina, 1999; Fratantoni & Schuh, 2003; Owyang & Wall, 2003 and Vargas-Silva, 2008). Further, for example Goodman (1998) states that as local housing markets almost certainly are non-linear, i.e. the marginal effect of a shock differs across regions, geographical aggregation of data might result in biased estimates. Even if Goodman (1998) admits that aggregate housing market estimation is far preferable over no estimation at all in the lack of disaggregated data, he strongly advocates the use of regional modeling to avoid an aggregation bias. Many of the studies on the heterogeneous effects of monetary policy study the impact of an interest rate shock on income, GDP, or the price level. For example Carlino & DeFina (1999) and Owyang & Wall (2003) test for heterogeneous effects of monetary policy shocks on regional income in the USA, Fielding & Shields (2006) study regional differences of the monetary transmission mechanism on price levels in South Africa, and Arnold & Vrugt (2002) examine the impact of monetary policy shocks on regional output in the Netherlands. Francis, Owyan, & Sekhposyan (2012) take the level of disaggregation one step further and estimate the responses of employment to a monetary policy shock on city-level in the US using a structural VAR. All authors find evidence that monetary policy shocks have different effects across regions.

Fratantoni & Schuh (2003), Vargas-Silva (2008), and Campbell, Yang, & Wang (2010) test for heterogeneous effects of monetary policy shocks specifically on the housing market. Fratantoni & Schuh (2003) argue that since both housing supply and demand are determined in distinct local markets, an analysis of the impact of monetary policy on house prices should be conducted using data at a disaggregated level. The authors employ data from 27 US regions between 1986 and 1996 in a modified panel VAR, and find that regional conditions indeed significantly influence the impact of monetary policy on house prices. Further the authors argue that housing is a good candidate to quantify the regional responses to monetary policy as it by its nature is a critical channel of monetary transmission.

Similar results of heterogeneous effects across regions are also found by Vargas-Silva (2008), who examine the impulse responses of house prices to a monetary policy shock on a regional level in the US employing a structural VAR identified by sign restrictions. Specifically, the results show that while the initial response to a contractionary monetary policy shock is negative across all regions, both the magnitude of the trough and the persistency of the impacts varies across regions. Campbell, Yang, & Wang (2010) investigate the heterogeneous effects of monetary policy on 20 regional units in Sweden between 1991 and 2002. The model they use is a multivariate persistent shock metric, where shocks are decomposed into an interest rate shock and a regional-specific shock component. Campbell, Yang, & Wang (2010) find significant regional effects of monetary policy on the regional housing markets, however, the results indicate that interest rates dominate the influence of local price innovations especially in the core economic regions in Sweden. Further, the authors claim that monetary policy has contributed significantly to the house price booms in the three biggest cities of Sweden, i.e. Stockholm, Gothenburg and Malmö. However, since no economic factors or micro information are included in this model, further analysis needs to be conducted controlling for these in order to see whether the results hold in a more detailed model.

Monetary policy and asset price booms

As touched upon in earlier sections, research has found that the effects of a monetary policy shock on asset prices are larger when asset prices are experiencing a boom (Crowe;Dell'Ariccia;Igan;& Rabanal, 2013; Goodhart & Hofmann, 2008). Usually, booms are defined as major, long-term deviations from a price trend. The underlying price trend is mostly estimated recursively using information available up to the time when the prediction was made. To capture low frequency, cumulative deviations, implicitly emphasizing the mean reversion tendencies, high smoothing parameters are used. Further, many papers include a period of approximately 10 years before the actual time frame to be able to observe meaningful trends (see e.g. Goodhart & Hofmann, 2008; Borio & Lowe, 2004; Adalid & Detken, 2007).

The exact definition and identification of a boom period varies across papers. Adalid & Detken (2007) define asset booms as a consecutive period of at least four quarters, when the current asset price index exceeds the estimated trend by at least 10%. Goodhart & Hofmann (2008), define a boom as a positive deviation of more than 5% from the smooth trend lasting for at least 12 quarters. In order to test for the differing impacts on monetary policy on house prices, the authors include a dummy variable in their panel-VAR model, which indicates if there is a house market boom or not. The authors identify house price booms for 18 countries between 1985 and 2006 and estimate for Sweden boom periods from 1987 Q4–1991 Q4 and 1998 Q1–2006 Q4. Agnello & Schuknecht (2009) refine the identification of boom and bust periods applying a so-called triangular methodology on the housing market in 18 countries, including Sweden, from 1980-2007. The authors start by de-trending the series to account for long term trends in a similar way as Borio & Lowe (2004), Adalid & Detken (2007), and Goodhart & Hofmann (2008) do, but instead of estimating the trends recursively, the whole sample, i.e. ex-post data is employed. Local peaks and local troughs define a phase, for which the duration and the magnitude is calculated. A phase is then classified as a boom or a bust if the cumulative change corresponds to the first quartile from the empirical distribution of the cumulative changes in the whole sample. According to this paper, Sweden experienced a housing boom at a comparable period as in Goodhart & Hofmann (2008), from 1986-1990 and from 1997-2007.

Englund (2011) has developed a fundamental valuation approach applicable for the Swedish housing market to estimate when house prices are "overvalued" or "undervalued". The core assumption of Englund's (2011) model is that the yield of investing in housing, approximated by rent-to-price, should in equilibrium be equal to the user cost of housing. In well-functioning rental markets where housing offerings are a good substitute for owner-occupied housing, the cost of rental housing is commonly seen as a natural benchmark for the price of owner-occupied housing. Hence, if owner-occupied houses are rationally priced, the cost of housing consumption should be fairly equal for both owner-occupied and rental housing, i.e. the discounted cost of owning a house (including the initial investment, interest payments, and cost of maintenance and taxes) and the discounted value of rents (assumed to grow at a constant rate and setting the discount rate is equal to the after tax interest rate on mortgage loans) should be equal. When the rent-to-price ratio is higher than the user cost of housing, the housing market is "undervalued" and when the user cost of housing is higher than the rent-to-price ratio the market is "overvalued". To account for the regulated rental market in Sweden, Englund (2011) assumes that the regulated, observable rents and the real value of housing in a rental setting remain constant over time. This implies that using the available rent index will only result in a measure of the rent-to-price that differs from the "true" measure by a constant. Englund (2011) concludes that even though fairly bold assumptions are necessary, the time variation of the two time series reflect qualitatively useful information of the broad patterns in the Swedish housing market. Using data between 1980 Q1 and 2010 Q3, he estimates that the housing market was overvalued from 2004 to 2010 Q3.

Speculations about whether there is a bubble in a housing market on either a national or a regional level is a hotly debated topic. It is difficult to determine when a boom turns into a bubble, and in fact, the term “bubble” is even rarely clearly defined (Case & Shiller, 2003). A broad definition of an economic bubble is if there is a difference between the current market value and the fundamental value (see e.g. Flood & Hodrick, 1990; Dillén & Sellin, 2003). Case & Shiller (2003), looking specifically at the housing market, further argue that in the widespread use “bubble” refers to excessive public expectations of future price increases, which then cause a temporary rise in prices; however, a rapid price increase does not automatically indicate the existence of a bubble (Case & Shiller, 2003).

3. Background

The first part of this section will briefly introduce the objectives of the Swedish central bank, the Riksbank, as well as discuss the recent developments in the Swedish monetary policy. The second part presents stylized facts about the Swedish housing market.

3.2 Monetary policy in Sweden - the Riksbank

The Swedish central bank, the Riksbank, was founded in 1668 and is considered to be the oldest central bank in the world, and has since 1999 had an independent status in the Swedish law (Sveriges Riksbank, 2011). The statutory objective of the Riksbank is to maintain price stability. However, without neglecting this overriding objective, monetary policy also aims to support the general economic policy to stabilize production and employment around long-run sustainable paths (Sveriges Riksbank, 2015a). This is commonly referred to as “flexible inflation targeting (Sveriges Riksbank, 2015a; Hallsten & Tägström, 2009).

The objective of price stability was introduced in 1993 in the aftermath of the turbulent economic period in 1992 when the Riksbank was forced to abandon the fixed exchange rate against the predecessor of the Euro. The current inflation target, defined as the annual rise in the consumer price index (CPI), is set to 2%, and became effective in 1995 (Sveriges Riksbank, 2015a; Sveriges Riksbank, 2011; Ingves, 2007). Interest rate decisions are taken six times a year by the independent Executive Board of the Riksbank consisting of six full-time members (Hallsten & Tägström, 2009). Financial stability is mainly fostered through information gathering and analysis, resulting in recommendations aiming to encourage agents to reduce their risk-taking in the financial sector as well as strengthen their resilience to shocks. It is stressed that monetary policy only acts as a complement for the most important factors for preventing imbalances in asset prices and indebtedness; effective regulation and supervision (Sveriges Riksbank, 2015a; Sveriges Riksbank, 2014b). In 2014 the main responsibility for financial stability in Sweden was, after several investigations and discussions, ultimately moved from the Riksbank to the financial supervisory authority, the Finansinspektion. However, there seems to be a wedge between how these two authorities perceive the need for stabilizing policies; the Finansinspektion regards the recovery of the economy to be too frail for introducing macroprudential policies, such as amortization requirements, whereas the Riksbank strongly recommends the introduction of such measures (SEB, 2014; Finansdepartementet, 2014).

From 2004 onwards, the Executive Board got increasingly concerned about the level of household debt and the rapidly growing house prices in Sweden, (Ingves, 2007; Crowe, Dell'Ariccia, Igan, & Rabanal, 2013; Svensson, 2013; Ekholm, 2014). Even though rising household indebtedness driven by increasing house prices was not explicitly mentioned as a reason for repo-rate decisions until October 2012, for example (Ekholm, 2014) claims that the repo rate had been set higher than was justified by strictly looking at inflation already for several years. An example of a “leaning against the wind” monetary policy, in which asset prices and household indebtedness are explicitly taken into account to counteract financial instabilities, could also be seen in 2013, when the repo rate was kept constant in spite of falling prospects for inflation (Ekholm, 2014). However, whether tighter monetary policy actually is a viable tool to curb household debt has recently been questioned by for example Svensson (2013) and Ekholm (2014). Svensson (2013) argues that tighter monetary policy leads to a fast decrease of the nominal price level and nominal GDP, but only to a relatively slower decrease of total debt. Hence, he argues that “leaning against the wind” as a way to reduce the household debt-to-GDP ratio is counterproductive (Svensson, 2013). Svensson’s (2013) arguments have been questioned by for example Lars Heikensten, governor at the Riksbank from 2003-2005 and Per Jansson, Deputy Governor at the Riksbank (Heikensten, 2014; Jansson, 2014), both arguing that looking at and evaluating “leaning against the wind” from just a debt-to-GDP perspective is too narrow, rather it should be seen and evaluated from a broad economic context comprising the stability of the whole financial system. Nevertheless, even though debt levels and house prices continue to increase, the Riksbank has recently lowered the interest rate considerably, indicating that currently, at least temporarily, “leaning against the wind” is off the table

(Sveriges Riksbank, 2014). This is especially interesting as the Riksbank at the same time states that “The low interest rates are also expected to lead to continued rise in housing prices and rising household debt” (Sveriges Riksbank, 2014).

3.1 The Swedish housing market

Real house prices in Sweden have increased trend-wise over the last 20 years with a compounded annual growth rate of about 5% (SCB, 2015). The main underlying reasons for this trend are related to the rise of urbanization, population growth as well as increasing income levels. In general the housing market is characterized by long cycles and well-defined peaks and troughs. Looking at Sweden the housing market experienced peaks in in 1979 and 1989-1991 and troughs in 1985-1986 and 1993-1996. (Englund, 2011). The current Swedish house price increase can be considered to have started in 1997, and in contrast to many other developed countries, Sweden did not experience a severe slowdown in the housing market in the recent financial crisis (Englund, 2011). Research from Agnello & Schuknecht (2009), comparing the housing markets in 18 OECD countries until 2007 even classifies the Swedish housing market boom as the longest and most pronounced of the sample. After 2007 Swedish house prices have continued to grow at an accelerating rate (Sveriges Riksbank, 2014), and especially noteworthy is that the continuing rise of housing prices is acknowledged to be at least partly fostered by the current low interest rate environment (Sveriges Riksbank, 2014). Other culprits that are mentioned for the continuing rise in house prices in Sweden are the low property taxation and the tax deductibility of mortgage interest payments. These two factors create a strong tax incentive for households to invest in housing rather than in other assets (IMF, 2014).

Although house prices are positively correlated across regions, urbanization drives a wedge between the house price developments within the country as the increasing number of high income households moving to cities increases the pressure of centrally located land. The role of land scarcity becomes especially evident when looking at price differences across regions in Sweden (see also Appendix 1, Real house price indices, Sweden total and all regions 1996 Q1-2014Q2). Comparing the price development of for example the three major metropolitan areas in Sweden, Stockholm, Malmö, and Gothenburg, with rural, more sparsely populated areas, reveals that whereas the price level in the three urban areas have increased around two and a half times from the levels in the early 1980’s, prices in the rural areas have hardly increased at all. It is also found that the amplitude of the cycles tend to be higher in expanding regions (Englund, 2011).

The Swedish rental market is highly regulated and official rents are set to “fair value” in negotiations with central organizations representing landlords and tenants. Consequently, rents in central locations in major cities are set at a fair value that is significantly below market rents. This implies that in many cases there is a difference in the official rent and the value of housing in a rental setting (Englund, 2011). The rental market in Sweden is further characterized by a lack of supply of rental housing especially in areas with high demand. Additionally, the willingness to invest in the construction of rental housing is poor due to the market conditions with uncertain profit calculations. Moreover, especially in Stockholm there is a trend to convert rental buildings into owner-occupied housing (Statens offentliga utredningar, 2012). The consequences of the weak supply of rental housing are long queueing times and a substantial black market for rental contracts. It has been argued that the lack of a well-functioning rental market contributes to an increased demand for owner-occupied housing, thus further pushing up the price level also increasing household indebtedness (IMF, 2014; European Commission, 2014). Even if some measures have been taken in the rental market, for example to make it easier for private individuals to sublet their apartments, these measures are seen as insufficient to address the underlying structural problems (European Commission, 2014).

A closely related concern to the increasing house prices in Sweden is the high household indebtedness. Swedish household debt has grown substantially in recent years and is now from both a historical and an international perspective high (Sveriges Riksbank, 2014; Winstrand & Ölcer, 2014). Mortgages constitute 95% of the total debt held by households, indicating that the booming housing market is also reflected in the high debt burden (Sveriges Riksbank, 2014; IMF, 2014). About 40% of households hold interest-only loans, which is considerably more than abroad, for example in France the proportion is 0.33%. Furthermore, 42% of Swedish households have not drawn up any amortization plan, and even those households who amortize do that at a very slow rate. According to a recent survey conducted by the Riksbank comprising 1.8 million Swedish households, paying off the mortgage would take at least 85 years for approximately half of the new mortgage takers. In contrast, in many other countries it is common for households to amortize the full mortgage in 20-40 years. Many households additionally choose variable mortgage rates, and for example during the fall of 2014, 73% of the new mortgage loans have been subscribed at a variable rate. (Sveriges Riksbank, 2014).

Despite efforts to reduce household indebtedness, for example through the introduction of a loan-to-value cap in 2010, household debt has continued to increase and currently mortgages are estimated to increase at an annual rate of approximately 6%, a faster increase than the increase in both GDP and the households' disposable income (Sveriges Riksbank, 2014). Concerns about the high household indebtedness and increasing house prices in Sweden have not only been raised on a national level; also several international institutions, for example the IMF, the OECD, and the European Commission, have expressed the need to take actions to dampen the growth of indebtedness and house prices in Sweden (IMF, 2014; Sveriges Riksbank, 2014). In November 2014 the Riksbank presented a new recommendation addressing the increasing household indebtedness and the growing house prices; all "new" mortgages should be amortized down to 50% of the value of the house. Since the recommendation is vague it is however not clear what effects exactly this suggestion will have. Additionally, addressing the amortization requirements alone is estimated not to be enough to curb the rise in house prices and household indebtedness (Sveriges Riksbank, 2014). The IMF (2014) is on the same track, suggesting several macro prudential policy measures to address the situation in the Swedish housing market. These include for example an introduction of a loan-to-value and debt-to-income cap, a reduction of interest rate deductibility and a tightening in the property taxation (IMF, 2014).

4. Methodology

In order to analyze the interdependence of macro variables, here specifically of monetary policy and house prices, a structural vector autoregressive (VAR) model is employed, thereby following e.g. Björnland & Jacobsen (2010), Robstad (2014), Musso, Neri, & Stracca (2010), Elbourne (2008), and many others. Structural VARs are one of the pillars in empirical macroeconomics and finance, and allow to study responses of variables to one-time shocks (Kilian, 2011). Especially for analyzing the effect of monetary policy on economic variables a structural VAR approach has usually been the common procedure (Björnland & Jacobsen, 2010). All analyses are done in MATLAB using the Oxford MFE Toolbox by Sheppard (2013), and the MATLAB algorithm provided by Binning (2013/14). The methodology section is structured in the following way: First, we introduce vector autoregressions generally; in the second part, the structural VAR model for Sweden on a national, and afterwards on a regional level is established. Lastly, boom and non-boom periods are identified on a national and a regional level and a dummy-extended structural VAR is introduced.

4.1 Vector autoregressive models

Vector autoregressive models in general, as introduced by Sims (1980), are multi-variable linear models in which each variable in turn is explained by its own lagged values, and current and past values of the other variables in the system. All variables are hence endogenous and the often difficult distinction between endogenous and exogenous variables does not need to be made. In a reduced form VAR every variable is estimated by an ordinary least squares (OLS) regression as a linear function of its own and the other variables' past values, and a serially uncorrelated error term. If the variables within the system are correlated, which is usually the case, also the error terms in the reduced form model will be correlated across the equations (Stock & Watson, March 2001). Due to the complexity of the VAR system it is very difficult to interpret the coefficients meaningfully; hence, rather the impact of shocks from one variable on the system and all model variables is examined in impulse response functions (IRF) and by forecast error variance decompositions, which shows how much of a change in a dependent variable is due to its own shock or rather attributable to other shocks (Brooks, 2008). However, only if the correlated reduced form VAR forecast errors are decomposed into structural shocks that are mutually uncorrelated and have an economic interpretation, the causal effects of these shocks on model variables can be assessed (Kilian, 2011). As the reduced form representation does not give enough information to pin down the structural parameters, econometricians need to impose identifying restrictions on how structural shocks impact variables within the model system, thereby transforming the reduced form VAR into a structural VAR model (Binning, 2013/14). Often restrictions on contemporaneous or short-run effects among variables are imposed, but also restrictions on long-run effects or imposing sign restrictions are common approaches (Kilian, 2011). Estimated results will be sensitive to the underlying assumptions, and hence it is advisable to use several identification schemes to ensure the robustness of the results (Rubio-Ramirez, Waggoner, & Zha, 2010).

