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Essays on Monetary Policies, Housing Markets, and International Capital Flows

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김 형 지

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지도 교수 이 재 원

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위 원 장 김 영 식 (인)

부위원장 이 재 원 (인)

위 원 김 소 영 (인)

위 원 허 준 영 (인)

위 원 홍 정 의 (인)

Abstract

Essays on Monetary Policies, Housing Markets, and International Capital Flows

Kim, Hyungji
Department of Economics
The Graduate School
Seoul National University

This dissertation consists of two articles on monetary poicies and housing markets and one article on international capital flows. Chapter 1 analyzes how the effectiveness of monetary policy shocks are affected by leveraged housing boom period, when housing prices have surged with excessive levels of leverage. Threshold SVAR models are estimated on three small open economies - Norway, Korea, and Canada - by using minimum of the standardized real house price gap and household credit gap as a threshold variable. For all countries, the effects of monetary policy shocks on real house prices and output turn out to be more significant and stronger during the boom regime when the both real house price gap and household credit gap are above the threshold value.

Chapter 2 expands the scope of discussion into rental housing markets. In terms of monetary policy transmission mechanisms, the role of homeownership decision channels, where households could decide between mortgaged housing and rental housing, is examined focusing on sticky responses of housing rents to monetary policy shocks. A New Keynesian model incorporated with homeownership channels shows that substitution of mortgaged housing with rental housing after interest rate hikes results in smaller short-term effects of monetary policies but more persistent long-term effects. Rent rigidity, on the other hand, amplfies the short-term effect of monetary policy by suppressing this substitution, but its quantitative effect is limited and temporary.

Chapter 3 examines the effectiveness of post-AFC reforms of AFC economies, which had tightened capital controls since the Asian

Financial Crisis (AFC) to decrease the volatilities from international capital flow shocks. By classifying ASEAN+3 economies into AFC economies and Non-AFC economies, Bayesian panel VAR models are estimated on three sub-groups: (i) AFC economies in the AFC episodes, (ii) AFC economies in the GFC episodes, and (iii) Non-AFC economies in the GFC episodes. For AFC economies, the negative effects of net capital outflow shocks on real GDP growth rate during the AFC period become weaker during the GFC period. Furthermore, during the GFC episodes, AFC economies are more resilient to net capital outflow shocks than Non-AFC economies. These findings support the effectiveness of post-AFC reforms to strengthen resilience to capital flow shocks in AFC economies.

Keyword : Housing boom, Household credit, Rental housing, Monetary policy, International capital flow, Financial crisis

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Chapter 1. The effects of monetary policy during leveraged housing booms

1.1. Introduction

How do leveraged housing booms affect the effectiveness of monetary policy shocks? Since the Great Recession, which was caused by collapse of mispricing housing markets with excessive credits, the potential risk and consequences of asset bubbles with credit booms have been widely investigated. A post–crisis consensus supported by a vast literature is that leveraged housing booms, when house prices have boomed with increasing leverage, could pose a severe risk on financial stability (Jord´a et al, 2015).

Furthermore, a series of paper came along debating how central banks should deal with leveraged asset price booms, which stepped away from old beliefs that policy makers should ignore asset price cycles since they would be self–stabilizing. As potential risk of leveraged housing booms has become an unnegligible factor of macroeconomic stability, some argue that central banks should take the proactive approach towards asset price booms using the policy rate or macro–prudential tools, which is called “leaning against the wind”. In this context, the effects of monetary shocks during leveraged housing booms have attracted attention, particularly trade–offs between economic growth and financial stability.

However, a few empirical studies investigated the effects of monetary policy shocks with leveraged housing booms. Although some studies estimated the effects of monetary policy shocks with period of either credit booms or housing price booms, the paring effects of credit and housing booms on the effects of monetary policies were rarely analyzed. Historically, credit booms and housing booms haven't always came at the same time. Since the joint effects of housing booms and credit booms can be different from the effects of either of them alone, the effects of leveraged housing booms on the impacts of monetary policies should be empirically investigated.

The distinguishing feature of leveraged housing booms is that they could induce a positive feedback loop of credit and asset prices fluctuations (Mishikin, 2008, 2009). This could make them distinct from either housing booms or credit booms alone in terms of monetary policy transmission mechanisms. In fact, modest correlations between house price gaps and household credit gaps were found empirically among the smaple countries except for Korea : −0.12 for Norway, 0.81 for Korea, and 0.07 for Canada.

What could drive this discrepancy between housing and credit booms? And why do housing or credit booms alone have different effects on monetary policies than leveraged housing booms? Note that credit booms are measured using household credit to GDP ratio to capture the ability to repay (Alpanda and Zubairy, 2020). In this context, output reduction following negative shocks such as the Great Recession could lead to credit booms without housing price booms. In that cases, housing prices would not be far above their fundamental values, weakening the home equity channels. Moreover, heavily indebted households would find it difficult to take out new loans despite of interest rate cuts, leading to curtailed effects of monetary policies (Alpanda and Zubairy, 2020). On the other hand, rising economic prosperity could cause a rise in housing prices without a rapid growth of credit to GDP ratio. In those circumstances, the collateral effects of housing are limited, resulting in restricted effectiveness of monetary policies.

Jord´a et al (2015) also distinguished leveraged bubbles from asset price bubbles without credit booms, arguing that only leverage−driven bubbles matter for financial stability. He concluded that not only credit growth alone, but the interaction of credit growth and housing booms would determine the extent of the risk to financial stability. Since it is empirically evident that interaction of housing booms and credit booms plays a role in the risk on financial and macroeconomic stability, their empirical effects on leaning against the wind policies also need to be analyzed.

Figure 1. 1. Transmission Channels of Monetary Policy Shocks

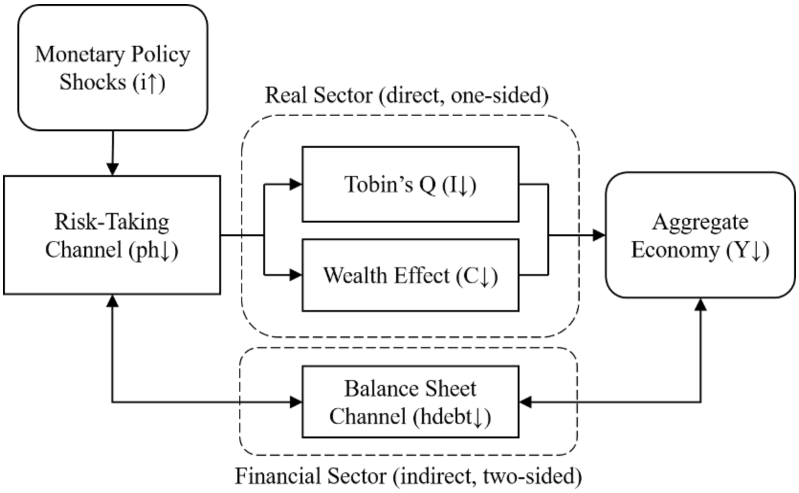


Figure 1.1 summarizes the transmission channels of monetary policies via housing markets. Housing prices serve as a medium for the transmission of monetary policies such as wealth effects or Tobin's Q (Boivin et al., 2010). There are two channels that amplify the effects of monetary policies on aggregate demands via housing prices, especially when leveraged housing booms have predominated in the housing markets: The first is a "risk taking channel". When housing prices have boomed, agents who expect future price increase are more likely to become over-leveraged with speculative motives. The excessive risk-taking behavior would pose potential risk on financial stability and trigger a sudden collapse of the housing markets for negative shocks. The second is a "balance sheet channel" (Iacoviello, 2005). In response to a fall in collateral values after monetary policy shocks, households with the high level of credits deleverage to ease debt burden, creating a vicious cycle of a further decline in house prices and aggregate demand.

The objective of this paper is to estimate the effects of monetary policy shocks during "the boom regime", where both the standardized real house price gap and the standardized household credit gap are above a certain threshold value, and compare them with those during "the normal regime", where either of them is below the threshold value. We use a dataset from three small open economies, Norway, Korea, and Canada for the period of inflation targeting framework activated.