4.2 Methodology for Sweden on a national level

Our estimations for the impact of monetary policy on real house prices in Sweden on a national level are based on a reduced form VAR of the following form:

$$y_t = C_0 + A_1 y_{t-1} + \cdots + A_l y_{t-l} + u_t \qquad (1.1)$$

where y_t is a vector of endogenous variables, C_0 is a constant, A_l are the coefficient matrices on the lags l, and u_t is a vector of error terms at time t. The variance-covariance matrix of the residuals is given by $E(u_t u_t') = \Sigma_u$ (see e.g. Kilian, 2011; Robstad, 2014; Binning, 2013/14;

Rubio-Ramirez, Waggoner, & Zha, 2010). Following Björnland & Jacobsen (2010) in aiming to capture the features of the dynamics in a small and open economy, we include real house prices (hp_t), the 3-month nominal domestic interest rate (i_t), a 3-month trade weighted nominal foreign interest rate (i_t^*), the real effective exchange rate (e_t), real GDP (g_t), and inflation (π_t) as variables in the model. All series are of quarterly frequency and in log differences; only the interest rates are in levels. Inflation is measured as the annual changes of the CPI because annual inflation is a more direct measure of the target rate mattering for policy makers than quarterly changes (Björnland & Jacobsen, 2010). Additionally, annual changes compared to annualized quarterly changes show less seasonal excess variability in the data implying that results might be more robust (Lindé, 2003). The nominal interest rate is chosen to capture the monetary policy shock, since central banks use interest rate instruments in the monetary policy setting (Björnland & Jacobsen, 2010). All variables are stationary according to an augmented Dickey Fuller test at a 5% level when a constant is employed.

To find the optimal number of lags to be included into the reduced form VAR the AIC, BIC and HQIC information criteria are used, see Appendix 2. Since for both the BIC and the HQIC information criteria a lag order one is optimal, i.e. minimizing the value of the information criterion[3], this lag order is chosen for the model in this paper. When estimating the reduced form VAR a Ljung-Box test of the error terms, i.e. residuals, confirms that the residuals are serially uncorrelated and hence that the model fit is good.

Before identifying a structural VAR based on the reduced form VAR in a next step, we first ensure the existence of a statistically significant link between real house prices and monetary policy using a Granger Causality test (Dokko, Doyle, Kiley, & Kim, 2011; Goodhart & Hofmann, 2008). The relation between the two variables is well documented in previous research (see *Literature review, monetary policy and house prices*) and this can with a p-value of 0.002 also be confirmed for our sample; the p-value indicates that the null hypothesis of monetary policy not influencing real house prices can be rejected at the 5% level (Dokko, Doyle, Kiley, & Kim, 2011). It hence seems appropriate to conduct a structural VAR analysis.

Our six variable reduced form VAR allows the identification of six structural shocks. However, we are primarily interested in the structural shocks to monetary policy (ε_t^i) and hence

[3] Information criteria embody a function of the residual sum of squares and some penalty for the loss of freedom when adding extra parameters; when adding an additional lag the residual sum of squares decreases, but the value of the penalty term increases (see e.g. Brooks (2008) for more information).

follow Björnland & Jacobsen (2010) in identifying the other shocks only loosely as real house price shocks (ε_t^{hp}), as inflation shocks (ε_t^{π}), output shocks (ε_t^{g}), exchanges rate shocks (ε_t^{e}), and foreign interest rate shocks ($\varepsilon_t^{i^*}$). Similar to Robstad (2014), and to ensure robust results, we will use three common identification methods to identify a monetary policy shock based on the reduced form VAR: the Cholesky decomposition of the variance-covariance matrix of the reduced form VAR residuals as in Sims (1980), short- and long-run zero restrictions as used by Björnland & Jacobsen (2010), and sign restrictions comparable to Uhlig (2005). Each method has its benefits and drawbacks, and there is no consensus about which types of identification assumptions are the most viable (Robstad, 2014). The identified structural VAR models have the following form:

$$B_0 y_t = D_0 + B_1 y_{t-1} + \cdots + B_l y_{t-l} + \varepsilon_t \qquad (2.1)$$

where B_0 is the matrix of contemporaneous restrictions, $B_0^{-1}B_l = A_l$, referring to the coefficient matrices of the lags l, and $B_0^{-1}C_0 = D_0$, the constant. ε_t is a vector of structural shocks with $u_t = B_0^{-1}\varepsilon_t$. The variance of the structural shocks ε_t is normalized to unity with $E(\varepsilon_t\varepsilon_t') = I_m$ (where I_m is an identity matrix of order m), so that a unit innovation in the structural shocks is an innovation of the size of one standard deviation. This normalization implies that[4] $B_0^{-1}B_0^{-1\prime} = \Sigma_u$. Many matrices Z might solve the requirement $ZZ' = \Sigma_u$, so additional information from economic theory is needed to impose meaningful zero restrictions on Z and to identify the structural VAR. (Kilian, 2011; Robstad, 2014; Binning, 2013/14; Rubio-Ramirez, Waggoner, & Zha, 2010; among others)

Identification of the structural VAR

For exact identification of the matrix Z and hence of the structural VAR a certain number of zero restrictions needs to be imposed on the matrix; this so called order condition was first introduced by Rothenberg (1971). It implies that the matrix Z with $m \times m$ elements (m equals the number of endogenous variables in the VAR) contains $m * (m-1)/2$ zero restrictions. In our case this means we need a total number of $6 * \frac{6-1}{2} = 15$ zero restrictions to ensure exact

[4] By construction $u_t = B_0\varepsilon_t$ and $E(\varepsilon_t\varepsilon_t') = I_m$; this means the variance of u_t is given by:

$$E(u_t u_t') = B_0^{-1}\,\mathrm{E}(\varepsilon_t\varepsilon_t')B_0^{-1\prime}$$
$$\Sigma_u = B_0^{-1}\,\mathrm{I}_m B_0^{-1\prime}$$
$$\Sigma_u = B_0^{-1}B_0^{-1\prime}$$

identification of the model. However, as Rubio-Ramirez, Waggoner, & Zha (2010) prove in their paper, the order condition is only a necessary condition but not sufficient for exact identification. For exact identification of the structural VAR the zero restrictions imposed on the matrix Z must rather fulfill the necessary and sufficient rank condition, whereby the imposed zero restrictions have to follow a certain pattern, equation by equation. Rubio-Ramirez, Waggoner, & Zha (2010) sort the structural shocks in the matrix Z according to the number of imposed zero restrictions q in descending order (i.e. a structural shock with 3 zero restrictions comes before a shock with only 1 restriction). They show then that Z is only exactly identified if the number of restrictions q imposed on each shock h respectively equals: $q_h = m - k$ for $1 \leq k \leq m$; where k is the position of the respective shock in Z after sorting all shocks in descending order (Rubio-Ramirez, Waggoner, & Zha, 2010).

The most common approach of identification is to use the lower triangular Cholesky decomposition [5] of the variance-covariance matrix Σ_u with $ZZ' = \Sigma_u$ (Robstad, 2014; Björnland & Jacobsen, 2010), as was introduced by Sims (1980). The system can be solved recursively: in the first equation only lagged values are considered, in the second equation, alongside all lagged values, the contemporaneous values of the variables which were solved at first, are used, and the process is reiterated until the whole equation system is solved (Stock & Watson, March 2001). Both the order condition and the rank condition for exact identification as explained above are fulfilled. A major challenge of this approach is in how to order the variables and thus how to identify the system; the order of the variables determines the possible contemporaneous effects and hence influences the results obtained.

In the literature a monetary policy shock is commonly identified by either sorting monetary policy last or by sorting house prices last implying that the respective last shock responds to all variables contemporaneously, but does not have any contemporaneous effect on the other model variables (compare Björnland & Jacobsen, 2010). However, both interest rates and house prices might react contemporaneously (i.e. within the quarter) to news. Economic theory usually assumes quick reactions of asset prices to monetary policy shocks (Iacoviello, 2005). Further, it also seems likely that policy makers use all current information when designing monetary policy (Björnland & Jacobsen, 2010). We will follow Robstad (2014) and sort interest rates

[5] The Cholesky decomposition produces the lower triangular L of a positive-definite matrix A and its transpose L' so that $A = L * L'$ (Geijn, 2011). It is the matrix analogue of computing the square root of a scalar (Kilian, 2011).

last, implying that the domestic interest rate responds to all variables contemporaneously, but that monetary policy does not have any contemporaneous effects on the other model variables. However, as a robustness check we will also identify the system by sorting house prices last. The other variables are sorted according to the traditional closed economy VAR literature (see e.g. Sims C., 1980; Christiano, Eichenbaum, & Evans, 2005; Christiano, Eichenbaum, & Evans, 1999), and small economy assumptions, thereby following Björnland & Jacobsen (2010) and Robstad (2014). The foreign interest rate is assumed to be only contemporaneously affected by foreign monetary policy and is thus sorted on top. GDP and inflation as macro variables react according to the standard restriction in the closed economy only with a lag to monetary policy, but influence monetary policy contemporaneously; they are sorted as second and third variable. House prices are assumed not to be influenced contemporaneously by an exchange rate shock and hence are sorted before the exchange rate. The matrix in Figure 1 *a)* summarizes the ordering of variables and the implied contemporaneous zero restrictions for the Cholesky identification.

The second approach of identification we employ follows Björnland & Jacobsen (2010), and removes the restrictions limiting the contemporaneous interaction between asset prices and interest rates by introducing restrictions on the long-run effects of monetary policy instead, as was introduced by Blanchard & Quah (1989). Specifically, it is assumed that monetary policy cannot have any long-run effects on the real effective exchange rate or on the real GDP. This is in line with standard neutrality assumptions for large classes of models in the monetary policy literature (see Blanchard & Quah, 1989; Clarida & Gali, 1994; Bekaert & Hodrick, 2012) and will not have big effects for our model specifically as we only include one lag. Imposing the restrictions in this way means that monetary policy and the real exchange rate can react contemporaneously to shocks in all other variables, and house prices can react simultaneously to monetary policy shocks. The model is again exactly identified according to the sufficient rank condition and the implied order condition (in total there are still 15 zero restrictions, which can be sorted as required). The restrictions are shown in Figure 1, b).

The last identification scheme employed to identify the structural VAR consists of sign restrictions, a method introduced by Uhlig (2005). Here only the direction of a structural shock on the other model variables can be restricted to be positive or negative. This identification approach does not allow for exact identification of the structural VAR model. Theoretical restrictions can be imposed while the main questions of interest are left open, hence no a priori theorization of what a reasonable result might be, is needed (Vargas-Silva, 2008). The matrix

Z is drawn randomly from the posterior of the reduced form VAR and only the results of draws fulfilling the imposed sign restrictions are stored. In accordance with Uhlig (2005) we repeat the estimation of the structural VAR until we have 200 draws fulfilling the imposed sign restrictions. Since we are mainly interested in the effect of monetary policy, we primarily impose sign restrictions on a monetary policy shock. Following Robstad (2014), Fry & Pagan (2011), Vargas-Silva (2008), and conventional economic theory (Uhlig, 2005) we define the impact of contractionary monetary policy to be not positive on real GDP and inflation; domestic interest rates on the other hand are assumed not to decrease and also the exchange rate is presumed not to fall. For all other shocks it is only assumed that their impact on themselves leads to an appreciation of their respective variable. In Figure 1, *c)* the imposed restrictions on the monetary policy shock can be seen.

Figure 1, Imposed restrictions for identification, three identification approaches

a) Cholesky decomposition

Contemporaneous restrictions

	ε^{i^*}	ε^{g}	ε^{π}	ε^{hp}	ε^{e}	ε^{i}
i^*	x	0	0	0	0	0
g	x	x	0	0	0	0
π	x	x	x	0	0	0
hp	x	x	x	x	0	0
e	x	x	x	x	x	0
i	x	x	x	x	x	x

b) Combined zero restrictions

Short-run restrictions

	ε^{i^*}	ε^{g}	ε^{π}	ε^{hp}	ε^{e}	ε^{i}
i^*	x	0	0	0	0	0
g	x	x	0	0	0	0
π	x	x	x	0	0	0
hp	x	x	x	x	0	x
e	x	x	x	x	x	x
i	x	x	x	x	x	x

Long-run restrictions

	ε^{i^*}	ε^{g}	ε^{π}	ε^{hp}	ε^{e}	ε^{i}
i^*	x	x	x	x	x	x
g	x	x	x	x	x	0
π	x	x	x	x	x	x
hp	x	x	x	x	x	x
e	x	x	x	x	x	0
i	x	x	x	x	x	x

c) Sign restrictions

Restrictions on the monetary policy shock

	ε^{i}
i^*	x
g	-
π	-
hp	x
e	+
i	+

"x" marks that no restrictions are imposed, "0" marks zero restrictions, and "+" / "-" mark positive / negative sign restrictions.

Impulse response functions

Following the literature, we analyze the system dynamics in all three identified structural VAR models using impulse response functions. This means that each structural one-unit shock is applied once to the system at time zero (while all other structural shocks are zero) and the impacts on itself and all other model variables are estimated in the following periods.

From (2.1) we can write the structural VAR as:

$$y_t = B_0^{-1}D_0 + B_0^{-1}B_1 y_{t-1} + \cdots + B_0^{-1}B_l y_{t-l} + B_0^{-1}\varepsilon_t$$

Now it is assumed that the steady system is shocked by a one unit structural shock of size one whereby all other structural shocks are zero, implying, for example for $h = 1$, that $\varepsilon_0^1 = [1 \quad 0 \quad 0 \quad 0 \quad 0 \quad 0]'$. The response of the system to a fundamental shock then is:

For $s = 0$, $$y_0 = B_0^{-1}\varepsilon_0^h \qquad (3.1a)$$

For every $s > 0$, $$y_s = B_0^{-1}B_s y_{s-1} + \cdots + B_0^{-1}B_l y_{s-l} \qquad (3.1b)$$

We generate impulse responses for all $h = 6$ structural shocks and the impulse responses are traced for $0 \leq s \leq 12$ quarters (Iacoviello, 2009; Luetkepohl, 2011; Vigfusson, 1999; Kilian, 2011). Since the structural VAR model identified with sign restrictions is not exactly identified, the impulse responses vary for every random draw of B_0^{-1}; only IRFs in line with the imposed sign restrictions are stored until in total 200 valid draws are generated (Uhlig, 2005; Binning, 2013/14; Rubio-Ramirez, Waggoner & Zha, 2010). Following Björnland & Jacobsen (2010), impulse responses are always presented with the 84th and 16th percentile probability bands. For the structural VAR models identified by the Cholesky decomposition and combined short- and long-run zero restrictions, the MATLAB MFE-Toolbox by Sheppard (2013) employs bootstrapping with draws from the posterior of the reduced form VAR including 1,000 replications to compute the probability bands. The same procedure is for example applied by Goodhart & Hofmann (2008). For the structural VAR identified by sign restrictions, the probability bands are calculated directly from the 200 valid draws kept (Uhlig, 2005; Vargas-Silva, 2008).

4.3 Methodology for Sweden on a regional level

On a regional level we analyze the impact of monetary policy on house prices for 21 Swedish regions to study differences and similarities within Sweden. Estimations are based on a reduced form VAR of the following form:

$$y_{j,t} = C_{j,0} + A_{j,1}y_{j,t-1} + \cdots + A_{j,l}y_{j,t-l} + u_{j,t} \qquad (1.2)$$

where j denotes the respective region in Sweden, t refers to the time period, $y_{j,t}$ is the vector of endogenous variables, $C_{j,0}$ is a constant, $A_{j,l}$ are the coefficient matrices on the lags l and $u_{j,t}$

is a regional error term at time t. The variance-covariance matrix of the residuals is given by $E(u_{j,t}u'_{j,t}) = \Sigma_{j,u}$. Apart from two variables, the variables included are the same as the ones on a national level; the 3-month nominal domestic interest rate (i_t), a 3-month trade weighted nominal foreign interest rate (i_t^*), the real effective exchange rate (e_t), and inflation (π_t). Only real house prices and real GDP consist now of regional data, $hp^*_{j,t}$ and $g^*_{j,t}$. The variables are in log differences (with inflation being measured annually) and only the interest rates are in levels. All variables that are non-stationary according to an augmented Dickey Fuller test at 5% are differenced to stationarity. Again a constant is employed in all models.

To find the optimal number of lags to be included into the reduced form VARs the AIC, BIC and HQIC information criteria are computed. Estimating the reduced form VARs with the different proposed lag lengths per region shows that the model fit in terms of serially uncorrelated residuals according to a Ljung-Box test seems best when the number of lags according to the HQIC information criterion is used; thus the HQIC lag length is used in all regional VAR models, which can be seen in Appendix 3.

Since the statistical link between monetary policy and house prices was already established earlier, structural VARs are directly identified for all regions based on the respective reduced form VAR. Only the combined short- and long-run zero restrictions, as employed by Björnland & Jacobsen (2010) and explained in detail above, are used. This choice is based on the fact that the other two methods previously discussed, Cholesky decomposition and sign restrictions, showed similar results as the combined short- and long-run zero restrictions on a national level (see *6.1 Impulse responses to a monetary policy shock, total sample period, National level*). Additionally, only the combined zero restrictions and not the Cholesky decomposition allow for the likely contemporaneous interaction between house prices and monetary policy, and as Fry & Pagan (2011) stress, the information content of sign restrictions is weak, thereby making results less reliable. Thus it seems appropriate to only use one identification approach on a regional level, thereby following Robstad (2014), who focusses as well on the results from the combined zero restrictions. The identified structural VAR models on a regional level then have the following form:

$$B_{j,0}y_{j,t} = D_{j,0} + B_{j,1}y_{j,t-1} + \cdots + B_{j,l}y_{j,t-l} + \varepsilon_{j,t} \qquad (2.2)$$

where j denotes the respective region in Sweden, $B_{j,0}$ is the matrix of contemporaneous restrictions, $B_{j,0}^{-1}B_{j,l} = A_{j,l}$ and $B_{j,0}^{-1}C_{j,0} = D_{j,0}$. $\varepsilon_{j,t}$ is a vector of structural shocks with $u_{j,t} = B_{j,0}^{-1}\varepsilon_{j,t}$. The variance of the structural shocks $\varepsilon_{j,t}$ is normalized to unity with $E(\varepsilon_{j,t}\varepsilon'_{j,t}) = I_m$

so that a unit innovation in the structural shocks is an innovation of size one standard deviation, implying that $B_{j,0}^{-1} B_{j,0}^{-1\prime} = \Sigma_{j,u}$. (Kilian, 2011; Robstad, 2014; Binning, 2013/14; Rubio-Ramirez, Waggoner, & Zha, 2010; among others)

Similar to the national level we compute impulse responses for all structural shocks h also for the regional structural VAR models. Comparable to equations (3.1a) and (3.1b) we get:

For s = 0, $$y_{j,0} = B_{j,0}^{-1} \varepsilon_{j,0}^{h} \tag{3.2a}$$

For every s > 0, $$y_{j,s} = B_{j,0}^{-1} B_{j,s} y_{j,s-1} + \cdots + B_{j,0}^{-1} B_{j,l} y_{\ j,s-l} \tag{3.2b}$$

Impulse responses are again traced for $0 \leq s \leq 12$ quarters.

4.4 Boom estimation at national level

In order to examine if the relation between monetary policy and house prices differs in times of asset price booms versus times not characterized by a boom, the structural VAR model is extended with a dummy variable for identified boom periods and for periods without a boom. Thereby we primarily follow Goodhart & Hofmann (2008). Booms are defined as a positive deviation from the smooth trend in real house prices following Borio & Lowe (2004), Adalid & Detken (2007), and Goodhart & Hofmann (2008).

Following Adalid & Detken (2007) we define a boom in real house prices as a 10% positive deviation from the smooth, national trend. In order for a period to qualify as boom, at least ten consecutive periods need to show the positive deviation from the trend. This period of ten consecutive quarters that we employ is longer than the four consecutive quarters Adalid & Detken (2007) use to identify a boom; however, the period is still shorter than the period employed by Goodhart & Hofmann (2008), who use a 5% positive deviation lasting for at least 12 quarters for boom identification. Real house prices in Sweden have been increasing strongly over the last 20 years (1996 Q1-2014 Q2, the growth was ~ 160%, see also Appendix 1), thus this rather long period combined with a high deviation seems appropriate for a realistic boom estimation and allows us to capture both boom and non-boom periods in the national sample. However, to ensure robustness of the results we also estimate booms employing the thresholds given by Adalid & Detken (2007), and Goodhart & Hofmann (2008).

The trend in real house prices is computed using a one-sided Hodrick-Prescott (HP) filter with a high smoothing parameter of 100,000 (Goodhart & Hofmann, 2008; Adalid & Detken, 2007). This means that the trend only adjusts slowly to new information, thus capturing low frequency, cumulative deviations, and implicitly emphasizing mean reversion tendencies. The trend is estimated recursively, using information available only up to the quarter when the boom estimation is made respectively. To come up with meaningful trends already in the first estimation period 1996 Q1, we use house price data starting from 1986 Q1 (Borio & Lowe, 2004). Quarterly real house prices, indexed at 1996 Q1 = 100 are used for the analysis.

The identified boom and non-boom periods are applied on the above described reduced form VAR model (1.1) as dummy variables; the dummy-extended reduced form VAR model has the following form:

$$y_t = C_0 + A_{B,1}y_{t-1} * D_1^B + A_{NB,1}y_{t-1} * D_1^{NB} + \cdots + A_{B,l}y_{t-l} * D_l^B + A_{NB,l}y_{t-l} * D_l^{NB} + u_t \qquad (4.1)$$

where D_t^B and D_t^{NB} are dummy variables; D_t^B is set to one during estimated house price boom periods t and set to zero in non-boom periods. D_t^{NB} on the other hand is set to one in estimated non-boom periods t and set to zero during boom periods. A structural VAR model is then identified based on the dummy-extended reduced form VAR as before using three identification methods and impulse responses for the boom and non-boom sample are computed. (Goodhart & Hofmann, 2008)

To test whether the impulse responses in times of booms and non-booms are statistically different we use, following Adalid & Detken (2007), the Wilcoxon-Mann-Whitney test. This test is a non-parametric rank-sum test for differences in population. The null hypothesis is that there is no difference between the two populations when the sum of the ranks of the two samples is relatively different (Adalid & Detken, 2007).

4.5 Boom estimation at regional level

To also see if the impact of monetary policy on house prices differs in boom periods versus non-boom periods on a regional level, we identify these different periods per region. A regional boom period is thereby defined as a 10% positive deviation from the smooth regional trend that lasts for at least 10 quarters. Technically the regional trends are computed in the same way as before in the national boom estimation apart from that now the underlying smooth trend is based on the regional house price development. The regional smooth trend rather than the national trend is used since house prices are usually set in a local context (see e.g. Van Soest & Niu, 2014; Case & Shiller, 2003). Furthermore, supply and demand for housing depend mainly on idiosyncratic and regional factors (see e.g. Vargas-Silva, 2008; Fratantoni & Schuh, 2003). A

HP-filter with a smoothing parameter of 100,000 is again used and the trend is estimated recursively (see *4.4 Boom estimation at national level*).

The dummy-extended reduced form VAR, which is again the basis for the structural VAR, has the following form:

$$y_{j,t} = C_{j,0} + A_{j,B,1}y_{j,t-1} * D_{j,1}^{B} + A_{j,NB,1}y_{j,t-1} * D_{j,1}^{NB} + \cdots + A_{j,B,l}y_{j,t-l} * D_{j,l}^{B} + A_{j,NB,l}y_{j,t-l} * D_{j,l}^{NB} + u_{j,t} \qquad (4.2)$$

where $D_{j,t}^{B}$ and $D_{j,t}^{NB}$ are dummy variables per region j; $D_{j,t}^{B}$ is set to one during estimated house price boom periods and set to zero in non-boom periods. $D_{j,t}^{NB}$ is set to one in estimated non-boom periods t, and set to zero during boom periods. Structural VAR models are identified as before through combined short- and long-run zero restrictions based on the dummy-extended reduced form VAR and impulse responses are calculated for boom and for non-boom samples, comparable to the methodology employed on a national level. Similar to the national level a Mann-Whitney-Wilcoxon test is also applied to the regional results to determine whether the difference in impulse responses during boom and non-boom periods is significant in each region respectively (see *4.4 Boom estimation at national level*).

5. Data

In the structural VAR analyses on a national and on a regional level quarterly data from 1996 Q1 until 2014 Q2 is used. This sample period is chosen since the Riksbank's objective of price stability became effective in 1995 (Sveriges Riksbank, 2011; Ingves, 2007) and it seems likely that after one year of adjusting policies to the new objective, i.e. from 1996 on, Sweden can be characterized by having a stable monetary policy regime. This is an important feature as a stable monetary policy regime is essential for meaningful analyses of the impact of monetary policy shocks (Robstad, 2014; Björnland & Jacobsen, 2010). Most of the series are accessed from Datastream, only regional house price data and regional gross domestic product (RGDP) data is directly retrieved from Statiska centralbyrån (SCB). A summary of the data and variables used can be found in Appendix 4.

The domestic interest rate corresponds to the nominal 3-month STIBOR rate (Stockholm Interbank Offered Rate), released by Sveriges Riksbank. The real effective exchange rate (REER) is the trade weighted average of bilateral SEK exchange rates adjusted by relative consumer prices computed by the Bank for International Settlement (BIS); it is based on the narrow index comprising 27 economies. An increase in the REER indicates a real appreciation of the SEK and a decrease shows a real depreciation (Bank For International Settlements, 2015). The foreign interest rate is calculated as the trade weighted average of the nominal 3-month interbank rates of Sweden's trading partners; thereby the same BIS trade weights as for the REER are applied. House prices both on a national and on a regional level refer to owner-occupied one-and two-dwelling buildings for permanent living and are based on the real estate price index, which is compiled by the SCB. The indices are adjusted to real values by the Swedish CPI. The CPI includes all items and is from the International Monetary Fund (IMF). When estimating boom periods in the Swedish housing market, real house prices are indexed (1996 Q1 = 100) and data from 1986 Q1 until 2014 Q2 is used. In Appendix 1 the regional real house price indices from 1996 Q1 onwards are displayed together with the national real house price index. Output is measured as real, seasonally adjusted GDP in SEK and is provided by the ECB. On a regional level regional GDP as provided by the SCB is used, adjusted by the CPI to account for inflation. Since this data is only available at an annual frequency the series are, following Goodhart & Hofmann (2008), interpolated to quarterly data using the Chow-Lin interpolation method. The Chow & Lin (1971) interpolation method allows for a linear interpolation of time series by a related series; here the Swedish national GDP is used as related series assuming that regional GDP follows the same underlying trends as the national GDP. The interpolation was executed employing the MATLAB Disaggregation Library by Quilis (2013).

6. Results

The results section is divided into two parts. The first part presents and discusses the results obtained for Sweden on an aggregated, national level and on a disaggregated, regional level for the total sample period. The second part focuses on the differences during boom versus non-boom periods, both on an aggregated and a disaggregated level.

6.1 Impulse responses to a monetary policy shock, total sample period

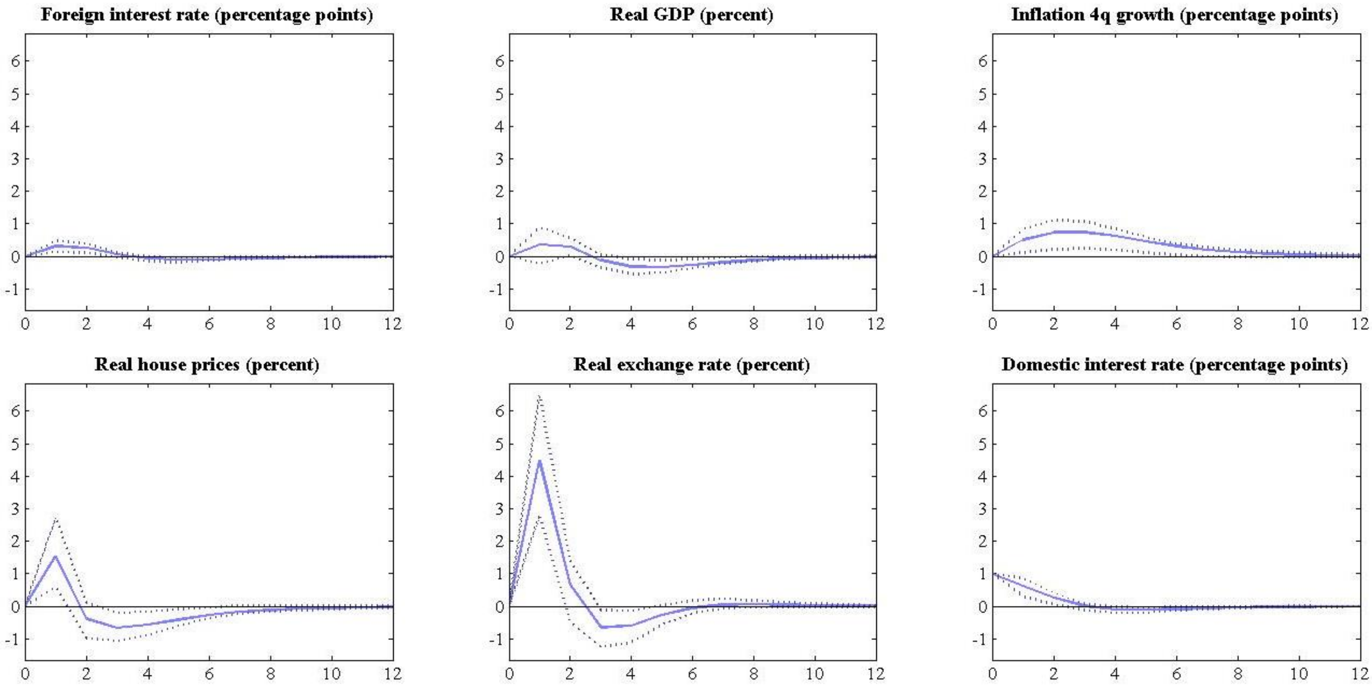
This section will present the effects of a monetary policy shock for the total sample first on a national level and thereafter on a regional level. The analyses will evaluate whether Swedish house prices react to a monetary policy shock in a similar way as has been found in previous literature, as well as whether a monetary policy shock has varying effects across regions in Sweden, similar to what has been found in the US.

National level

As previously described, for Sweden on a national level monetary policy shocks are identified using three different identification approaches; the Cholesky decomposition, combined short- and long-run zero restrictions, and sign restrictions. The respective impulse responses per identification method for the foreign interest rate, real GDP, inflation, real house prices, the real exchange rate and the domestic interest rate to a contractionary monetary policy shock are reported in Figure 2, Figure 3, and Figure 4. All impulse responses are normalized to a 1% monetary policy shock and are recorded for 12 quarters.

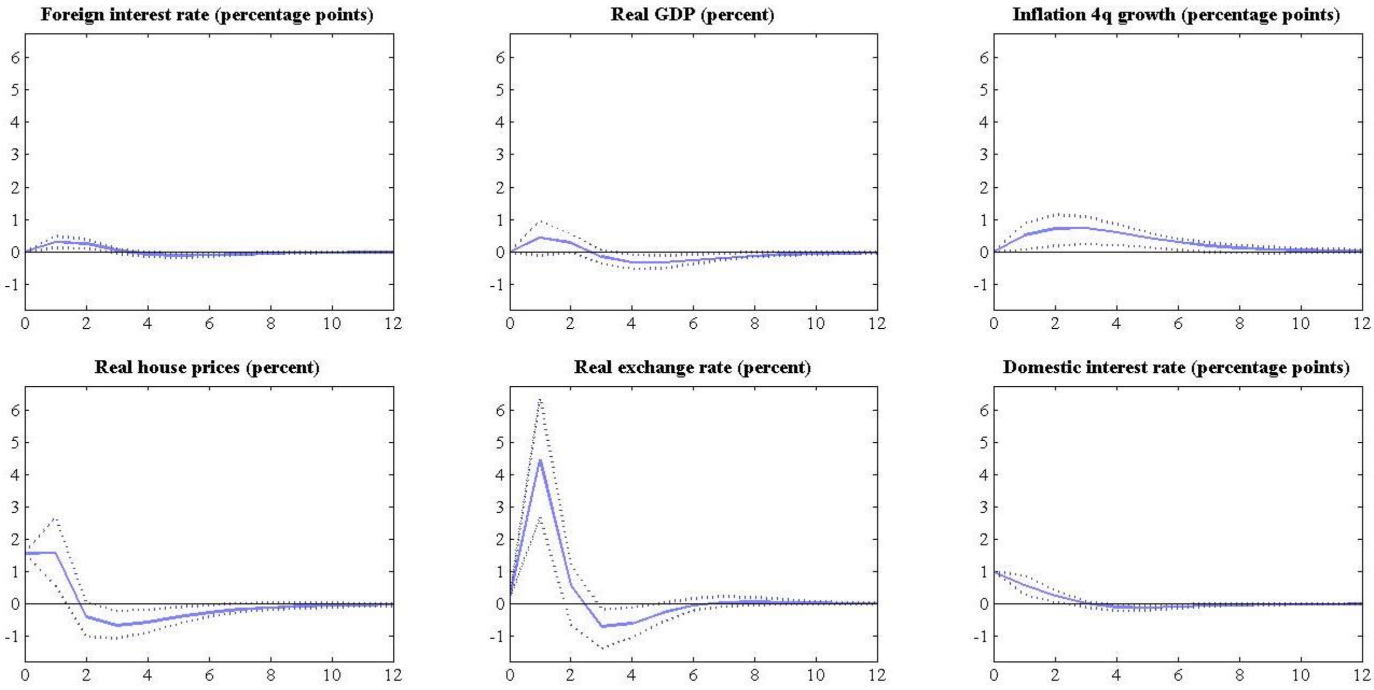
The comparison of the overall patterns in the impulse responses to a monetary policy shock for the different identification methods shows similar results for all three identification approaches, indicating overall robustness of the estimated effects. This is further confirmed by the similar impulse responses for an identification through the Cholesky decomposition employing a different ordering of the variables, see Appendix 5. For the Cholesky decomposition all contemporaneous effects of monetary policy are by construction restricted to be zero, whereas for the combined short- and long-run zero restrictions some contemporaneous effects are allowed. The identifications through the Cholesky decomposition and through combined zero restrictions show not only very similar patterns, but are also alike in the magnitude and persistency of the impacts. This holds as well for the persistency of a monetary policy shock on real GDP and on the real exchange rate even though in the combined short- and long-run zero restrictions the long-run effect of monetary policy is restricted to zero. However, this might be caused by the fact that only one lag is included into the VAR model and that hence the effect of this long-run restriction is very small. The impulse responses for the identification through sign restrictions are less pronounced than for the other two identification methods, particularly of house prices and of the real exchange rate. This might be caused by the weaker information content of sign restrictions leading to less clear and reliable results (Fry & Pagan, 2011).