As Jord´a et al (2015) pointed out, a bubble episode is too rare to estimate significant empirical results without being interrupted by sample size limitations. We approach this problem by dealing with a housing boom episode which is a larger concept including the bubble episode. The boom reigme is identified endogensously in our Threshold SVAR model. Instead of using the credit gap or the house price gap alone as a threshold variable, the minimum of the standardized house price gap and the credit gap is employed as a threshold variable to capture the leveraged housing boom epsiodes.

Our key finding is that the responses of real housing prices as well as output are significantly amplified during leveraged housing booms. For all sample countries, the adverse effects of monetary policy shocks on real housing prices are stronger and more significant during the boom regime. The responses of output are also more significant during the boom regime, but only in the short run. There are, however, a number of country-specific effects. First, Norway, where household credit declines severely during the boom regime, shows a more significant and persistent decrease in output than other countries where there is a temporary and insignificant decline in household credit. Second, the difference in the output responses across the regimes is much less pronounced for Canada, where most of mortgages are fixed rates. The disparity between Canada and other sample countries where variable rate mortgages are popular demonstrates that not only the quantitative but also the qualitative aspects of housing finance are important for the macroeconomic effects of monetary shocks.

To identify which channel is more important in amplifying the effects of monetary policies during the boom regime, two extended models are estimated by dividing output into each component, namely consumption and fixed capital, which capture the wealth effect and Tobin's Q effect respectively. The primary monetary policy transmission channel during the boom regime varies by country. In Norway, the stronger response of output during the boom regime is not fully explained by either fixed capital or private consumption. Instead, the key mechanism for the significant response of output in the boom regime appears to be the fall in the total level of household credit, i.e. the balance sheet effect. Meanwhile, the paring of the wealth effect and the balance sheet effect are the main channels of demand shocks during the boom regime in Korea, where a substantial proportion of household assets are in real estate. Lastly in Canada, the Tobin's Q effect accompanied with a sharp drop in real house prices is the main channel of demand shocks during the boom regimes.

To the best of our knowledge, this paper is the first to document the effects of monetary policy shocks during leveraged housing booms which are identified jointly by both the real house price gap and household credit gap. The identification of the boom regime using the pairing of the real house price and household credit gap makes an additional contribution to other studies that simply employ either of them alone as a threshold variable to estimate the effects of monetary policy shocks. In the previous literature, the impact of monetary policy shocks would become insignificant with the high level of household indebtedness. See Alpanda and Zubairy (2019) and Aikman et al. (2020). Also, stronger effects of monetary policy shocks are attained during housing booms identified by the level of house prices, but the difference in the impulse responses is not statistically significant. See Goodhart and Hofmann (2008). However, when we constrain the scope of credit booms or housing booms into leverage driven housing booms, the effects of monetary policy shocks on housing prices and output turn out to be more significant and stronger.

This paper builds upon several strands of the literature. The empirical findings of this paper, which show that housing markets become more vulnerable to monetary shocks during the boom period, support empirical literature that pointed to the interaction of credit booms and housing bubbles as the primary cause of financial instability (Jord´a et al, 2015). Furthermore, the results of this paper that output decreases more significantly with leveraged booms align the theoretical literature with respect to collateral effects of housing values in monetary policy transmission channels (Iacoviello, 2005). The country-specific difference with respect to the effect of monetary policies is consistent with the literature that structure of mortgage finance would play a role the spillover effects of housing burst into aggregate demand (Calza et al., 2013). Meanwhile, the structural identification of this paper is primarily based on Bjørnland and Jacobsen (2010), augmented with household credit to GDP ratio to incorporate credit conditions.

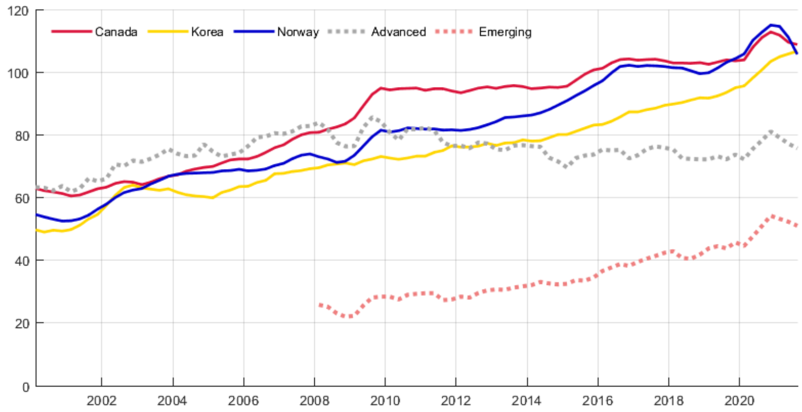
The rest of this paper is constructed as follows. In the section 1.2, data and empirical methodologies are described. In the section 1.3, the boom regimes of three small open economies identified with T-SVAR model and the state-dependent impulse responses of monetary shocks are discussed. In addition, extended models are also estimated to identify the specific channels of monetary policy transmission mechanism via housing markets during the boom regime. Section 1.4 describes some robustness test with respect to a threshold variable. Then we conclude in the section 1.5.

1.2. Data and Empirical Methodologies

1.2.1. Data and Empirical Methodologies

The cross-country analysis of this paper focuses on the three small open economies (SOEs) – Norway, Korea, and Canada. As SOEs, these economies tend to be subject to fluctuations in exchange rates and capital outflows, so that we use the structural identification of Bjørnland and Jacobsen (2010) for small open economies to estimate the effects of monetary policy shocks. Beyond the small open economy characteristic, however, it is notable that the countries have common features regarding the recent developments in the housing markets and monetary policy stance toward them.

Figure 1. 2. Household Debt to GDP Ratio, 2000:Q1~2021:Q3



First of all, they are the countries with rapidly growing household credit and housing prices. Figure 1.2 provides the household debt to GDP ratio across the countries from 2000:Q1 to 2021:Q3. The figures make it clear that these countries' household debt to GDP ratios grow at a considerable pace as they surpass the advanced economies' average since the early 2010s. Korea, in particular, has experienced the fastest growth in terms of household credit to GDP ratio since mid-2010s.

Table 1.1 shows the quarter to quarter average growth rate of real residential property prices across three economies. For Norway and Canada, house prices have increased more rapidly than the average of advanced economies sin the post-2010 sample. Although not as pronounced as in these two countries, the speed of

Table 1. 1. Average Growth Rate of Real Residential Property Prices (QoQ, %)

	Norway	Korea	Canada	Advanced	Emerging
2010:Q1 – 2021:Q3	0.81	0.42	1.04	0.63	0.43
2015:Q1 – 2021:Q3	0.72	0.72	1.41	1.02	0.37

Source: BIS.

the rising house prices in Korea stands out more for the post−2015 period than the previous one.

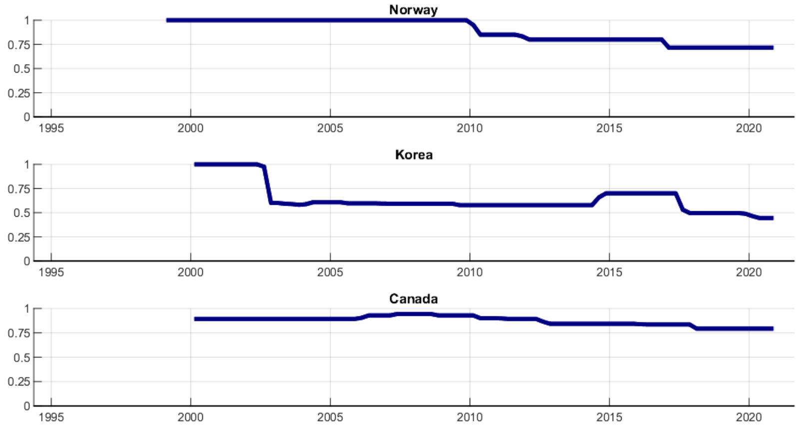
Second, they share common characteristics in terms of monetary policy framework. Specifically, these are the countries with financial stability as one of the mandates of the central bank. Monetary policy is conducted so as to mitigate the build−up of financial imbalances (Norges Bank's monetary policy strategy statement), to maintain financial stability while pursuing price promote the economic and financial welfare (the Bank of Canada Act). In this vein, there is a plethora of evidence in the monetary policy statements by these central banks regarding rising house prices and household debt have been an obvious concern for policy decision makings.