Figure 2, Monetary Policy Shock - Cholesky decomposition, monetary policy ordered last



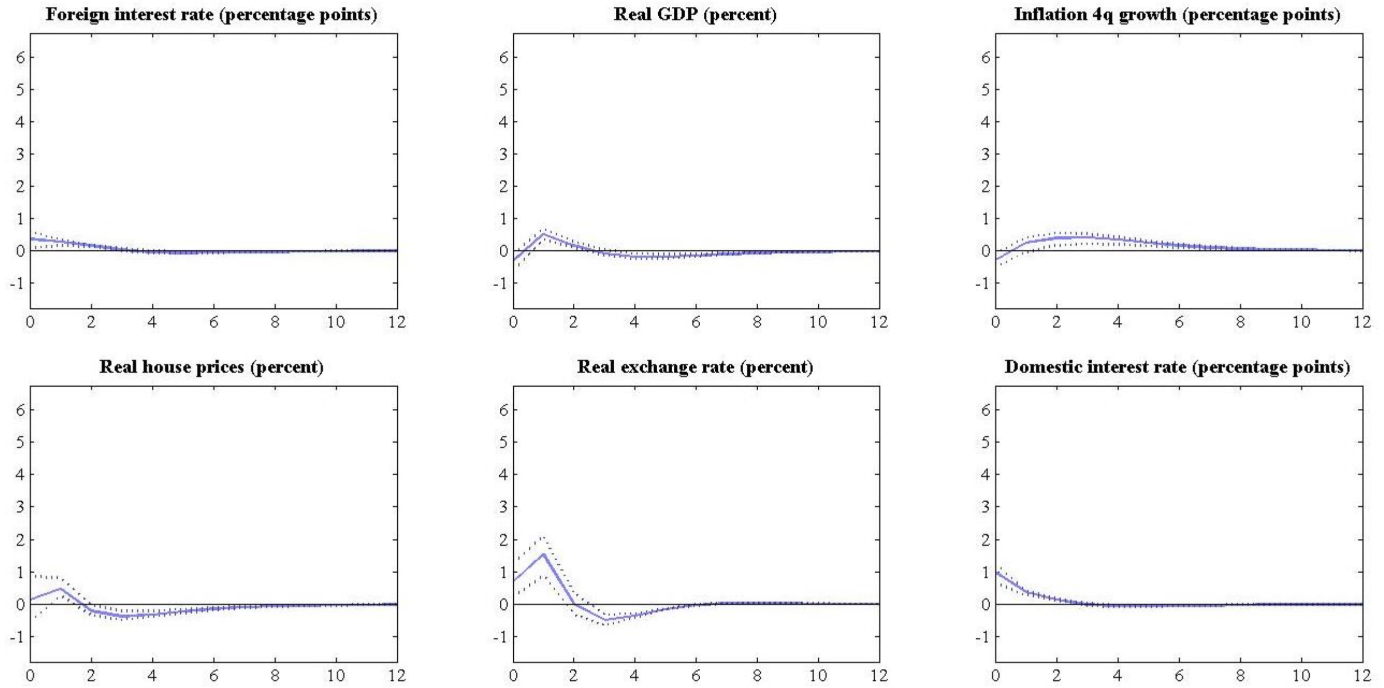
The impulse responses (y-axis), recorded for 12 quarters (along the x-axis), are normalized to a 1% monetary policy shock. Identification is achieved through a Cholesky decomposition with monetary policy ordered last. Solid lines are the impulse response estimates whereas the dotted lines correspond to the 16th and 84th percentile probability bands.

Figure 3, Monetary Policy Shock - combined short- and long-run zero restrictions



The impulse responses (y-axis), recorded for 12 quarters (along the x-axis), are normalized to a 1% monetary policy shock. Identification is achieved through combined short- and long-run zero restrictions. Solid lines are the impulse response estimates whereas the dotted lines correspond to the 16th and 84th percentile probability bands.

Figure 4, Monetary policy shock - sign restrictions



The impulse responses (y-axis), recorded for 12 quarters (along the x-axis), are normalized to a 1% monetary policy shock. Identification is achieved through sign restrictions. Solid lines are the median estimates based on 200 valid draws whereas the dotted lines correspond to the 16th and 84th percentile probability bands.

When we examine the impulse responses of house prices to a monetary policy shock in detail we see that for all three identification methods house prices exhibit a significant positive impulse response in the first period. For both the Cholesky decomposition and the combined zero restrictions as identification approaches the impulse responses show a positive spike of about 1.6% in period one; for sign restrictions the initial positive spike is of about 0.5%. However, the initial, contemporaneous responses vary; for identfication methods not restricting the simultaneous impact, i.e. the combined short- and long-run zero restrictions and sign restrictions, the contemporaneous impulse responses are positive, albeit this impact is only significant at the $84^{th}/16^{th}$ percentile probability bands when employing the combined zero restrictions. For the Cholesky decomposition the contemporaneous impulse response is by construction zero. The impulse responses for all three identification schemes become negative starting in quarter two, and bottom out at -0.5% in quarter three before reverting towards zero until quarter eight. This impulse response pattern of house prices to a monetary policy shock, exhibiting a positive impulse response in period one and only afterwards becoming negative for about eight quarters, differs from most previous research; most studies find completely negative impulse responses and also the persistency is commonly found to be longer than eight quarters (see e.g. Robstad, 2014; Jarocinski & Smets, 2008; Elbourne, 2008).

As for example outlined by Elbourne (2008) and Mishkin (2007), an immediate drop in house prices to a monetary policy shock would follow standard economic reasoning; as soon as the interest rate increases, potential house buyers have to afford a higher initial interest payment and a higher consequent interest rate burden. Hence, the amount a potential house buyer is willing and capable to pay for a house decreases. This means in broader terms that the demand for housing is negatively related to interst rates since interest payments represent such a large share of the costs of buying a house (Elbourne, 2008; Mishkin, 2007). In comparable previous research on the effects of monetary policy on house prices in Sweden, both Björnland & Jacobsen (2010) and Laseén & Strid (2013) find, in contrast to us, fully negative impulse responses lasting for considerably longer than eight quarters. Nevertheless, neither results are directly comparable to our estimations; Björnland & Jacobsen (2010) use a sample period from 1986 Q1 until 2006 Q4, which might account for the more pronounced effect of monetary policy on house prices with a trough of -4.5% in quarter eight. Laseén & Strid (2013) find, using a sample period from 1995 Q1-2013 Q2, a trough of about -1.2% in quarter five. The trough is of a comparable magnitude as the one we find, but since Laseén and Strid's (2013) model and the choice of variables, driven by their research question of the impact of monetary policy on household debt levels, differs considerably from our variables and model, also their results are not directly comparable to ours.

In a similar way to us, Musso, Neri, & Stracca (2010) and McCarthy & Peach (2002) observe in their studies that house prices during the first periods increase as response to a monetary policy shock and decrease only afterwards. In Musso, Neri, & Stracca's (2010) paper the impulse response in quarter one is positive before turning negative for their Euro sample excluding Germany, just as in our sample. McCarthy & Peach (2002) find that house prices in the US only turn negative in response to a monetary shock in the third quarter after the shock. McCarthy & Peach (2002) refer to this phenomenon as the "home prize puzzle", and argue that the effect is firstly related to households' loss aversion as found by Genesove & Mayer (2001) and to loan-to-value (LTV) constraints as discussed by Genesove & Mayer (1997), and secondly to the expectations of potential further interest rate increases in the future. Loss aversion refers to the fact that potential house sellers are reluctant to realize losses in form of lower prices. This implies that ask prices are not adjusted downwards in accordance to the lowered fundamental value when interest rates increase, meaning that the sale is delayed, or the offer is withdrawn altogether from the market (see Genesove & Mayer, 2001; McCarthy & Peach, 2002). Often house sellers might in addition not be able to afford to sell at lower prices during a downturn in the housing market due to their already high loan-to-value ratio. A high LTV ratio implies that if the house is sold at a lower price than the value presumed in the mortgage contract, the LTV ratio increases further, which might make it impossible for the seller to afford an equivalent house with the proceeds of the sale. This makes potential house sellers become less willing to sell their houses compared to what they would be without the shock being present (see Genesove & Mayer, 1997; McCarthy & Peach, 2002). Both the effect of loss aversion and the effect of the LTV constraints reduce the supply in the housing market compared to the level of supply in the absence of a shock, thereby potentially increasing the prices for housing.

The second line of reasoning for a house price increase as reaction to a monetary policy shock refers to a sustained or even growing demand for housing despite increased costs. If potential house buyers expect further increases in interest rates they might perceive it to be better for them to buy a house immediately at the only slightly increased interest rate costs, than to wait longer, since postponing the transaction might make the financing even more expensive. As a result house buyers might rush to the housing market, thereby sustaining or even increasing the demand for housing and thus potentially also increasing house prices. Particularly the interplay of the reduced supply and the sustained, or even increased demand of houses, might explain why we see an initial increase in house prices as reaction to a contractionary monetary policy shock (McCarthy & Peach, 2002).

For all three identification methods the impulse responses of real GDP to a monetary policy shock are positive in the first two quarters before becoming negative in period three. Even though the positive responses are only significant for the identification through sign restrictions, the responses do not follow the conventional view of economic theory predicting that real GDP should drop immediately on a contractionary monetary policy shock and not only in subsequent periods (Vargas-Silva, 2008; Uhlig, 2005). Exactly this notion is also imposed in form of a negative sign restriction for the contemporaneous impact of monetary policy on real GDP. Noteworthy is, however, as Uhlig (2005) shows in his paper particularly examining the effect of monetary policy on real GDP, the impact of monetary policy does not necessarily have to be negative but can also be positive, indicating that the current consensus in the literature should not be taken for granted. A similar unexpected positive effect to a shock in monetary policy can be observed for inflation. Here it is even more pronounced than for real GDP; for all three identification methods the impulse responses are significantly positive for about six to eight quarters even though again for sign restrictions the contemporaneous impact is restricted to be negative. However, previous research, starting with Sims (1992) and thereafter confirmed by many subsequent papers (see e.g. Barth & Ramey, 2001; Björnland & Jacobsen, 2010; Robstad, 2014), has found similar patterns in the impulse responses of inflation, and commonly refer to this as the “price puzzle”. The puzzle is related to the cost channel of monetary policy and can be explained by a short-run cost-push, where firms pass on the increased borrowing costs (resulting from the increased domestic interest rate) to the prices of goods; inflation thus increases in the short run (Barth & Ramey, 2001; Ravenna & Walsh, 2006; Chowdbury;Hoffmann;& Schabert, 2006). Additionally, the price puzzle might also explain the positive real GDP in the first periods, since firms, by passing on increased borrowing costs to prices, can avoid to immediately reduce their production, thereby delaying a drop in real GDP to later periods.

For the foreign interest rate the almost zero impact of a monetary policy shock is in line with the assumption of Sweden as a small open economy (Björnland & Jacobsen, 2010). Sweden’s monetary policy has hence none or only a negligible influence on the interest rate of its main trading partners. However, whereas the contemporaneous impulse response for the Cholesky decomposition and the combined short- and long-run zero restrictions is zero by construction, the identification through sign restrictions leads to a slightly positive and significant simultaneous impact of about 0.4%. This might reflect Sweden’s close links with its main trading partners, especially with the Eurozone, in that an increase in the STIBOR coincides with an increase in a foreign interest rate.

The impulse response of monetary policy on the real effective exchange rate shows a significant and highly positive spike in quarter one of about 4.5% for the Cholesky decomposition and for the combined zero restrictions and of about 1.5% for sign restrictions, indicating a real appreciation of the Swedish Krona. Contractionary monetary policy leads to an increase in the domestic interest rate while the foreign interest rate is kept constant; this leads to an increased demand for the Swedish Krona in the short run (since it becomes relatively more attractive to invest in Sweden), thereby driving up the exchange rate (Bekaert & Hodrick, 2012). The lower impulse response for the identification through sign restrictions might be explained by the contemporaneous increase in the foreign interest rate as response to a contractionary monetary policy shock in our model. This would dilute the effect of increased currency demand and hence lead to a smaller effect in the real exchange rate. The to 1% normalized impulse response of the domestic interest rate to a monetary policy shock exhibits a persistency of about four quarters, which is in line with previous research (Björnland & Jacobsen, 2010).

In conclusion we find that real house prices in Sweden on a national level drop after an initial increase approximately -0.5% as a response to a positive interest rate shock of 1%. Our model estimates reasonable effects of a contractionary monetary policy shock also on the other five variables in the model. The results seem to be robust since all three employed identification methods yield similar impulse responses. However, for example Goodman (1998) argues that using national aggregates when looking at the effects of monetary policy on the housing market might lead to biased estimates. This is caused by a differing marginal reaction of local housing markets to shocks (Goodman, 1998). Furthermore, previous literature has found substantially varying effects of monetary policy across regions, thereby supporting the claim of potentially biased estimates for geographically aggregated data. Especially the housing market might be affected, since supply and demand depend heavily on idiosyncratic and regional factors (see e.g. Vargas-Silva, 2008; Fratantoni & Schuh, 2003). To ensure an appropriate estimate of the actual effects of monetary policy on the housing market in Sweden, in the next section hence impulse responses on a regional, disaggregated level are analyzed.

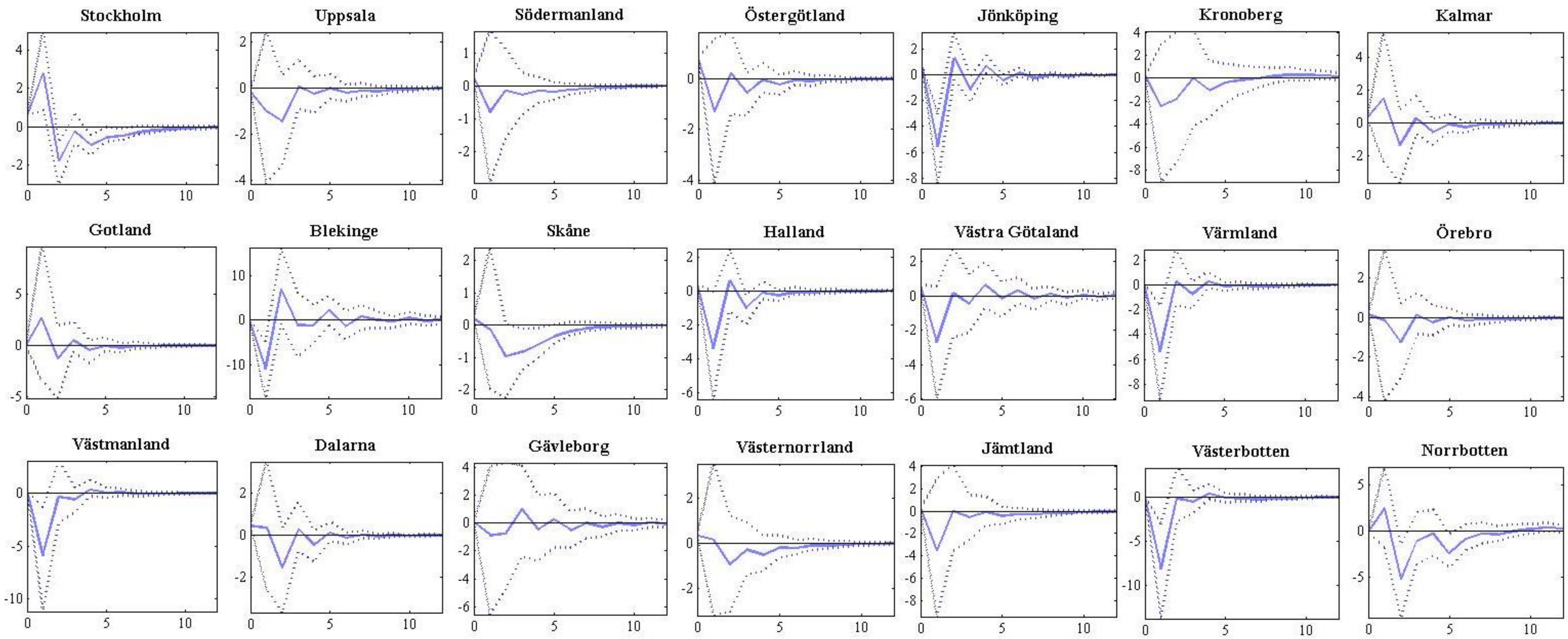
Regional level

On a disaggregated level monetary policy shocks are identified only using combined short- and long-run zero restrictions. Regional real GDP and regional real house prices are used as substitute for the national series. The impulse responses of real house prices to a contractionary monetary policy shock, estimated for the 21 Swedish regions in our sample, are displayed in Figure 5.

Comparing the overall patterns of regional impulse responses reveals a great variation in the effect of monetary policy across regions both in the initial response, as well as in the timing and magnitude of a negative trough, and in the overall persistency of the monetary policy shock. Notable is that the magnitude of the maximum negative impact seems for many regions to be considerably larger than for Sweden on a national level, indicating that the impact of a monetary policy shock might be diluted when analyzed on an aggregated, national level. Of the seven regions having significant impulse responses, Stockholm is the only region exhibiting similar, but more pronounced, impulse responses as Sweden on a national level; in quarter one impulse responses are significantly positive before they turn negative in the following period. As one of Sweden's core regions, it is reasonable that Stockholm has a considerable influence on the aggregated national results and drives the observed effects. Real house prices in all other regions with significant impulse responses (Jönköping, Blekinge, Halland, Värmland, Västmanland, and Västerbotten) show a negative response to a monetary policy shock already in the first quarter after the shock. When taking all regions into account, and thus also including the non-significant results at the 84th/16th percentile error bands, the house price responses for nine out of 21 regions experience a trough in quarter two whereas for the rest of the regions the maximum downward adjustment occurs already in quarter one. Only six of the regions show a positive impulse response in quarter one before turning negative. The magnitude of the maximum drop varies between about 1% (for example Uppsala, Skåne, and Örebro) to over 5% (Jönköping, Värmland, Västmanland, and Västerbotten). This implies that for most regions in Sweden monetary policy has a merely negative and quite strong impact on house prices in line with the theory of decreasing demand for housing as soon as financing costs increase. Only in six regions, notably including Stockholm, effects of loss aversion, LTV constraints or expectations of further interest rate increases seem to affect house prices initially.

In conclusion monetary policy seems to have varying effects across regions in Sweden, and hence our results are in line with the findings of for example Vargas-Silva (2008) and Fratantoni & Schuh (2003). As previous literature has mainly been conducted on the United States, where regions are arguably more independent than in Sweden because of the federal system, it can be considered a valuable insight to confirm that we see heterogeneous effects of monetary policy on house prices also in Sweden. Furthermore, the results suggest that impulse responses on a national level are mostly driven by developments in Stockholm. The aggregated estimate thus does not represent fairly actual developments across regions in Sweden. To further see if the heterogeneous effects are driven by different phases of business cycles to which the regions are exposed, we will later estimate impulse responses in house price boom versus non-boom periods.

Figure 5,Monetary policy shock on real house prices, regional level - combined short- and long-run zero restrictions



The percentage impulse responses of real house prices per region (y-axis), recorded for 12 quarters (along the x-axis), are normalized to a 1% monetary policy shock. Identification is achieved through combined short- and long-run zero restrictions. Solid lines are the impulse response estimates whereas the dotted lines correspond to the 16th and 84th percentile probability bands.

6.2 Impulse responses during boom versus non-boom periods

In the second part of the results section we will analyze whether house prices react differently to a monetary policy shock depending on the underlying growth in the housing market; this means that we differentiate between house price boom and non-boom periods. Comparable to section *6.1 Impulse responses to a monetary policy shock, total sample period*, this analysis will be done first on a national and then on a regional level.

National level

As discussed before in greater detail, previous research finds that the impact of a monetary policy shock on house prices is larger when house prices are booming (see e.g. Goodhart & Hofmann, 2008). We test this for Sweden on a national level and compute impulse responses for all three identification methods during estimated boom and non-boom periods. Based on our chosen thresholds of 10% deviation for at least ten quarters for identifying a house price boom in Sweden, we find a national boom period from 1998 Q2-2007 Q4, see Figure 6. When using the thresholds introduced by Goodhart & Hofmann (2008) of a 5% deviation for 12 quarters or Adalid & Detken (2007) of 10% deviation for four quarters, equivalent boom periods are identified from 1998 Q2-2008 Q3 and 1998 Q2-2007 Q4 respectively, see Appendix 6, Appendix 7. This is also broadly in line with what has been identified as boom periods for Sweden by other researchers. For example Agnello & Schuknecht (2009) identify a boom from 1997-2007, and Goodhart & Hofmann (2008) estimate a boom for their shorter sample between 1998 Q1 and 2006 Q4.

Figure 7 and Figure 8 show the impulse responses for boom and non-boom periods respectively using the combined short- and long-run zero restrictions for identification. To ensure robustness of the results the structural VAR for both boom and non-boom periods is also identified through the Cholesky decomposition and sign restrictions, see Appendix 8 & Appendix 9, and Appendix 10 & Appendix 11 for the results. The impulse responses look very similar for all three identification approaches, and merely the IRFs for sign restrictions are again less pronounced, similar to the results for the total sample period. In the following paragraphs we will hence only discuss in detail the impulse responses identified through combined short- and long-run zero restrictions.

Figure 6, Identified boom periods, Sweden at national level, 10% deviation for at least 10 quarters

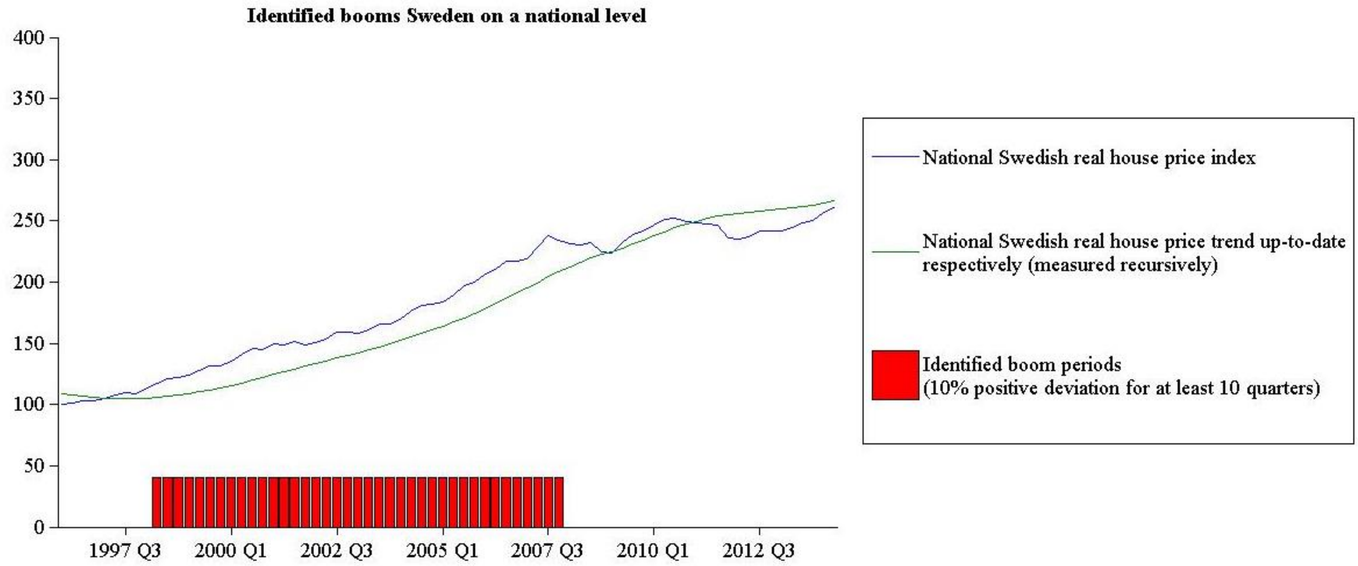
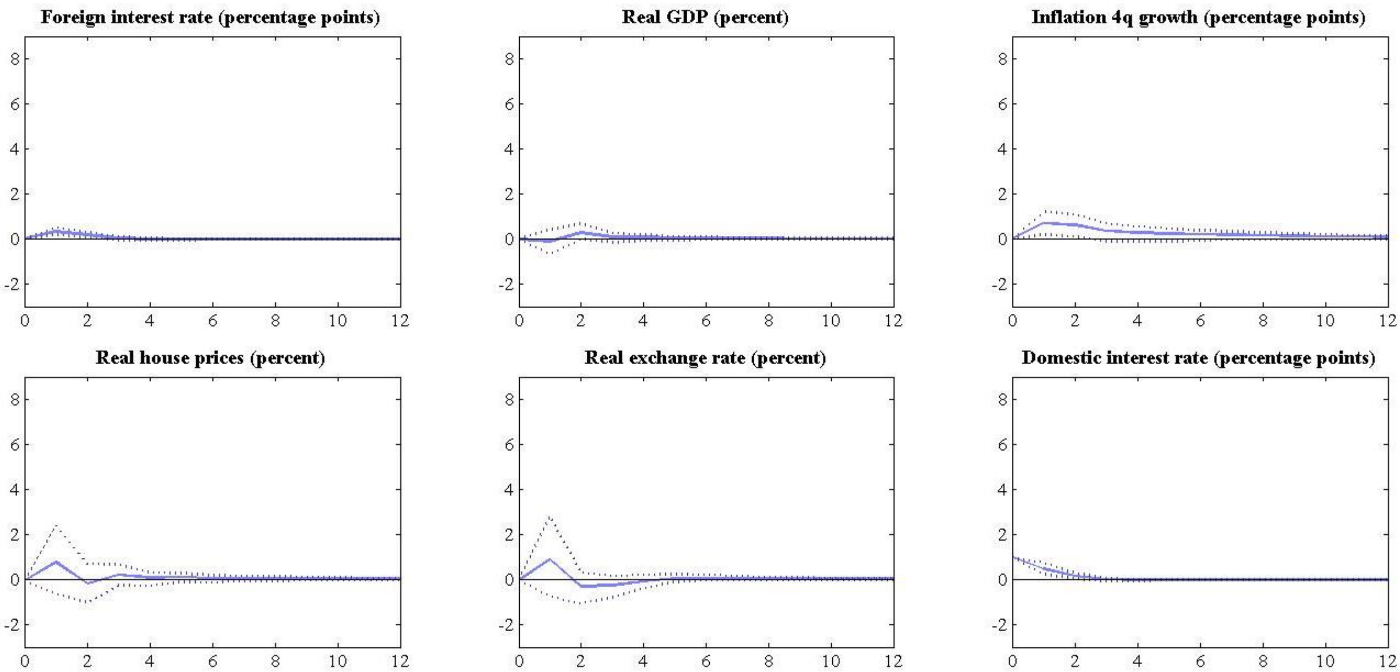
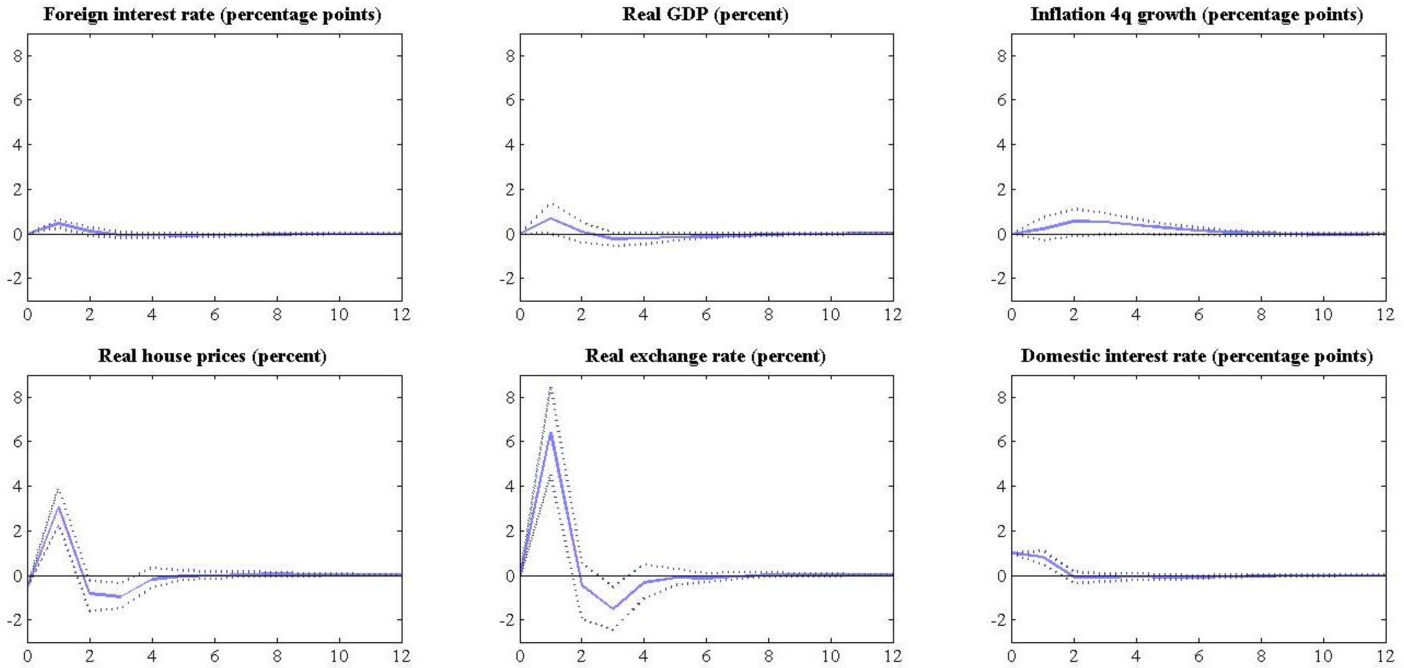


Figure 7, Monetary policy shock - boom periods (10q-10% deviation), combined zero restrictions



The impulse responses during boom periods (y-axis), recorded for 12 quarters (along the x-axis), are normalized to a 1% monetary policy shock. Identification is achieved through combined short- and long-run zero restrictions. Solid lines are the impulse response estimates whereas the dotted lines correspond to the 16th and 84th percentile probability bands.

Figure 8, Monetary policy shock - non-boom periods (10q-10% deviation), combined zero restrictions



The impulse responses during non-boom periods (y-axis), recorded for 12 quarters (along the x-axis), are normalized to a 1% monetary policy shock. Identification is achieved through combined short- and long-run zero restrictions. Solid lines are the impulse response estimates whereas the dotted lines correspond to the 16th and 84th percentile probability bands.

A comparison of the impulse responses for boom versus non-boom periods indicates that the general pattern for all variables looks quite similar for boom versus non-boom periods. However, during booms the impulse responses are considerably lower and insignificant for real house prices, the real exchange rate and real GDP. During non-boom periods, on the other hand, a monetary policy shock leads to an intensified effect on house prices following the same pattern as for the whole sample period. The results do not change when applying the slightly differing boom period between 1998 Q1 and 2008 Q3, identified by the thresholds proposed by Goodhart & Hofmann (2008), see Appendix 12, Appendix 13. A Mann-Whitney-Wilcoxon (MWW) test applied on the impulse responses of real house prices confirms at a 5% significance level with a p-value of 0.018 that the boom and non-boom impulse responses do not come from the same underlying population.

Our results showing a muted impact of a monetary policy shock in boom periods seem to go against the findings of Goodhart & Hofmann (2008), who observe that the impact of monetary policy is larger in boom periods than in periods not characterized by booms. A larger decrease in house prices during boom periods would follow from the above described housing demand mechanism as for example stated by Elbourne (2008). During a house price boom, the initial interest rate payment and the total interest rate burden are based on a higher underlying value of the house than during non-boom periods. An interest rate shock in this environment might make it unattractive or impossible for a relatively higher fraction of potential buyers to afford a house, leading to a relatively stronger decrease in demand for housing. To meet the decreased demand and lower willingness to pay, sellers need to lower their ask prices more severely, and consequently house prices drop more strongly. The fact that house prices during boom periods are not significantly impacted by a monetary policy shock, on the other hand, seems hard to reconcile with rational explanations. However, the insignificant responses might be explained by an interplay of demand and supply factors. This means that the decreased demand might be counterbalanced by a similarly decreased supply caused by loss aversion and LTV constraints, as mentioned by McCarthy & Peach's (2002). During boom periods both home sellers' anticipated losses and their LTV constraints could be larger than in non-boom periods, especially if for example a potential seller has already bought the house at a very high price close to the peak. This would decrease the housing supply to a larger extent during boom periods and might mitigate a strong decrease in demand for housing; house prices would stay stable. Another explanation for our results might be related to the aggregation bias found by Goodman (1998) and described in section *6.1 Impulse responses to a monetary policy shock, total sample period, National level*. According to that the observed results might just be the consequence of using national housing price data for actually locally operating and reacting housing markets. Furthermore, the muted impact of a monetary policy shock in times of booms could also be related to an endogeneity bias as outlined by Crowe, Dell'Ariccia, Igan, & Rabanal (2013). If central banks tighten the interest rates as reaction to a house price appreciation, on average higher interest rates would coincide with faster house price growth. This would tend to reduce the size and significance of the regression coefficients, leading to an underestimation of the effectiveness of monetary policy (Crowe, Dell'Ariccia, Igan, & Rabanal, 2013).

During non-boom periods the more pronounced impact of monetary policy on house prices, where the initial spike is even larger compared to the total sample (around 3% compared to approximately 1.5%, see Figure 3), might be related to expectations of further interest rate increases (see McCarthy & Peach, 2002). Particularly in non-boom periods, i.e. in times of average house price growth, interest rate increases might be interpreted as a signal of continuing economic growth. If the Riksbank is expected to have superior information about the future economic development of the country, the first interest rate increase during average growth times might be interpreted as a signal of confidence in the economy. The signal could then be seen as a factor increasing the likelihood of future increases of the interest rate in order to keep inflation stable. This is also supported by the fact that central banks would hardly raise interest rates if a downturn was expected shortly thereafter. The concern of future interest rate increases, making financing of a house potentially more expensive, thus might stabilize or even increase the demand for housing at least for some time, thereby explaining the sharp initial spike.

To sum up, during boom periods house prices seem to respond rather unexpectedly to a monetary policy shock in not showing significantly negative impulse responses. However factors like loss aversion and LTV constraints in an interplay with decreased housing demand might explain these reactions. During non-boom periods, on the other hand, the impact of monetary policy seems to be intensified compared to the total sample IRFs, potentially caused by a signaling effect of future economic growth. Nonetheless, whether our results are really related to these factors or if rather an aggregation bias can explain the effects will be analyzed in the following section by estimating the impact of monetary policy on house prices in boom and non-boom periods on a disaggregated, regional level.

Regional level

As we have seen previously, monetary policy has heterogeneous effects across regions. This could be driven by different phases of business cycles the single regions are exposed to. To examine this, we estimate the impulse responses of house prices to a monetary policy shock during house price boom and non-boom periods per region. We thereby follow Campbell, Yang & Wang's (2010) suggested theme for further research regarding putting insights into the effects of monetary policy in different phases of business cycles on a regional level. The analysis might additionally allow making inferences regarding the results observed on a national level for boom versus non-boom periods.

The fact that house prices are usually set in a local context (see e.g. Van Soest & Niu, 2014; Case & Shiller, 2003), and that economic agents when forming expectations, are often influenced by the nearby environment (Easaw & Mossay, 2015), motivates the estimation of regional booms as a deviation from each region's smooth trend. Households might not look at the national price level when acting on the regional house market, but rather take into account only regional conditions. In other words, the reaction of house prices to a monetary policy shock is more likely to depend on the relative regional house price growth rather than the regional house price growth in relation to the national trend. Figure 9 (a), (b) depict the periods in which a region experiences a house price boom relative to its regional trend respectively.