It is also noteworthy that these are the countries where unconventional monetary policies have not been implemented on a large scale such that interest−rate based policies have been practiced after the adoption of inflation targeting framework, which is essential to our structural identification which relies on short term rates to identify domestic monetary shocks.

Heterogeneous institutional characteristics in terms of housing finance across the countries would enrich our cross−country analysis. Since the effects of monetary policy shocks become more pronounced in countries with more flexible mortgage markets (Calza et al., 2013), we first focus on the prevailing interest−rate structure of mortgage contracts in each country. Fixed rate mortgages are dominant in Canada where 72 percent of total mortgages are fixed rates in February 2022. Whereas variable rate mortgages are prevailed Norway and Korea where the share of variable−rate household credit has been above 90 percent and 60 percent respectively since 2005.

Different level of macroprudential regulation for property mortgage loans across the countries might also affect the effectiveness of monetary policy shocks during the boom regime. Figure 1.3 shows the simple average of the regulatory loan−to−value (LTV) limits in each country from January 2000 to December

Figure 1. 3. LTV Ratio Cap of the Three Countries

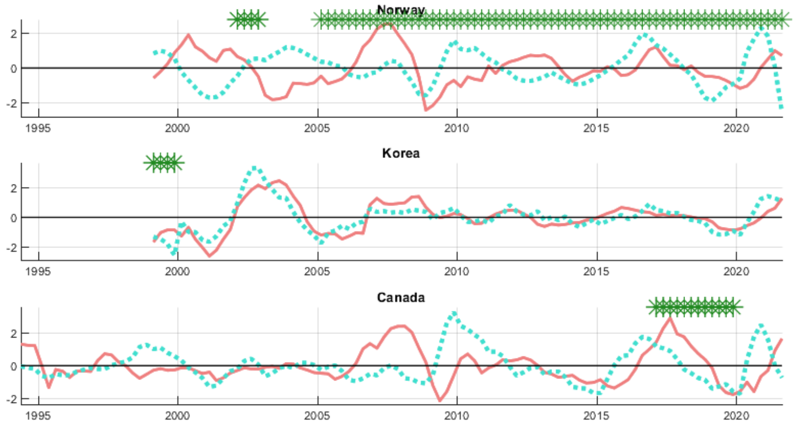


2020. While Norway and Canada have relatively high LTV limits ranging from 70 to 100 on average, Korea's LTV limits have tightened, decreasing from 100 in January 2000 to 44.4 in March 2020. Tighter LTV limits of Korea might be a result of the recent rapid increase in its household debt to GDP ratio. Lower LTV would weaken the household credit contraction during the boom regime as they would enhance borrowers' tolerance to negative housing price shocks (Jácome and Mitra, 2015).

Figure 1.4 shows the simple average measure of real estate inflow regulation for each country from 1995 to 2019, which ranges from zero to one, with a value closer to zero indicating greater openness of housing markets to foreign capital inflows (Schindler, 2009). Norway has the strictest regulation on foreign real estate inflows, which have remained unchanged at one from 2005 to 2019. Canada has recently tightened regulation on foreign real estate inflows, which became one in 2017 and maintained afterwards. Korea, in contrast, has allowed foreign capital inflows into its housing market since 2000.

Our samples run from 1999:Q1 until 2021:Q3 for Norway and Korea, and 1994Q2−2021:Q3 for Canada, following the move in each country towards an inflation targeting regime. For Canada, the starting point of the sample period is constrained by the availability of data.

Figure 1. 4. House Price Gap, Household Credit Gap and the Timing of Real Estate Inflow Restriction



Notes: In each panel, the solid and dashed lines indicate the standardized house price gap and standardized household credit gap, respectively. In each panel, the asterisks (*) indicate the period of real-estate inflow restrictions are practiced.

The benchmark VAR model in this paper is that from Bjørnland and Jacobsen (2010), which consists of six endogenous variables. In order to incorporate interactions between household credit and changes in monetary policies, we augment the model with the growth of household debt to GDP ratio. Thus our benchmark VAR model has seven endogenous variables: (i) US 10-year treasury rate (i_t^*), (ii) real GDP (Δy_t), (iii) core CPI inflation (QoQ, π_t), (iv) household credit as a percentage of GDP ($\Delta hdebt_t$), (v) real house prices (Δph_t), (vi) real effective exchange rate (ΔER_t), and (vii) policy interest rate (i_t). Notice that in order to account for the effects of US unconventional monetary policies following the global financial crisis of 2008-2009, the US 10-year Treasury rate is employed as the foreign interest rate (i_t^*), rather than the Federal Funds rate as in Bjornland and Jacobsen (2010). The detailed sources of data are reported in Appendix 1.A.

Following Bjornland and Jacobsen (2010), y_t, ph_t, ER_t and $hdebt_t$ are included in log first-differenced form for stationarity. It should be highlighted that $hdebt_t$ is measured as a percentage of GDP rather than an absolute amount. This is because the household debt to GDP ratio is more likely to capture borrowers' repayment capacity and resilience against negative house price shocks than levels. Measuring credit as a percentage of GDP is also consistent with a strand of empirical research on the effects of financial conditions (Schryder and Opitz, 2020; Büyükkarabacak and Valev, 2010; Rubaszek and Serwa, 2014).

1.2.2. Reduced-Form Threshold VAR (TVAR) model

This paper explores empirically the state-dependent effects of monetary policy shocks across two distinct housing market conditions defined as follows: (i) the boom regime, characterized jointly by house prices and household credit exceeding certain threshold values; and (ii) the normal regime, which is a complement to the boom regime. For this purpose, we employ a TVAR model as established in Tsay(1998). A reduced-form TVAR model with the boom and normal regimes is specified as follows.

$$Y_t = \left[\alpha^b + \sum_{j=1}^{l} B_j^b Y_{t-j} + u_t^b\right] S_t + \left[\alpha^n + \sum_{j=1}^{l} B_j^n Y_{t-j} + u_t^n\right](1 - S_t) \tag{1.1}$$

where the superscripts b and n denote the boom and normal regime, respectively. Y_t is the vector of endogenous variables, and $B_j^{r\prime}$s with j = 1,⋯, l and r = {b, n} are the matrices of the reduced-form VAR coefficients associated with the endogenous variables where l denotes the lag length of the VAR model. $\alpha^{r\prime}$s denote the regime-dependent constant terms, and u_t^{r}' s are reduced-form errors with $\mathrm{E}(u_t^r) = 0, \mathrm{E}(u_t^r u_t^{r\prime}) = \sum_{u,t}^{r}$, $\mathrm{E}\left(u_t^r u_s^{r\prime}\right) = 0$ for $s \neq t$ and $r \in \{b, n\}$. The TVAR model in (1.1) allows for two regimes, where the regime is determined by the level of a threshold variable (Z_t) with a delay d, Z_{t-d}, relative to an unobserved threshold level Z^*. In particular, the regimes are determined as follows:

$$S_t = 1 \iff Z_{t-d} \geq Z^* \tag{1.2}$$

In practice, the threshold value Z^* is estimated endogenously in order to minimize the residual sum of squares of (1.1). Notice that to prevent the sample size for a certain regime from being too small to estimate structural VAR models, the number of samples for each country is trimmed under certain percentages (22% for Norway, 25% for Korea, and 27% for Canada). For all the countries, the lag length is set to be two (i.e., l =2) and threshold delay is assumed to be one (i.e., d=1) following the existing literature using TVAR models (e.g., Balke(2000) and Afonso et al. (2018), among many others).

In order to identify leveraged housing booms associated jointly with house price and household credit growths, we use the minimum of the standardized real house price gap and household credit gap as the threshold variable Z_t. More specifically, the threshold variable is defined as follows:

$$Z_t = min\left[\frac{ph\,gap_t - \mu_{ph\,gap}}{\sigma_{ph\,gap}}, \frac{hdebt\,gap_t - \mu_{hdebt\,gap}}{\sigma_{hdebt\,gap}}\right] \qquad (1.3)$$

where $ph\,gap_t$ and $hdebt\,gap_t$ are the HP detrended real house prices and the household debt to GDP ratio, respectively. μ_x and σ_x denote the mean and standard deviation of a variable, respectively. Following Jord´a et al(2015), the two-sided HP filter with the smoothing parameter λ = 1,600 is employed①.

Three points are worth noting for the threshold variable as above. First, although there seems to be a lack of consensus in the existing literature regarding how to define housing booms or credit booms, a plethora of research characterizes housing price booms (credit booms) as deviation of real house prices (the household

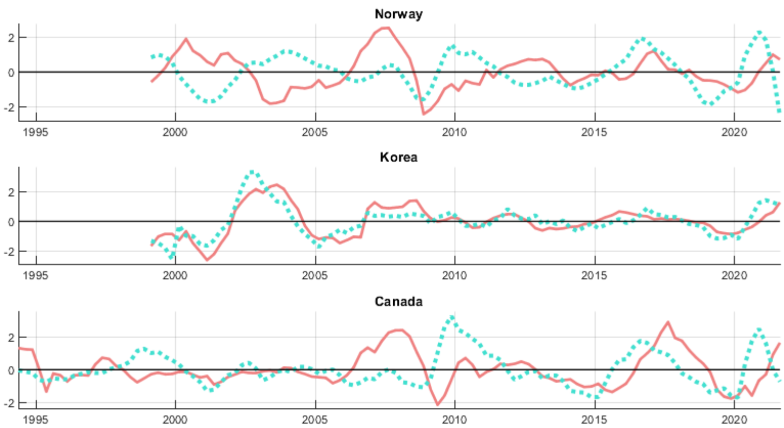
① Data used to calculate country-specific two-sided HP filters is from 1975:Q1 to 2021:Q3 and 1970:Q1 to 2021:Q3 for the household credit gap and house price gap in Norway, 1962:Q4 to 2021:Q3 and 1986:Q1 to 2021:Q3 for the household credit gap and house price gap in Korea, and 1969:Q1 to 2021:Q3 and 1970:Q1 to 2021:Q3 for the household credit gap and house price gap in Canada.

debt to GDP ratio) above some specific threshold relative to an HP filtered trend (Borio and Lowe, 2002; Detken and Smets, 2004; Goodhart and Hofmann, 2008; Jord ́a et al, 2015; Alpanda and Zubairy, 2019; Aikman et al., 2020). Thus, the use of $ph\,gap_t$ and $hdebt\,gap_t$ in (1.2) is consistent with existing literatures. Second, employing the standardized variables, instead of the variables themselves, is to prevent the threshold variable form being dominantly determined by a variable with higher volatility.

Lastly and most importantly, by using the minimum as the threshold variable, identified boom regimes are likely to rule out episodes with rising house prices or credit only. For example, the period following the Great Depression, when household credit soared due to a series of financial defaults whereas housing markets collapsed in many countries, would not be classified as the boom regime under our criteria. This criterion would help in distinguishing the effects of leveraged housing booms on financial instability from housing booms without credit booms or credit booms without housing booms.

Figure 1.5 shows the standardized state variables for each country. As made explicit in the figure, it is notable that Korea displays a comovement pattern between the house price and

Figure 1. 5. House Price Gap and Household Credit Gap



Notes: In each panel, the solid and dashed lines indicate the standardized house price gap and standardized household credit gap, respectively.

household credit gaps quite different from the other countries. For Korea, the two series tend to move together over time, whereas such a comovement is hardly observed for Norway and Canada. For instance, the correlation coefficients for each country are 0.81 for Korea, −0.12 for Norway, and 0.07 for Canada.

Based on the distinctive correlation structures between the two series across the countries, the number of regimes in the TVAR model is imposed to be country−specific. For Korea associated with a strong comovement among them, the number of regimes is set to be two. Meanwhile, three regimes are assumed for Canada and Norway. This is because, given the low or negative correlation, assuming two regimes may be too simple to capture boom regime characterized jointly by the two series. Thus, for Norway and Canada, we identify thee regimes ex ante, and classify the highest regimes as boom regime and the rest as the normal regime.

1.2.3. Structural Identification

For each regime, monetary policy shocks are identified by the methodology proposed in Bjørnland and Jacobsen (2010). The key setup of Bjørnland and Jacobsen (2010) is to allow for contemporaneous interactions between asset markets and change in monetary policies by utilizing both short− and long−run restrictions. Allowing for a full simultaneous relationship between housing prices and monetary policies could capture forward−looking behavior in asset markets, which is especially important when studying the role of house prices in the monetary transmission mechanism.

We extend the methodology so as to be suitable for augmentation with household credit as an additional variable. The crux of our identification assumption is that foreign interest rate (i_t^*), output (y_t), inflation (π_t), and household credit ($hdebt_t$) do not respond contemporaneously to monetary policy shocks, but real house prices (ph_t) and the real effective exchange rate (ER_t) do. The reduced−form VAR model for each regime with the matrix of short−run restriction $\mathbf{S}$ and the vector of structural shocks is as follows:

$$
\begin{bmatrix} i_t^* \\ \Delta y_t \\ \pi_t \\ \Delta hdebt_t \\ \Delta ph_t \\ \Delta ER_t \\ i_t \end{bmatrix} = B^r(L) + \begin{bmatrix} S_{11}^r & 0 & 0 & 0 & 0 & 0 & 0 \\ S_{21}^r & S_{22}^r & 0 & 0 & 0 & 0 & 0 \\ S_{31}^r & S_{32}^r & S_{33}^r & 0 & 0 & 0 & 0 \\ S_{41}^r & S_{42}^r & S_{43}^r & S_{44}^r & 0 & 0 & 0 \\ S_{51}^r & S_{52}^r & S_{53}^r & S_{54}^r & S_{55}^r & 0 & S_{57}^r \\ S_{61}^r & S_{62}^r & S_{63}^r & S_{64}^r & S_{65}^r & S_{66}^r & S_{67}^r \\ S_{71}^r & S_{72}^r & S_{73}^r & S_{74}^r & S_{75}^r & S_{76}^r & S_{77}^r \end{bmatrix} \begin{bmatrix} e_{i^*} \\ e_y \\ e_\pi \\ e_{hdebt} \\ e_{ph} \\ e_{ER} \\ e_i \end{bmatrix} \quad (1.4)
$$

$$\text{where } r \in \{b, n\}$$

Allowing for the contemporaneous responses of ph_t and ER_t to monetary policy shocks requires two more identification restrictions on the matrix of long-run restrictions F : (i) Monetary policy shocks have no effect on the long-run level of the real exchange rate. ($\text{i.e.}, F_{67} = 0$) (ii) Monetary policy shocks have no effect on the long-run level of real GDP. ($\text{i.e.}, F_{27} = 0$) These are the standard long-run neutrality assumptions conventionally employed in existing literature with respect to monetary policies (e.g., Obstfeld, 1985; Blanchard and Quah, 1989; Clarida and Gali, 1994).

$$
F^r = \begin{bmatrix} F_{11}^r & F_{12}^r & F_{13}^r & F_{14}^r & F_{15}^r & F_{16}^r & F_{17}^r \\ F_{21}^r & F_{22}^r & F_{23}^r & F_{24}^r & F_{25}^r & F_{26}^r & 0 \\ F_{31}^r & F_{32}^r & F_{33}^r & F_{34}^r & F_{35}^r & F_{36}^r & F_{37}^r \\ F_{41}^r & F_{42}^r & F_{43}^r & F_{44}^r & F_{45}^r & F_{46}^r & F_{47}^r \\ F_{51}^r & F_{52}^r & F_{53}^r & F_{54}^r & F_{55}^r & F_{56}^r & F_{57}^r \\ F_{61}^r & F_{62}^r & F_{63}^r & F_{64}^r & F_{65}^r & F_{66}^r & 0 \\ F_{71}^r & F_{72}^r & F_{73}^r & F_{74}^r & F_{75}^r & F_{76}^r & F_{77}^r \end{bmatrix} \quad (1.5)
$$

$$\text{where } r \in \{b, n\}$$

1.3. Empirical Results

1.3.1. Identified Regimes

Table 1.2 reports the summary statistics across the identified regimes for the three small open economies. The threshold values for Norway, Korea, and Canada are −0.49907, −0.174255, and −0.402671, respectively. The number (percentage) of samples identified as the boom regime is 39 (44%) for Norway, 44 (49%)

Table 1. 2. Summary Statistics Across the Identified Regimes

	Threshold value	Boom regime			Normal regime		
		Number of obs	Mean		Number of obs	Mean	
			Ph gap	Hdebt gap		Ph gap	Hdegt gap
Norway	−0.50	39	0.48%	0.20%	50	−0.36%	−0.19%
Korea	−0.17	44	0.72%	0.64%	45	−0.65%	−0.57%
Canada	−0.40	48	0.48%	0.24%	60	−0.43%	−0.19%

Source: Author calculations.

for Korea, and 48 (44%) for Canada, which is consistently less than half of the overall sample. Korea has the highest percentage of samples in the boom regime despite of the highest threshold value among three countries, caused mainly by the strong positive correlation between the house price gap and household credit gap.