Figure 9(a), Identified boom periods (10q-10% deviation), regional level Sweden

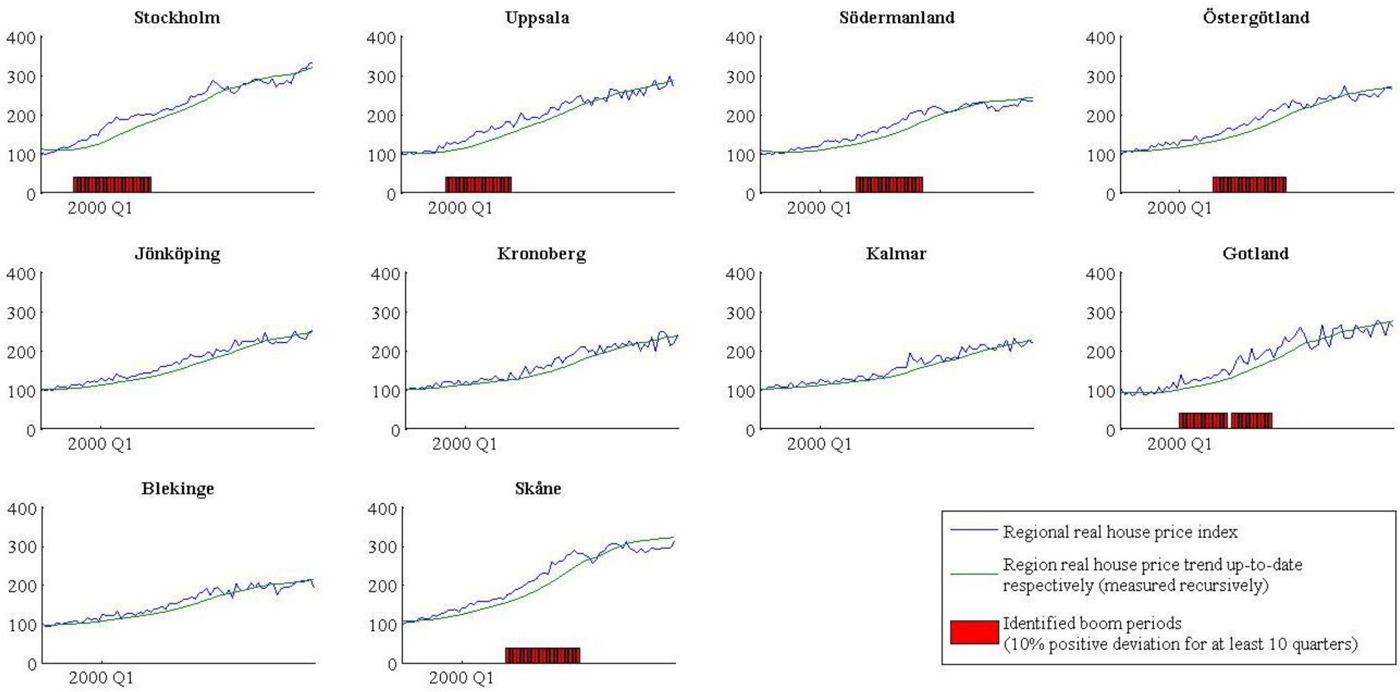
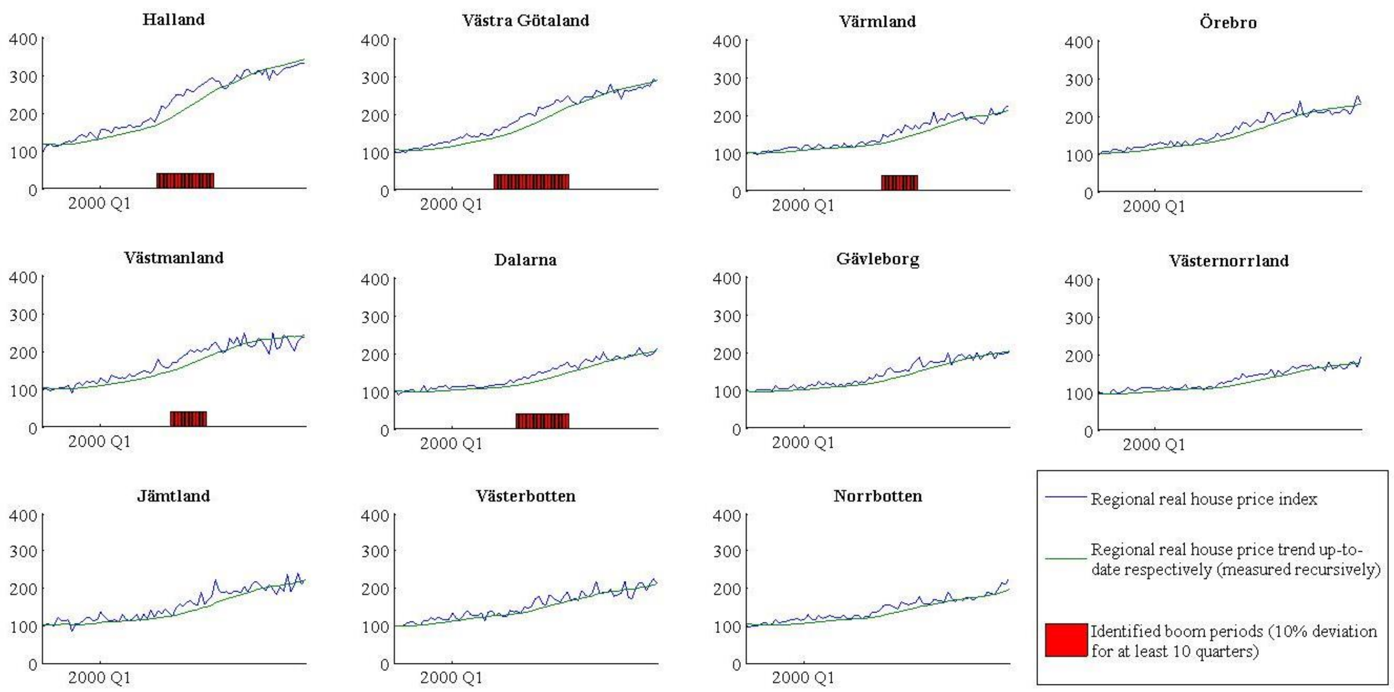


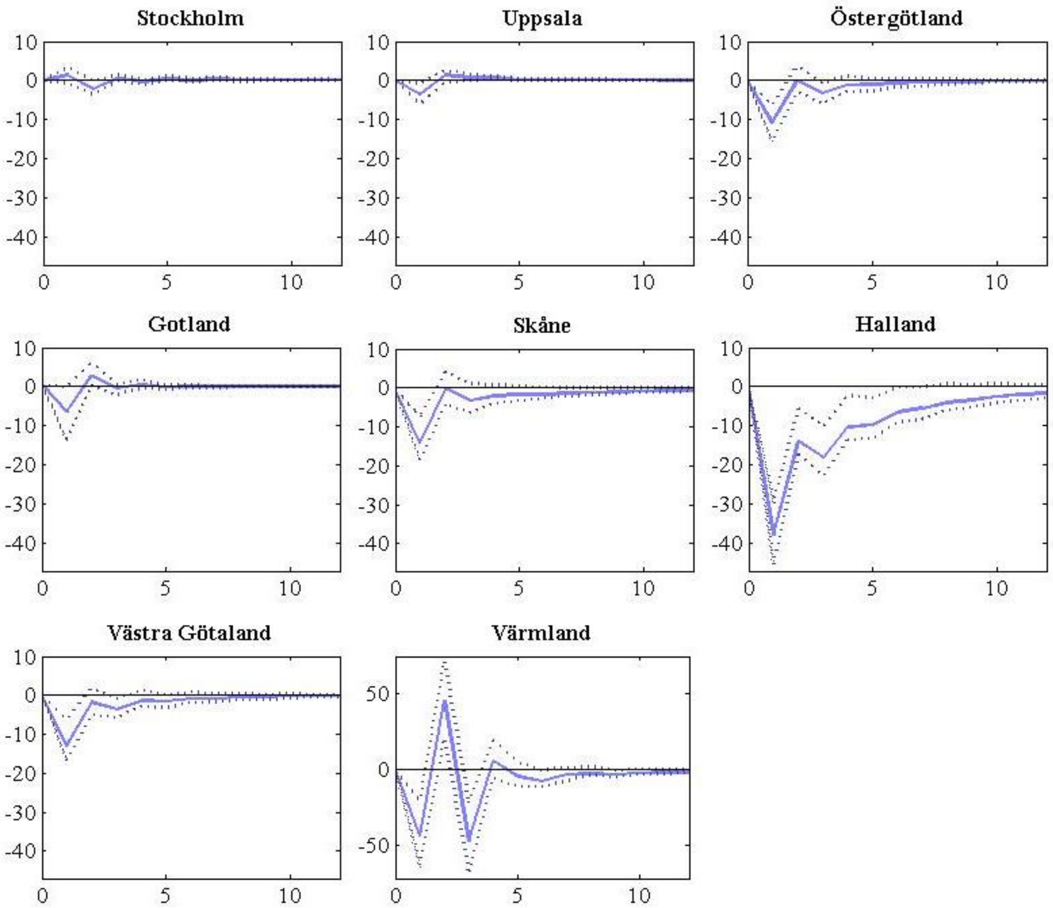
Figure 9 (b), Identified boom periods (10q-10% deviation), regional level Sweden



In 11 of the 21 regions we identify regional house price boom periods, whereas in the other regions the house price growth has tracked the smooth regional trend throughout our sample period. The impulse responses during boom versus non-boom periods for the eight regions Stockholm, Uppsala, Östergötland, Gotland, Skåne, Halland, Västra Götaland and Värmland, where both the boom and the non-boom periods allow for an impulse response analysis respectively, are shown in Figure 10 and Figure 11**.**

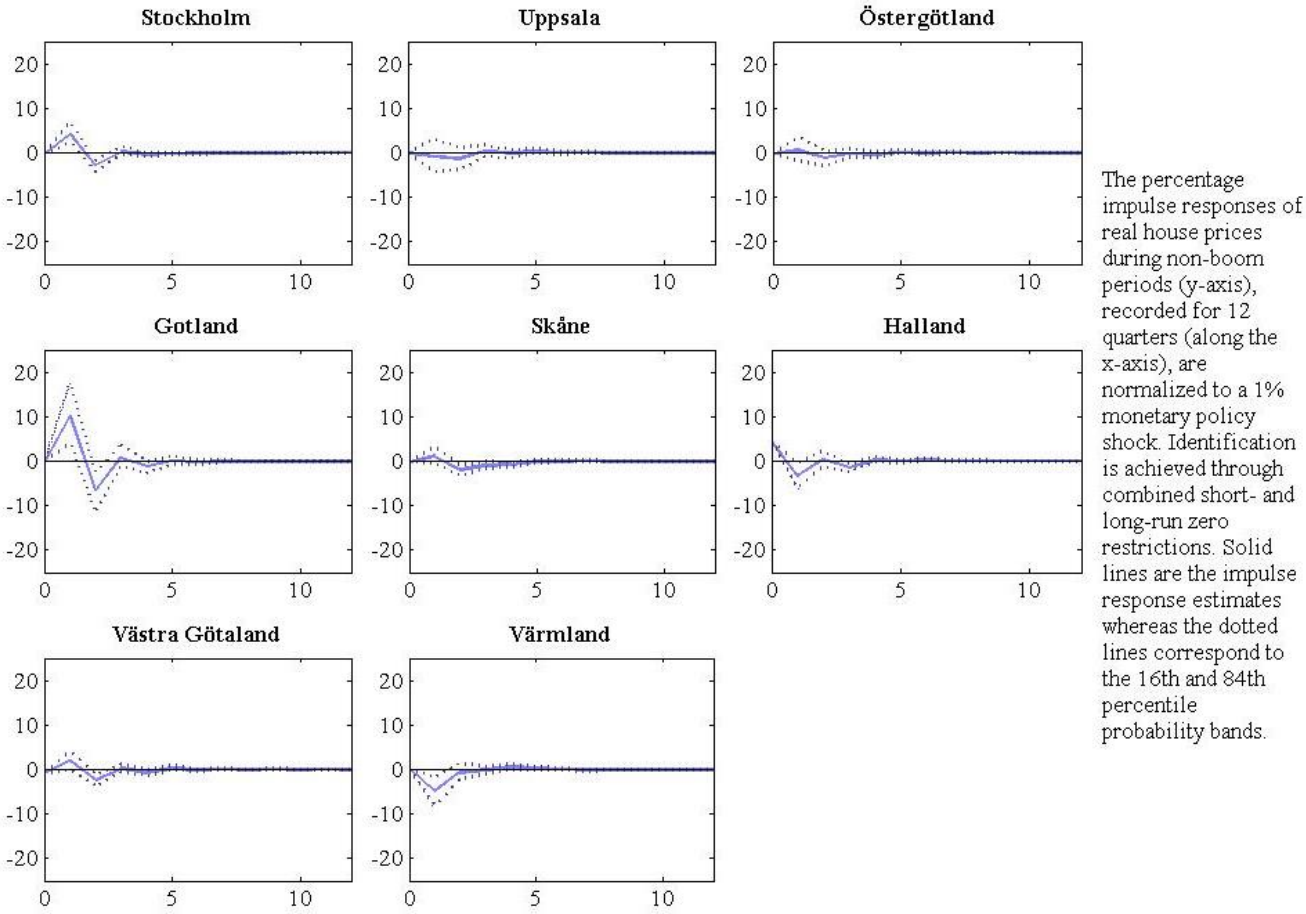
During boom periods (Figure 10) seven of the eight regions exhibit a significant negative reaction during the first quarter. The magnitude varies between -3% in Uppsala and around -40% in Värmland and Halland. Stockholm is the only region exhibiting an initial increase in house prices, but this is not significant. Compared to the impulse responses of regional house prices to a monetary policy shock for the total sample (Figure 5), a substantially larger fraction of the impulse responses in times of booms seems to be significant. Additionally, the negative downturn is also more pronounced, with a larger drop during the boom periods than for the whole sample.

Figure 10, Monetary policy shock on real house prices, boom periods (10q-10%deviation), regions - combined zero restrictions



The percentage impulse responses of real house prices during boom periods (y-axis), recorded for 12 quarters (along the x-axis), are normalized to a 1% monetary policy shock. Identification is achieved through combined short- and long-run zero restrictions. Solid lines are the impulse response estimates whereas the dotted lines correspond to the 16th and 84th percentile probability bands.

Figure 11, Monetary policy shock on real house prices, non-boom periods (10q-10%deviation), regions - combined zero restrictions



The impulse responses during non-boom periods (Figure 11) differ considerably from the ones during boom periods (Figure 10). A MWW test confirms that at a 10% level the IRFs for all eight regions are statistically not from the same population and hence significantly different, and at a 5% level at least for six of the regions. The p-values are for Stockholm 0.036; Uppsala 0.008; Östergötland 0.036; Gotland 0.081; Skåne 0.046; Halland 0.001; Västra Götaland 0.065; Värmland 0.101. During non-boom periods only house prices in Värmland fall significantly already during the first quarter; in Stockholm and Gotland, on the other hand, the response in the corresponding quarter is positive and significant, and thus comparable to the effects on a national level. Halland exhibits a contemporaneous positive response, but house prices turn negative already during the first quarter. In general, house prices in none of the regions show persistent impulse responses to the contractionary monetary policy shock during non-boom periods. Further, the magnitude of the troughs is around -2% to -5%, and thus substantially smaller than during boom periods. Hence, the regional analysis shows that a contractionary monetary policy shock does not seem to have any clear impact on real house prices during non-boom periods.

The impulse responses of house prices during boom versus non-boom periods are for most regions in line with the findings of Goodhart & Hofmann (2008), showing that monetary policy has a larger impact on house prices during boom-periods than during non-boom periods. The stronger and significantly negative effect of monetary policy during boom periods would thereby follow from the previously described more severe decrease in demand for housing during boom periods (see *6.2 Impulse responses during boom versus non-boom periods, National level*). During boom periods the initial interest payment and the following interest rate burden are based on a higher underlying value. Hence the amount that a potential house buyer is willing, and capable, to pay for a house decreases more severely during boom than during non-boom periods, leading to a stronger decline in house prices. Furthermore, an interest rate increase in boom periods might rather be interpreted as a warning signal that the economy is overheating, than as a signal for further economic growth opening up for future interest rate increases. This interpretation of the interest rate hike would then intensify the decrease in demand and prices for housing during booms.

Only in Stockholm house prices show a muted response to a monetary policy shock during boom periods and a more pronounced response in non-boom periods than for the total sample period. This is similar to the impulse responses on a national level, indicating that Stockholm is the main driver for the national results in boom versus non-boom periods. During boom periods a stronger decrease in demand for housing in Stockholm might be counterbalanced by a stronger decrease in supply of housing, motivated by loss aversion and LTV constraints. These opposite effects would then mitigate the potential impacts of monetary policy (see *6.2 Impulse responses during boom versus non-boom periods, National level*). During non-boom periods, on the other hand, an interest rate increase might be interpreted as a signal of continuing economic growth, evoking future interest rate increases to keep inflation stable. The fear of further interest rate increases then, at least initially, sustains or even boosts the demand for housing in the market. This "home prize puzzle"-effect is not only visible in Stockholm, but also in Gotland and Halland, where house prices show a significant increase as a response to an interest rate shock during non-boom periods.

To sum up, the regional impulse responses during boom and non-boom periods are mostly in line with the findings of Goodhart & Hofmann (2008); monetary policy has a larger impact on house prices during boom periods than during non-boom periods. Only Stockholm shows similar reactions as Sweden on a national level with muted IRFs during boom periods and more intensified responses during non-boom periods. The observed differences in the impact of monetary policy on house prices in boom versus non-boom periods might explain, at least partly, the heterogeneous results on a regional level for the total sample period. Most likely not all regions experience the same phases of a house price cycle in the total sample period, which in turn would trigger the heterogenenous effects of a monetary policy shock across regions.

7. Discussion and conclusion

Asset price fluctuations can have substantial effects on both the real economy and inflation. However, there is an ongoing debate on whether central banks should include asset price fluctuations in their monetary policy decisions. The Swedish Riksbank has been one of the first central banks to explicitly include rising house prices and increasing household indebtedness into interest rate decisions, thereby following a "leaning against the wind" approach. However, while both house prices and household indebtedness continue to grow, the repo rate has been decreased considerably since the second half of 2014, indicating that the latest interest rate decisions no longer reflect this stance of monetary policy. This poses the question of how monetary policy actually influences real house prices in Sweden, and hence how effective the instrument of monetary policy is in promoting financial and macroeconomic stability.