The last six columns in Table 1.2 contrasts these two regimes by providing the summary statistics across the regimes of each country. It is noteworthy that for all the countries, the boom regimes are associated with positive house price gap and household credit gap on average. The normal regime, on the other hand, has negative mean values for both series. This joint disparity between the boom and normal regime resulting from the minimum threshold variable makes it easier to analyze the distinct effects of monetary policy shocks across the regimes.

Figure 1.6 shows the timing of the identified regimes from the TVAR model, where the boom regimes are associated with the shaded area. The characteristics of the identified regimes vary somewhat between Korea and the other two countries, originated in part by the comovement pattern of the state variables. The boom regimes are observed less frequently but with longer durations for Korea with the high degree of correlation between the house price gap and household credit gap. An opposite pattern emerges for Norway and Canada with the weak comovement between the variables, where the relatively frequent booms last shorter.

Figure 1. 6. House Price Gap, Household Credit Gap
and the Identified Boom Regime

Notes: In each panel, the solid and dashed lines indicate the standardized house price gap and standardized household credit gap, respectively. In each panel, the shaded areas indicate the timing of the boom regime identified from the TVAR model.

1.3.2. State−Dependent Impulse Response Function

Figure 1.7 shows the impulse response functions to a 100 basis points (bps) monetary policy tightening shocks associated with boom and normal regimes in Norway, Korea, and Canada. The median and one−sigma confidence intervals are reported calculated by Bayesian inferences by the Gibbs sampler introduced in Chen and Lee (1995) with 5,000 posterior draws. Cumulative responses are reported except for foreign interest rate (i_t^*), and short−term rate (i_t), which are not log first−differenced to stationarity.

It is noteworthy that we measure household credit ($hdebt_t$) as a percentage of nominal GDP, which would control borrowers' resilience against negative house price shocks better than the total level measure. Thus, the impulse response of $hdebt_t$ in our model inevitably incorporates those of output (y_t). To separate the responses of household credit from those of output, we back−calculate the responses of the total level of household credit ($hcredit_t$) by adding the responses of the y_t to the responses of the

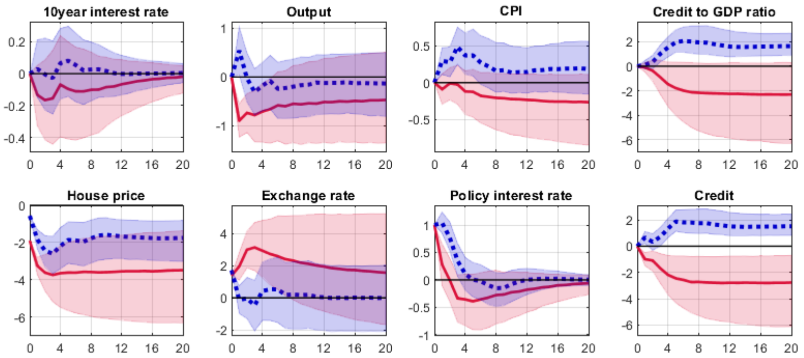
$hdebt_t$ using the relationship $hdebt_t = hcredit_t - y_t$ (in logarithmic terms) as follows.

$$Responses\ of\ hcredit_t = Responses\ of\ hdebt_t + Responses\ of\ y_t \quad (1.6)$$

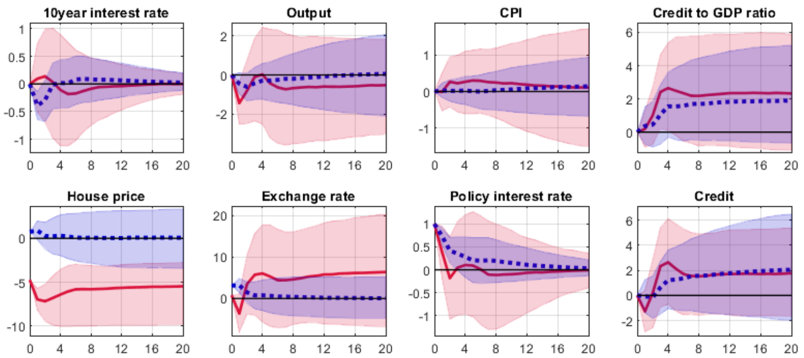
We concentrate on the response of output (y_t), real house prices (ph_t), household credit as a percentage of GDP ($hdebt_t$), and the total level of household credit ($hcredit_t$) as our primary focus is on how the effects of monetary policy shocks, particularly the trade−offs of output and house prices, are affected by the presence of leveraged housing booms. The most divergent response of output to deflationary monetary policy shocks between the boom and normal regime can be observed in Norway, where over 90 percent of mortgage loans have been variable rate since 2005. In Figure 1.7.(a), output decreases by 1 percent and remains significantly negative for a time throughout the boom regime. In contrast, during the normal regime, output rises insignificantly and only temporarily, then declines over time. Although the difference between the boom and normal regimes in terms of the median response of real home prices is only 2 percentage points, which is not that large in comparison to other countries, the response of real house prices is still more pronounced in the boom regime.

It is noteworthy that only in Norway does household credit as a percentage of GDP falls by 2 percent during the boom regime in terms of median responses. The total level of household credit also falls slightly more than 2 percent during the boom regime. In contrast, household credit as a percentage of GDP significantly rises by 2 percent during the normal regime, while the responses of total level of household credit are less than 2 percent and insignificant. We conjecture that household credit may play a role in the spillover effects of the housing bust into aggregate demand during the boom regime since that Norway has the biggest disparity between the boom and normal regime in terms of responses of output as well as those of household credit.

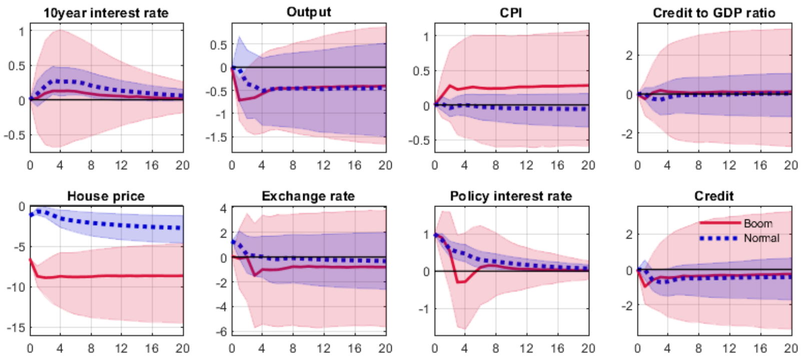
Figure 1. 7. Impulse Responses to Monetary Policy Shocks across the Regimes



Panel (a): Norway



Panel (b): Korea



Panel (c): Canada

Notes: In each panel, impulse responses associated with the boom regime (solid line with the shaded area for the median and 68% band estimates) and with the normal regime (dashed line with the shaded area) are reported. The x−axis measures quarters.

The gap between the boom and normal regimes in terms of the responses of output and real house prices is pronounced in Korea, which is characterized by a large share of variable mortgage rates and a strict level of LTV regulation. As shown in Figure 1.7.(b), the median responses of real house prices fall by 8 percent significantly and persistently, and those of output also decrease by 1 percent significantly shortly following the impact shock during the boom regime. In contrast, during the normal regime, real house prices rise modestly in the very short run and remain insignificant, with the median response hovering around 0 percent. The declining output responses under the normal regime are also less than 1 percent and insignificant over the entire period.

Household credit as a percentage of GDP increases insignificantly regardless of the regimes, which is slightly more pronounced for the boom regime. When it comes to responses to the total level of household credit, we find short-run but insignificant decreases that are consistent across both regimes. The insignificant response of household indebtedness to deflationary monetary shocks may reflect institutional characteristics of Korea, where the degree of LTV regulation had tightened over the sample periods.

Lastly, we consider the state-dependent impulse responses in Canada, which has the highest share of fixed-rate mortgage loans among three economies. As shown in Figure 1.7.(c), the median responses of output and real house prices fall for both regimes but with a different pattern. In particular, real house price declines are more pronounced in the boom regime, ranging from − 1 to 3 percent during normal regime and from − 5 to 10 percent during the boom regime. In terms of output response, output declines significantly during the boom regime, but only in the short run. In fact, the median responses of output across the regimes are similar at −0.5 percent.

Meanwhile, the responses of household credit as a percentage of GDP across the regimes are insignificant and quite similar in terms of the median estimates. In terms of the response of total level of household credit, it declines insignificantly at − 0.5 percent for the both regimes. The insignificant responses of household credit across the regimes in Canada might result from institutional characteristics of the Canadian mortgage loan market where about half of the mortgage loans consist of fixed−rate mortgages. When the fixed−rate mortgages are common, fewer borrowers would be restricted by borrowing constraints after tightening monetary policy shocks, leading to the smallest gap between the boom and normal regimes in terms of the responses of output and household credit in Canada.

Overall, we can easily find common results for the effects of monetary policy shocks during the boom and normal regimes across countries despite of some country−specific heterogeneity. First, the adverse effects of monetary policy shocks on real housing prices are stronger and more significant during the boom regime. The percentage difference in minimum response of real house prices from the boom to the normal regime in Norway, Korea, and Canada is −1.14 percentage point, −7.20 percentage point, and −6.15 percentage point, respectively. Second, the impacts of monetary policy shocks on output are more significant during the boom regime, but only in the short run for Korea and Canada. In particular, the percentage gap in the minimum response of output from the boom to normal regime in Norway, Korea, and Canada is −0.58 percentage point, −0.83 percentage point, and −0.20 percentage point, respectively.

While the effects of monetary policy shocks on output and house prices are significantly magnified in the presence of credit−fueled housing booms, the percentage loss of output per percentage change in housing prices can differ across the countries. Thus, we calculate sacrifice ratio suggested by Pascal (2020) for the boom regime, which is defined as the median percentage change in output divided by the median percent change in housing prices three years after monetary policy shocks. The sacrifice ratio during the boom regime is 0.14 for Norway, 0.11 for Korea, and 0.05 for Canada. It is noteworthy that countries with a high proportion of variable interest rates in mortgage lending have a high sacrifice ratio. This finding suggests that the trade−offs of leaning against the wind policy would be dependent not just on the prevalence of leveraged booms, but also on the structure of mortgage finance.

There are a variety of country−specific effects that can only be observed in a particular country. To begin with, Norway, where household credit declines significantly during the boom regime not only as a percentage of GDP but also at the total level, exhibits a significant and persistent decrease in output compared to other countries where household credit responds insignificantly to monetary policy shocks. Our finding that a fall in household indebtedness after house price depreciation may amplify the deflationary effects of monetary policy shocks on aggregate demand is consistent with what Iacoviello (2005) called balance sheet channels or collateral effects of housing values.

Second, in Korea, output falls stronger and more significantly during the boom regime. However, the reduction in output was only transitory despite of significant and persistent decline in real house prices for the boom regime. The short−run decline of output coincides with the transient decrease in the total level of household credit during the boom regime. Given that Korea has the strictest level of LTV regulation among the three economies across the sample period, we hypothesize that tightening LTV limits could help to attenuate the balance sheet channels that would amplify the effects of monetary policy shocks in the boom regime.

Lastly, the divergence in the output responses across the regimes is much less pronounced for Canada, where 72 percent of total mortgages are fixed rates in February 2022. In Canada, the gap between the regimes in terms of household credit as a percentage of GDP as well as at the total level is also smallest and insignificant. This finding contrasts with the significant difference in output and household credit responses across the regimes observed in Norway, where the majority of mortgages are variable rate. The disparity between Canada and Norway implies not only would the amount of household credit, but also the institutional characteristics of housing finance, play a role in the monetary policy transmission channel, as in Calza et al (2013).

1.3.3. Extended models with private consumption and fixed capitals

We show in the previous subsection that the effects of monetary policy shocks on real housing prices and output would be magnified during the boom regime. The stronger response of real house prices during the boom regime would be due to the risk-taking and balance-sheet channels, which would amplify the efficacy of tightening monetary policy shocks on housing markets by creating a vicious cycle of housing price crash and credit crunches. However, it is unknown which output components are the main sources of the stronger decline in output following monetary policy shocks during the boom regime.

In this subsection, we decompose output into two components: private consumption and fixed capital, and then estimate two extended VAR models augmented with each component to identify the main sources of stronger declines in output during the boom regime. The fluctuations in private consumption and fixed capital following monetary policy shocks could identify the two primary housing market channels of transmission mechanism in terms of output: the wealth effect and the Tobin's Q effect. The wealth effect describes the monetary policy transmission mechanism in which a housing market collapse after monetary tightening diminishes lifetime financial resources of households and reduces consumption expenditure by permanent income hypotheses (Mohanty and Turner, 2008). Meanwhile, a decline in housing prices dissuades corporations from investing more in housing construction as it costs more than its market value, which is known as "Tobin's Q effect" (Nocera and Roma, 2017).

We estimate two additional eight-variable structural VAR models where private consumption (C_t) or fixed capital (I_t) is ordered after output. Endogenous variables in the extended models are $Y_t^1 = [i_t^*, \Delta y_t, \Delta C_t, \pi_t, \Delta hdebt_t, \Delta ph_t, \Delta ER_t, i_t]$ and

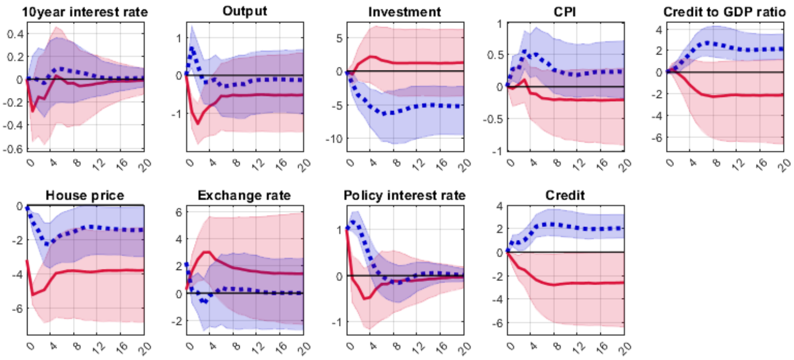
$Y_t^2 = [i_t^*, \Delta y_t, \Delta I_t, \pi_t, \Delta hdebt_t, \Delta ph_t, \Delta ER_t, i_t]$. C_t and I_t are log first-differenced and standard recursive zero restrictions are assumed on them like y_t. Despite the use of a different set of endogenous variables, the regime specification of the extended models for TVAR is identical to that of the baseline models for comparison.

By examining whether the impulse responses of the extended model with C_t or I_t reproduces similar results to the baseline model, we could identify which channel, wealth effect or Tobin's Q effect, is more important in the housing market channels of transmission mechanisms on output during the boom regime. And the rest of the results that C_t or I_t cannot explain would be relevant to financial accelerator mechanism via balance sheet channels. For each country, we find that only one of the two extended models produces the stronger responses of real house prices and output during the boom regime as the baseline model.