In this paper we analyze the impact of a monetary policy shock on house prices in Sweden on both a national and a regional level using a structural VAR. On a national level we find that house prices show, after an initial positive response, a significant decrease of -0.5% as response to a 1% interest rate increase. The negative response is in line with previous research, explainable by a decreased demand for housing due to the increased financing costs. The initial spike is only found by a few other papers, and might be attributable to potential house sellers' loss aversion and loan-to-value constraints, reducing the supply in the market, as well as to expectations of future interest rate increases, boosting the demand. On a disaggregated level, on the other hand, the reactions to a monetary policy shock differ considerably across regions. Even though monetary policy is directed in the same way to all regions, regional conditions seem to determine how effective interest rate increases are on a disaggregated level. Of the regions analyzed, only Stockholm exhibits a similar response pattern as observable on a national level. As one of Sweden's core regions, Stockholm probably has a considerable influence on the aggregated, national results, implying that the aggregated results might not give a representative picture of Sweden as a whole and might not capture the regional effects appropriately. The magnitude of responses to a monetary shock on a regional level seems to be larger than on a national level, indicating that in the aggregated data the actual impact of monetary policy on real housing prices might be underestimated.

When distinguishing between the impacts of an interest rate shock in boom versus non-boom periods we find a significant difference in the impulse responses of house prices. In times of booms house prices on a national level seem to be unaffected by an interest rate shock, whereas in non-boom periods the pattern is comparable to the total sample response with a positive initial spike before the impulse responses turn negative. Especially the muted effect in times of booms is surprising, as households are likely to be more highly levered during boom periods and thus to be more exposed to interest rate changes. However, it is also possible that the national results are subject to some aggregation bias as was found by Goodman (1998). When turning to a disaggregated level we find, in contrast to the national results, evidence that house prices indeed seem to be more sensitive to monetary policy shocks during house price booms than in non-boom periods, where the effect is almost muted. A contractionary policy shock during boom periods leads to a significant decrease in real house prices on a regional level.

Our findings contain some valuable insights for policy makers. First, monetary policy seems to affect real house prices in Sweden, implying that it is a tool the Riksbank could potentially use to influence the development of real house prices in combination with increasing household debt levels in Sweden. Second, the impact of a monetary policy shock varies substantially across regions; hence looking only at the response to a monetary policy shock on an aggregated, national level, might induce improper conclusions. More specifically, even though monetary policy might seem ineffective on an aggregated level, mainly driven by the diverging effects in Stockholm, it might still have a stabilizing impact on the housing market in the majority of the remaining regions. Third, a contractionary monetary policy shock seems to have only a considerable negative impact on regional house prices during local boom periods, i.e. when it is most needed, while the impact during non-boom periods is almost non-existent. This insight might be of value especially given the recent debate about whether the Riksbank or the Finansinspektion should have the main responsibility for financial stability in Sweden.

Our results are robust across all of our three employed identification methods including a Cholesky decomposition, combined short- and long-run zero restrictions, and sign restrictions. However, the sample length containing only quarterly observations is rather small. A study employing monthly data might be able to more precisely capture the impact of monetary policy decisions as these occur more frequently than on a quarterly basis. Additionally, analyses based on house prices comprising more types of housing than just one- and two-dwelling buildings for permanent living, as currently is available for a sufficiently long time series, might give more generalizable results. Since factors that influence in a considerable way the transmission channel of monetary policy, such as taxes and regulations regarding interest rate deductibility, differ substantially across countries, the results of this study should not be generalized outside of Sweden. One of the criticisms passed on to using monetary policy as a tool to mitigate asset price booms is related to the likely negative side effects it has on the economy as a whole (Jansson, 2014; Crowe, Dell'Ariccia, Igan, & Rabanal, 2013). In our model, we do not see a negative impact of the interest rate increase on GDP. However, this would have to be tested explicitly in a new model setup to allow for robust conclusions.

The paper opens up for further research on specific regional characteristics that determine the impact that monetary policy will have on a regional level. Especially the rather puzzling findings in some regions, such as in the core business region of Stockholm, might be explained by such an analysis. Additionally, further research regarding the impact of monetary policy on the fundamental value versus a potential bubble value of housing might give further insights into the effectiveness of monetary policy in Sweden. Extending our results by answering these additional questions will possibly bring forward the debate of whether and how monetary policy should be employed for promoting financial and macroeconomic stability in Sweden.

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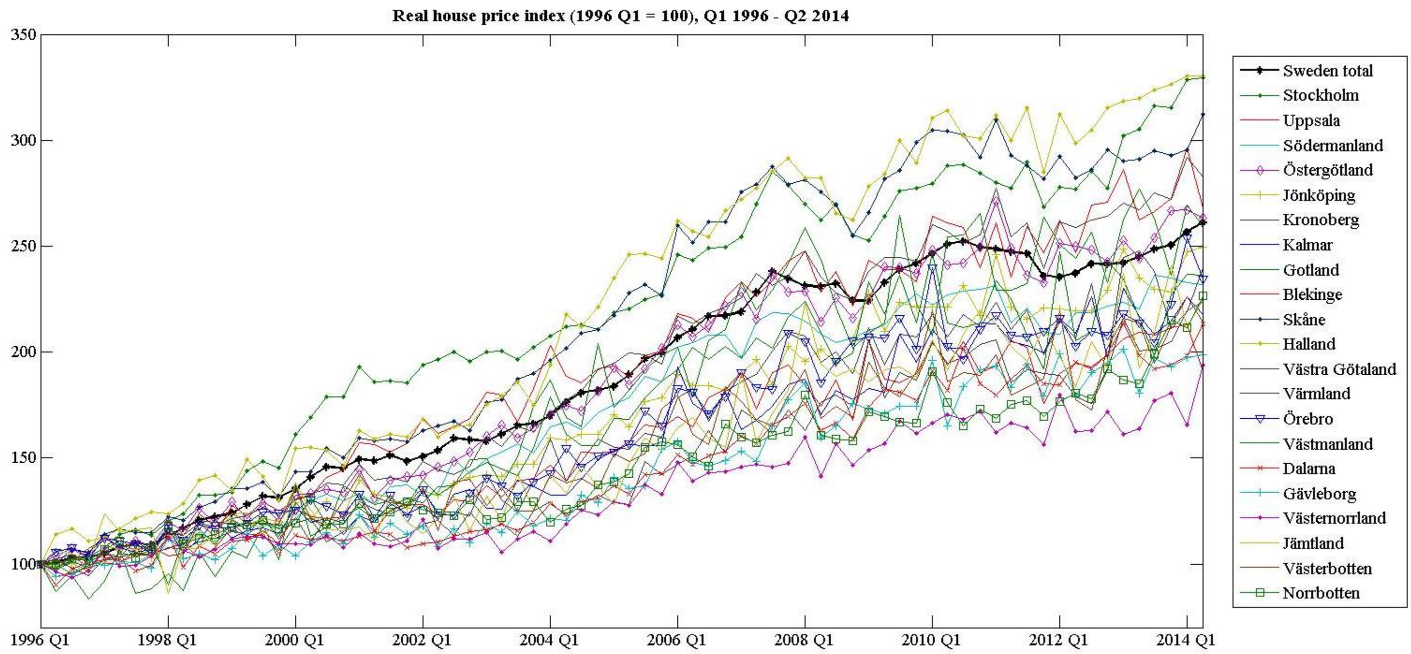
Appendix

Appendix 1, Real house price indices, Sweden total and all regions



Appendix 2, Lag length selection with information criteria, Sweden total

Information Criteria	Lag 1	Lag 2	Lag 3	Lag 4	Lag 5
Akaike Information Criterion (AIC)	-60,64	-60,82	-60,62	**-61,14**	-61,10
Schwarz Information Criterion (BIC)	**-59,51**	-58,54	-57,18	-56,52	-55,27
Hannan-Quinn Information Criterion (HQIC)	**-60,19**	-59,91	-59,25	-59,31	-58,78

Appendix 3, Regional lag length according to HQIC information criterion

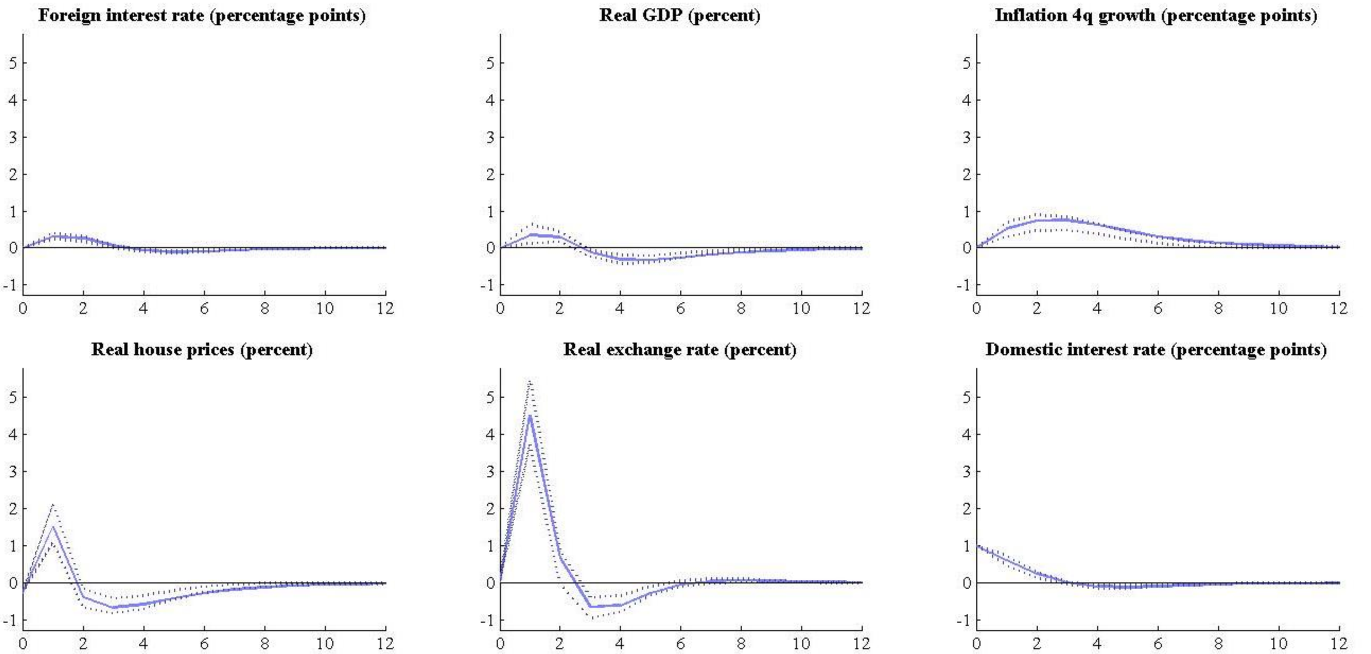
Region	Lags included according to HQIC	Region	Lags included according to HQIC
Stockholm	**1**	*Västra Götaland*	**1**
Uppsala	**1**	*Värmland*	**1**
Södermanland	**1**	*Örebro*	**1**
Östergötland	**1**	*Västmanland*	**1**
Jönköping	**1**	*Dalarna*	**1**
Kronoberg	**2**	*Gävleborg*	**1**
Kalmar	**1**	*Västernorrland*	**1**
Gotland	**1**	*Jämtland*	**1**
Blekinge	**2**	*Västerbotten*	**1**
Skåne	**1**	*Norrbotten*	**2**
Halland	**1**		

Appendix 4, Data used

Data	Defintion	Source
Domestic interest rate	3-month nominal STIBOR	Sveriges Riksbank
Real effective exchange rate (REER)	Trade weighted average of bilateral exchange rates adjusted by relative consumer prices. Based on the narrow index comprising 27 economies[1]	Bank of International Settlement (BIS)
Foreign interest rate	Trade weighted average of 3 month nominal interbank rates. Calculated applying the same trade weights as for REER[1]	Respective National Banks[1]
GDP	Real GDP, seasonlly adjusted	ECB
Regional GDP	Nominal regional GDP adjusted by the consumer price index. Interpolated to quarterly by Chow-Lin interpolation method using the Swedish national GDP as the underlying trend	Statistiska Central Byrån (SCB)
Consumer Price Index	Including all items	IMF
House prices	Owner-occupied one-and two-dwelling buildings for permanent living. Adjusted to real values using the Swedish consumer price index	Statistiska Central Byrån (SCB)

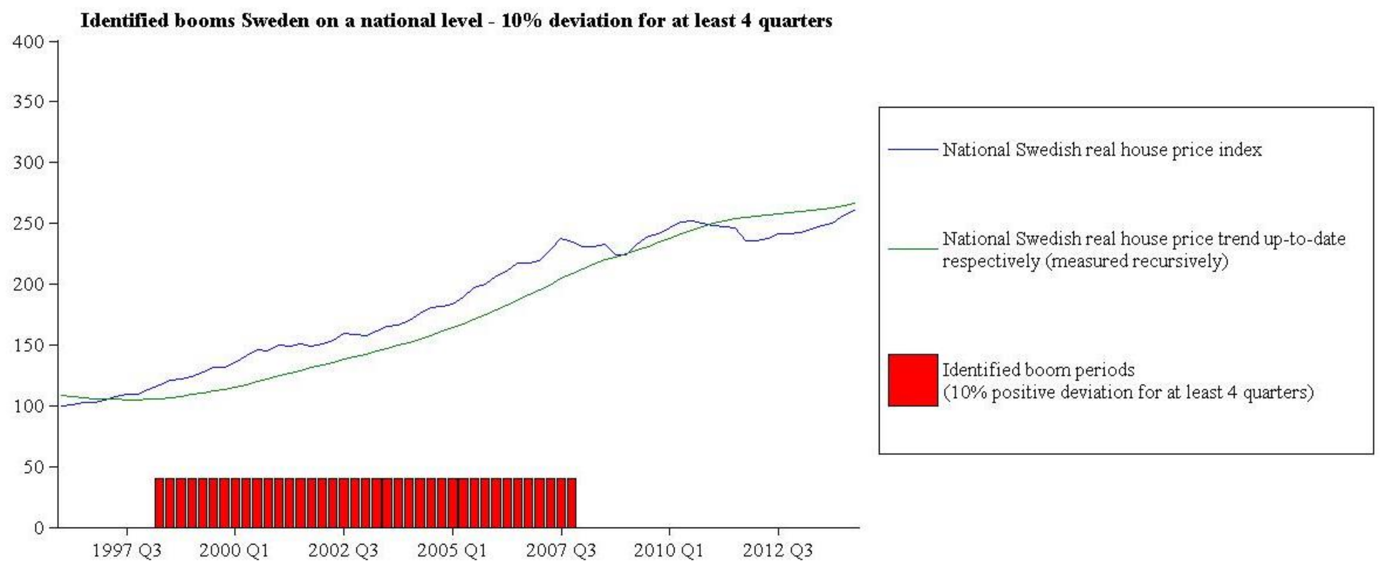
[1] Bilateral trading partners of Sweden, narrow index: Australia, Canada, Chinese Taipei, Denmark, Euro Area, Hong Kong, Japan, Korea, Mexico, New Zealand, Norway, Singapore, Swizerland, UK, USA

Appendix 5, Monetary Policy Shock - Cholesky decomposition, real house prices ordered last

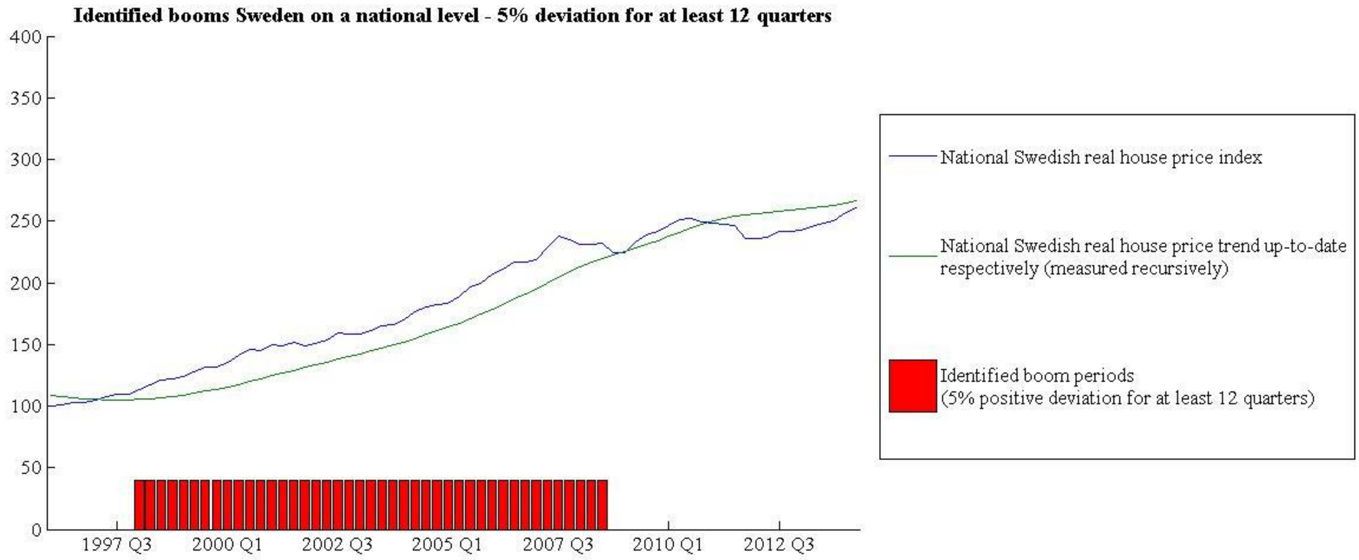


The impulse responses (y-axis), recorded for 12 quarters (along the x-axis), are normalized to a 1% monetary policy shock. Identification is achieved through a Cholesky decomposition with real house prices ordered last and domestic interest rate ordered penultimate. Solid lines are the impulse response estimates whereas the dotted lines correspond to the 16th and 84th percentile probability bands.