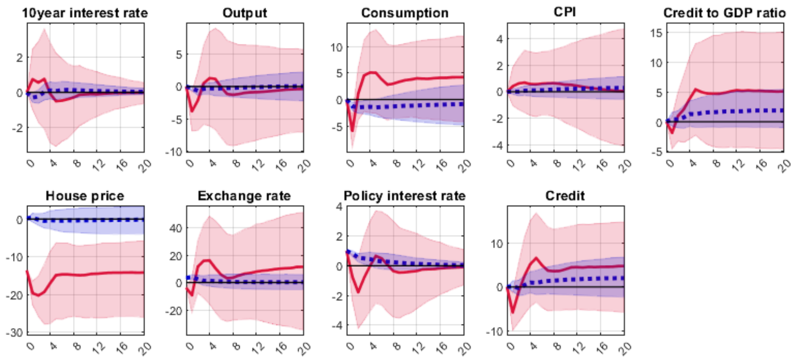
Figure 1.8 depicts the impulse response functions of the extended model, which produces the most similar results to the baseline models, in response to 100 basis point (bps) monetary policy tightening shocks associated with boom and normal regimes in Norway, Korea, and Canada. For Norway, the extended model augmented with fixed capital (I_t) provides the stronger responses of real house prices and output during the boom regime as in the baseline model. However, despite of significant decline of output during the boom regime, the negative response of fixed capital is transitory and insignificant. Instead, as in the baseline model, the total level of household credit declines solely in the boom regime, indicating that the effect of amplifying demand shocks via the balance sheet channel is more relevant in explaining a bigger decrease in output during the boom regime than the Tobin's Q effect.

In Korea, the extended model with private consumption (C_t) shows more significant responses of real house prices and output as in the baseline model. For the boom regime, the response of private consumption to monetary policy shocks is similar to that of output, which shows a sharp and significant short-term decrease. The extended model also makes the negative response of the total level

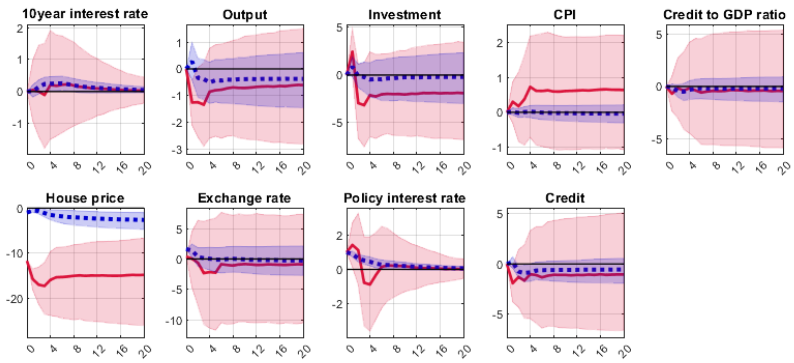
Figure 1. 8. Impulse Responses to Monetary Policy Shocks across the Regimes, Emerged from Extended Models with GDP Components



Panel (a): Norway



Panel (b): Korea



Panel (c): Canada

Notes: In each panel, impulse responses associated with the boom regime (solid line with the shaded area for the median and 68% band estimates) and with the normal regime (dashed line with the shaded area) are reported. The x−axis measures quarters.

of household credit during the boom regime more significant, albeit in the short run. In fact, 51.3 percent of Korea's total household assets were in real estate in 2018, exceeding Canada's 42.4 percent. These findings imply that in Korea, where real estate makes up the majority of household assets, the paring of wealth effects and balance sheet effects would serve as the main channels for the monetary policy transmission mechanism of demand shocks during the boom regime.

In terms of output and real housing prices, the impulse responses from the extended model with fixed capital (I_t) are the most similar to those from the baseline model in Canada. In contrast to insignificant responses of the total level of household credit during the boom regime, fixed capital declines significantly, which is consistent with the responses of output. The insignificant response of total household credit in Canada might be correlated with the fixed-rate mortgage structure as in the baseline model. This finding implies that Tobin's Q effect accompanied with a sharp house price decline would explain the majority of more significant effects of monetary policy shocks on output in Canada during the boom regime.

To summarize, we estimate the expanded models augmented with private consumption (C_t) and fixed capital (I_t) to identify the key housing market channels of the monetary policy transmission mechanism in terms of output during the boom regime. The result varies per country. The balance sheet effect is most important in Norway with respect to amplified output declines after monetary policy shocks during the boom regime. The pairing effect of the wealth effects and the balance sheet effect is at the root of the greater but transient responses of output during the boom regime in Korea, where real estate accounts for the majority of household total assets. Lastly, the Tobin's Q effect, combined with a sharp drop in housing prices, would account for more significant output responses to monetary policy shocks in Canada during the boom regime.

1.4. Robustness Tests

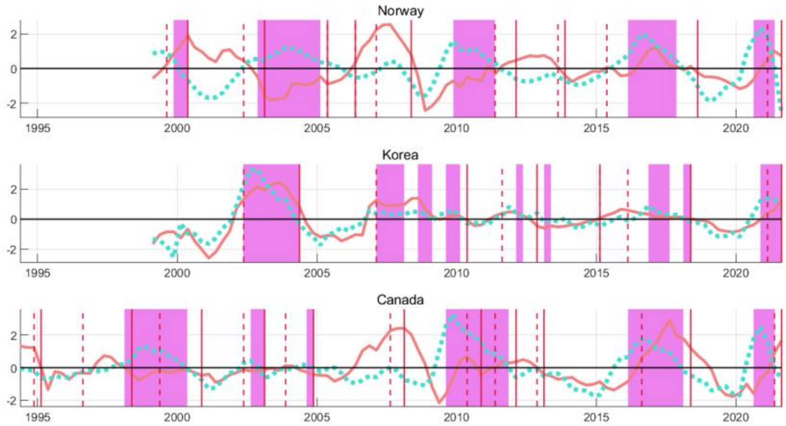
In this section, we report a variety of robustness tests, ranging from using the household credit gap alone or house price gap alone as a threshold variable instead of minimum of the standardized real house price gap and household credit gap to applying one−sided HP filter instead of two−sided HP filter to compute the house price gap and household credit gap. Overall, the robustness results offer support for our findings that the spill−over effects of monetary policy shocks on aggregate demand are significantly reinforced during the boom regimes identified by the pairing of the house price and household credit gap.

1.4.1. Using the household credit gap alone

The key result of this paper relies on using minimum of the standardized real house price gap and household credit gap as a threshold variable to identify the boom regime with leveraged booms. In fact, Jord ́a et al (2015) discovered that the coefficients on the pairing of credit growth and asset price bubbles for the probability of financial recessions are much larger than those on credit growth alone. In this context, we estimate the state-dependent impulse responses using the T-SVAR model again, this time with the household credit gap alone as a threshold variable, to confirm how the results change from those obtained by employing both the real house price gap and the household credit gap.

Figure 1.9 shows the boom regimes identified by the T-SVAR model with the household credit gap alone as a threshold variable. The detailed explanation of the model specification is provided in Appendix 1.B. The most pronounced difference from the baseline model is whether the boom regime includes the Great Recession or not. In Norway and Canada, the household credit gap surged as income fell after 2008, while the real house price gap dropped. Thus, with the household credit gap alone, the period around the Great Recession is identified as the boom regime, but not when the

Figure 1. 9. Household Credit Gap and the Identified Boom Regime

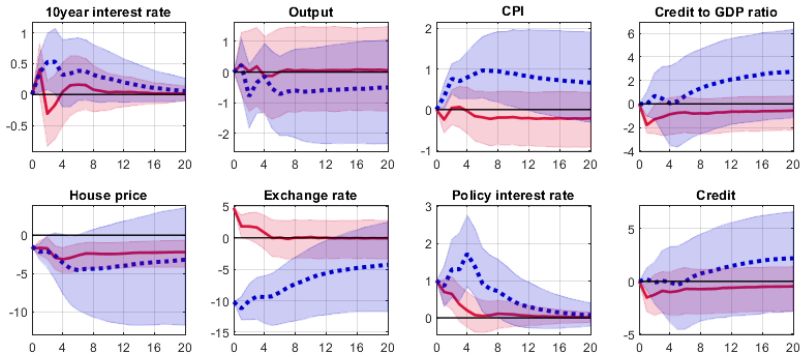


Notes: In each panel, the dashed line indicates the household credit gap. In each panel, the shaded areas indicate the timing of the boom regime identified from the TVAR model based on the household credit gap only, while the dashed and solid vertical lines indicate the starting and ending dates of the boom regime, respectively, identified from the TVAR model based both on the house price gap and household credit gap.