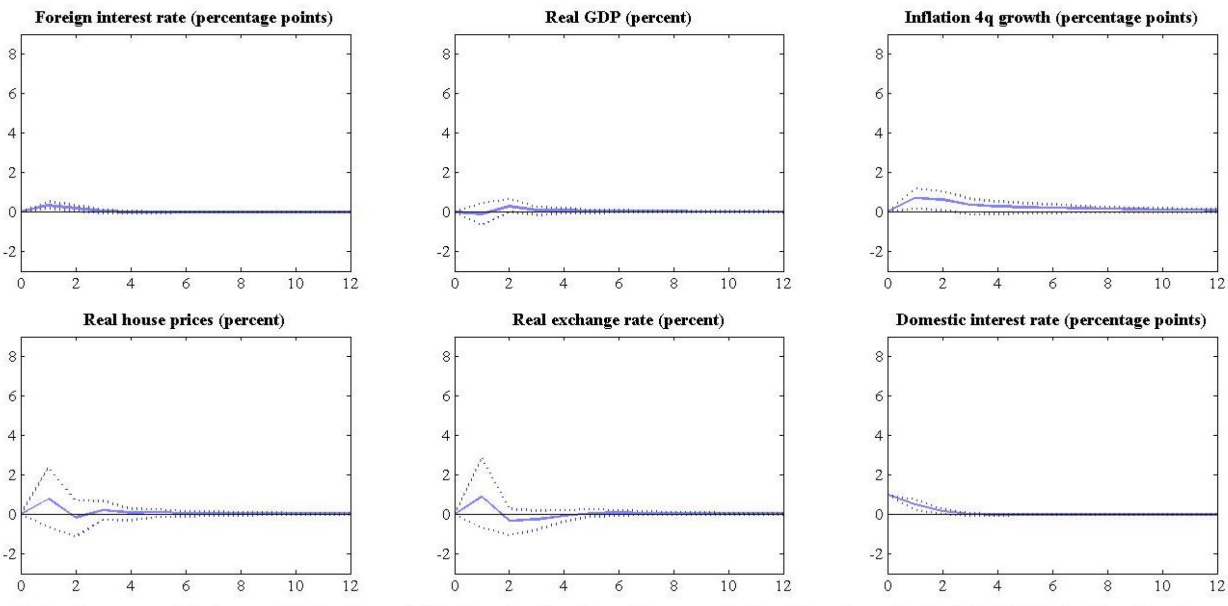
Appendix 6, Identified boom periods, Sweden at national level, 10% deviation for at least 4 quarters



Appendix 7, Identified boom periods, Sweden at national level, 5% deviation for at least 12 quarters

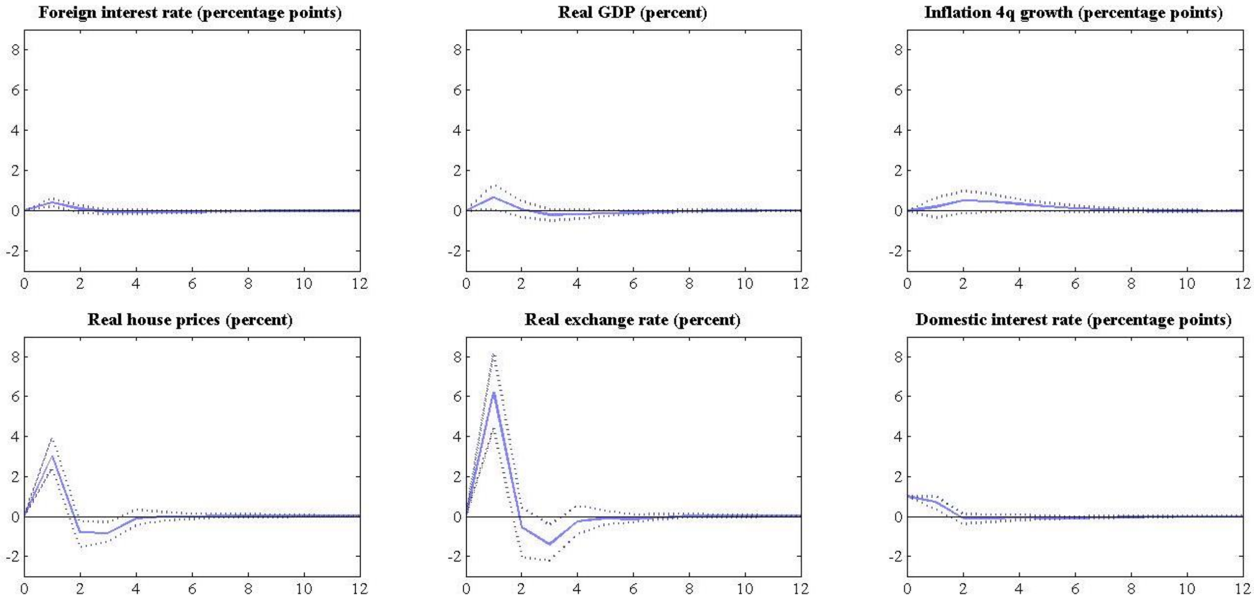


Appendix 8, Monetary policy shock - boom periods (10q-10% deviation), Cholesky decomposition, monetary policy ordered last



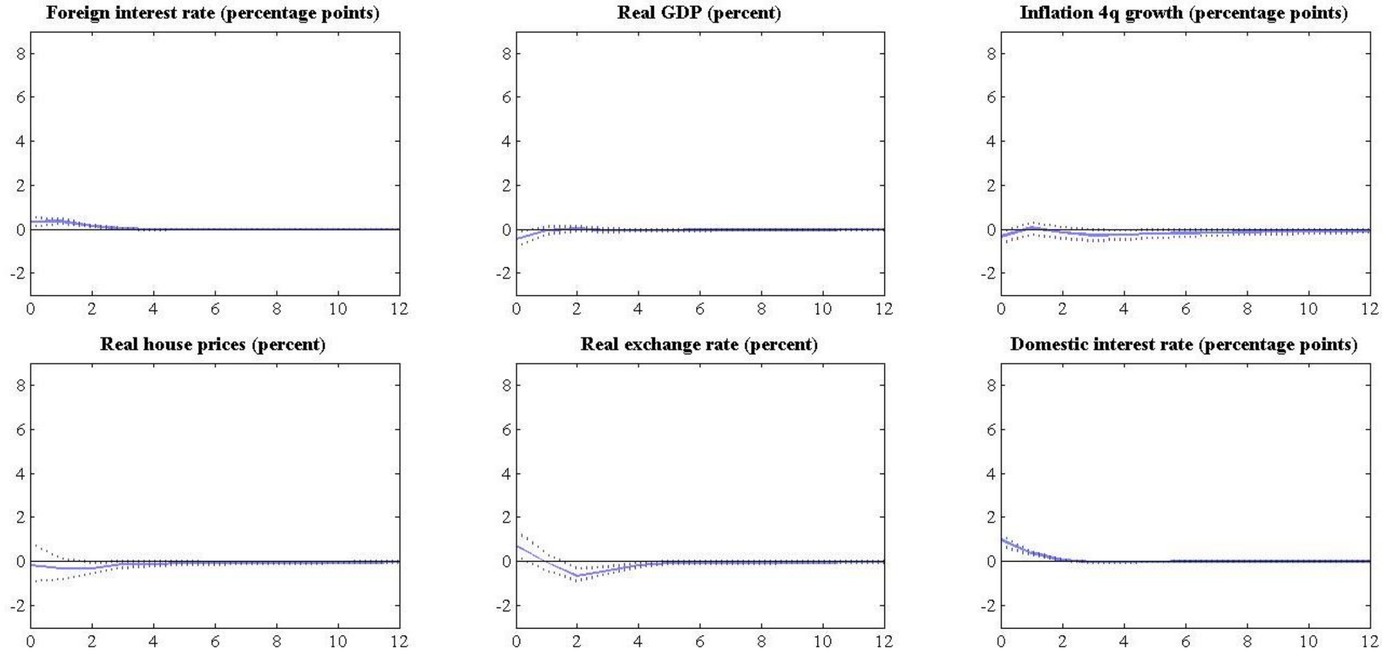
The impulse responses during boom periods (y-axis), recorded for 12 quarters (along the x-axis), are normalized to a 1% monetary policy shock. Identification is achieved through a Cholesky decomposition with monetary policy ordered last. Solid lines are the impulse response estimates whereas the dotted lines correspond to the 16th and 84th percentile probability bands.

Appendix 9, Monetary policy shock - non-boom periods (10q-10% deviation), Cholesky decomposition, monetary policy ordered last



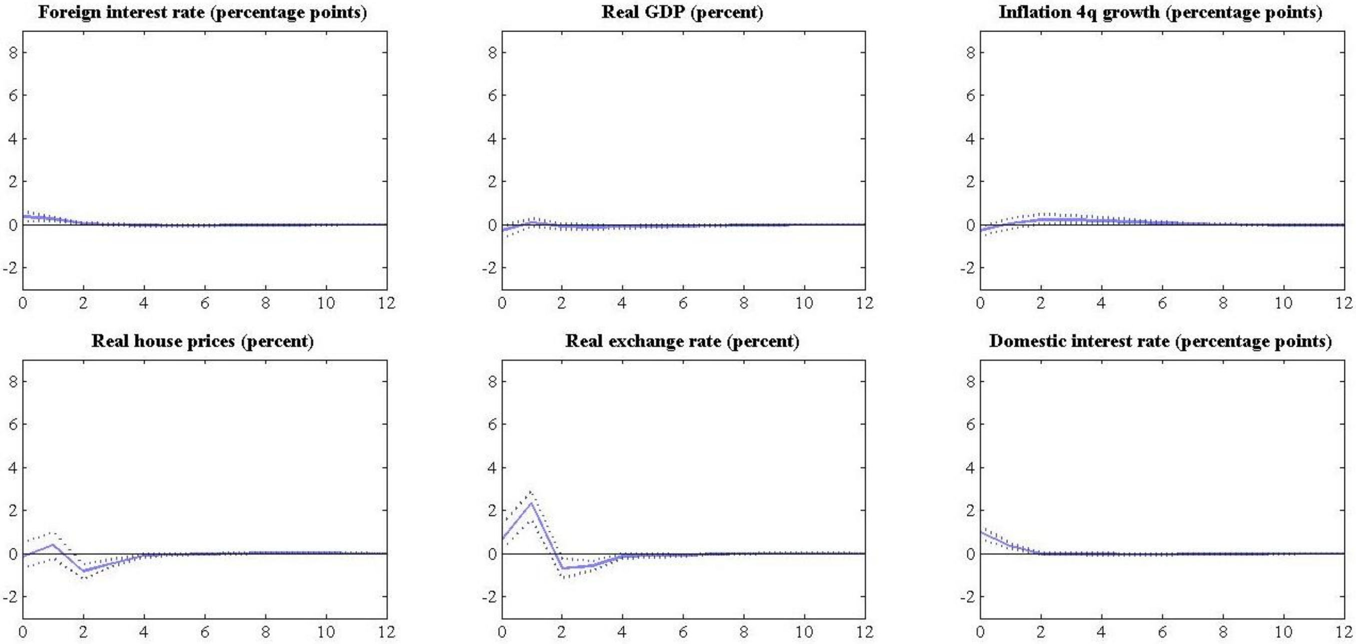
The impulse responses during non-boom periods (y-axis), recorded for 12 quarters (along the x-axis), are normalized to a 1% monetary policy shock. Identification is achieved through a Cholesky decomposition with monetary policy ordered last. Solid lines are the impulse response estimates whereas the dotted lines correspond to the 16th and 84th percentile probability bands.

Appendix 10, Monetary policy shock - boom periods (10q-10% deviation), sign restrictions



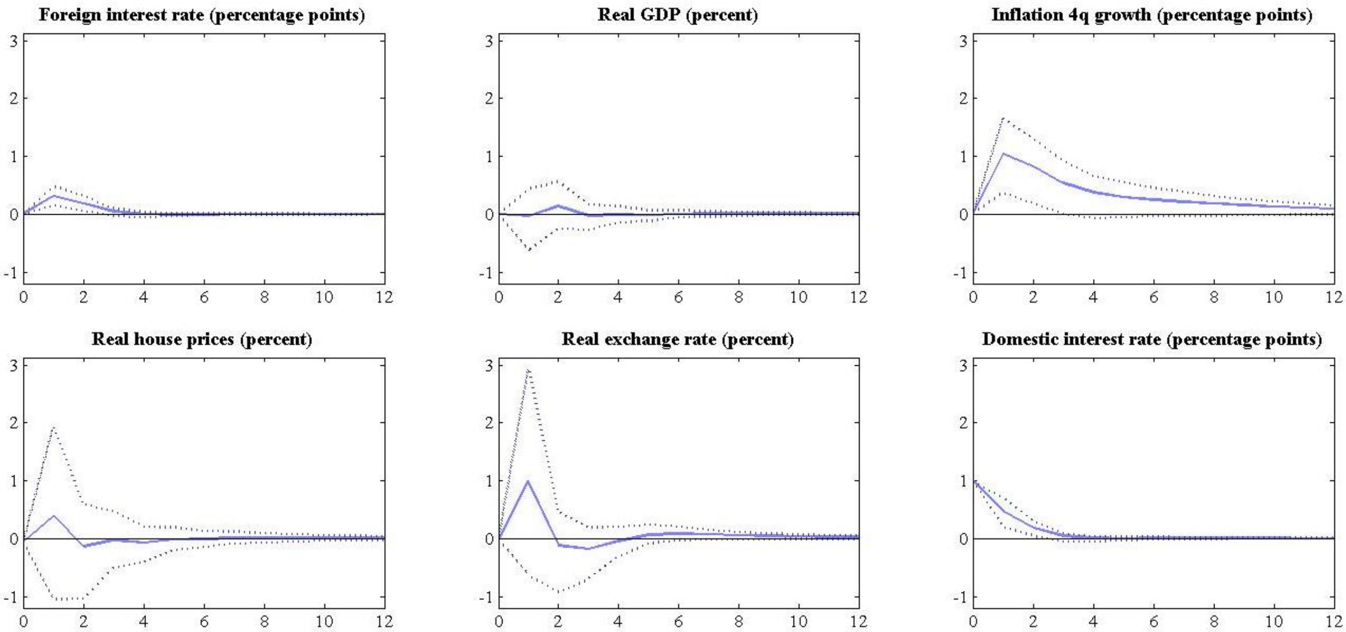
The impulse responses during boom periods (y-axis), recorded for 12 quarters (along the x-axis), are normalized to a 1% monetary policy shock. Identification is achieved through sign restrictions. Solid lines are the median estimates based on 200 valid draws whereas the dotted lines correspond to the 16th and 84th percentile probability bands.

Appendix 11, Monetary policy shock - non-boom periods (10q-10% deviation), sign restrictions



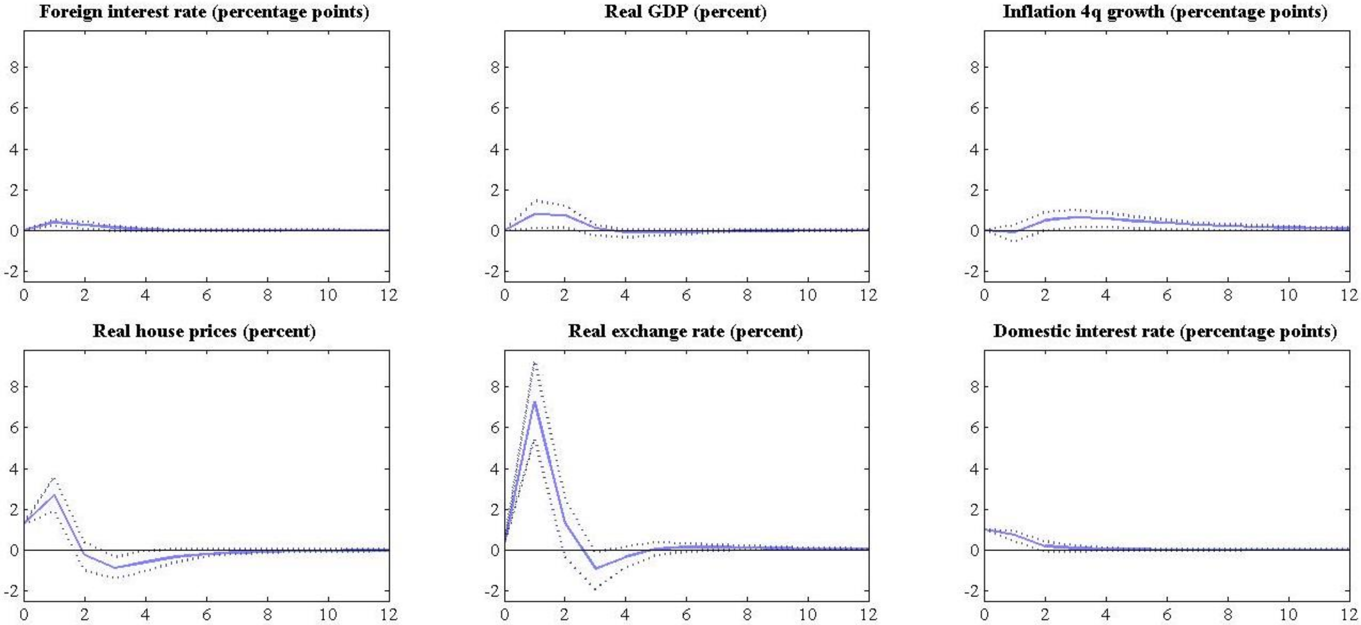
The impulse responses during non-boom periods (y-axis), recorded for 12 quarters (along the x-axis), are normalized to a 1% monetary policy shock. Identification is achieved through sign restrictions. Solid lines are the median estimates based on 200 valid draws whereas the dotted lines correspond to the 16th and 84th percentile probability bands.

Appendix 12, Monetary policy shock - boom periods (12q-5% deviation), combined zero restrictions



The impulse responses during boom periods (y-axis), recorded for 12 quarters (along the x-axis), are normalized to a 1% monetary policy shock. Identification is achieved through combined short- and long-run zero restrictions. Solid lines are the impulse response estimates whereas the dotted lines correspond to the 16th and 84th percentile probability bands.

Appendix 13, Monetary policy shock - non-boom periods (12q-5% deviation), combined zero restrictions



The impulse responses during non-boom periods (y-axis), recorded for 12 quarters (along the x-axis), are normalized to a 1% monetary policy shock. Identification is achieved through combined short- and long-run zero restrictions. Solid lines are the impulse response estimates whereas the dotted lines correspond to the 16th and 84th percentile probability bands.