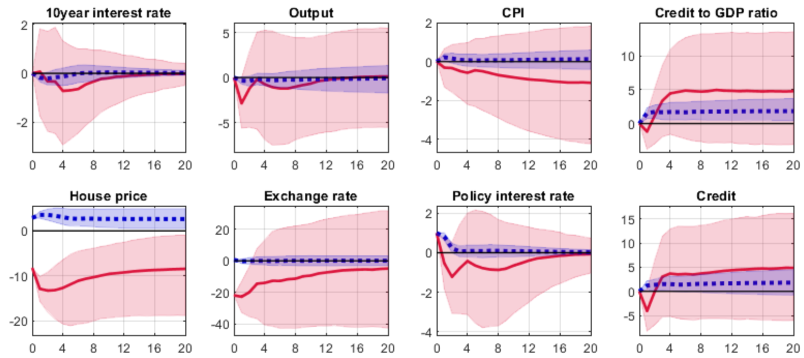
household credit gap and house price gap are combined. Meanwhile, in Korea, the disparity of the regime identification across the threshold variable is less pronounced due to high correlation of the real house price gap and household credit gap.

Figure 1.10 provides the state-dependent impulse responses of monetary policy shocks during the boom and normal regime, as identified only by the household credit gap. Arguably, the responses of output during the boom regime are insignificant for all three countries. Moreover, the median responses of output during the boom regime are much weaker than those during the normal regime for Norway and Canada. In Korea, where the correlation between the real house price gap and the household credit gap is high, the median responses of output in the boom regime decrease more dramatically, but not significantly. Meanwhile, the stronger response of real house prices during the boom regime is only observed in Korea, which is by 10 percent. For Norway and Canada, However, real house prices drop only by 2 and 3 percent respectively, which is not significantly different from the normal regime.

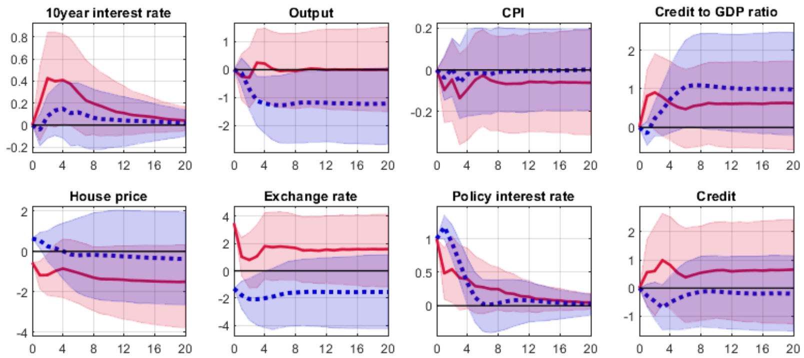
Figure 1. 10. Impulse Responses to Monetary Policy Shocks across the Regimes, when the Regimes are Identified only by the Household Credit Gap



Panel (a): Norway



Panel (b): Korea



Panel (c): Canada

Notes: In each panel, impulse responses associated with the boom regime (solid line with the shaded area for the median and 68% band estimates) and with the normal regime (dashed line with the shaded area) are reported. The x−axis measures quarters.

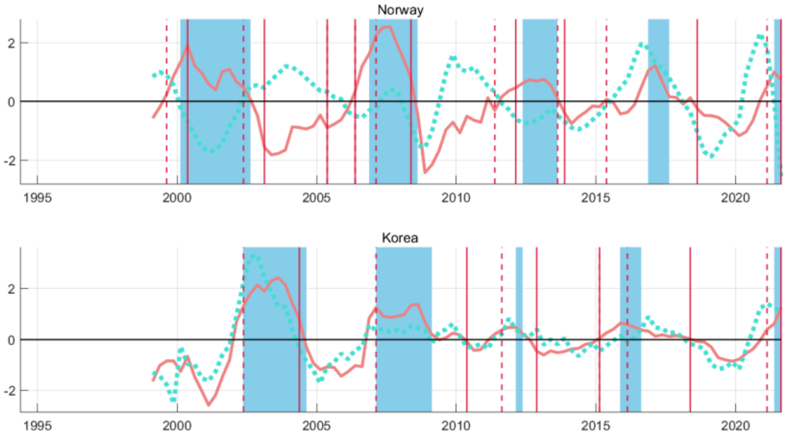
Overall, our finding suggests that when the boom regime is identified only by the level of the household credit gap alone, monetary policy shocks have insignificant impacts, especially on output. This finding is consistent with empirical research showing that the effects of monetary shocks were limited during the high debt period. See Alpanda and Zubairy (2019) and Aikman et al. (2020). Furthermore, household credit booms accompanied by asset price booms seems to have distinct effects on monetary policy transmission mechanism than the credit booms without asset booms. This finding highlights the importance of distinguishing credit booms with and without asset price booms when analyzing their effects on real economy.

1.4.2. Using the house price gap alone

A boost in economic growth may result in housing booms that are not accompanied by rapid increases in the credit to GDP ratio. A positive feedback between asset values and credit is absent in housing booms without credit booms, resulting in distinct impact on monetary policies compared to that of leveraged housing booms. We estimate the state-dependent impulse responses using the T-SVAR model with the house price gap alone as a threshold variable in order to compare the impulse responses to monetary policy shocks with those obtained by including both the real home price gap and the household credit gap.

Figure 1.11 shows the boom regimes identified by the T-SVAR model based on the house price gap alone as a threshold variable. We do not report the identified regimes for Canada, where house price increases after a contractionary monetary policy shock for any model specification. The detailed explanation of the model specification is provided in Appendix 1.C. From 2000:Q2 to 2001:Q4, Norway underwent housing booms without credit booms, as demonstrated by positive house price gaps and negative household credit gaps. Note that Norway's real GDP growth rate in 2000 was 3.2 percent, 1.2 percentage points higher than the growth

Figure 1. 11. House Price Gap and the Identified Boom Regime

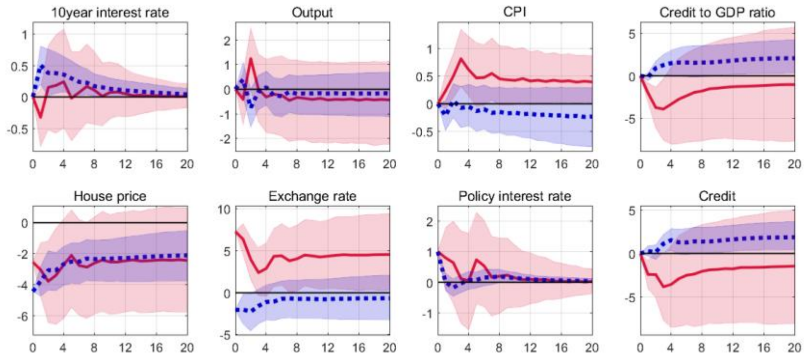


Notes: In each panel, the solid red line indicates the house price gap. In each panel, the shaded areas indicate the timing of the boom regime identified from the TVAR model based on the house price gap only, while the dashed and solid vertical lines indicate the starting and ending dates of the boom regime, respectively, identified from the TVAR model based both on the house price gap and household credit gap.

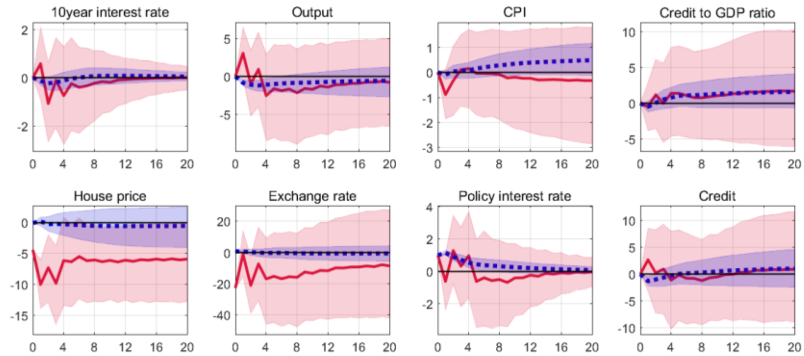
rate in 1999. In contrast, the boom regimes identified in Korea by the house price gap alone almost belong to those identified in the baseline model due to high correlation between the household credit gap and housing price gaps.

Figure 1.12 provides the state-dependent impulse responses of monetary policy shocks during the boom and normal regime, identified by house price gap alone. For all the countries, the impacts of tightening monetary policy shocks on output during the boom regime are insignificant, in contrast to the results of the baseline model where the effects of monetary policy shocks more significant during the boom regime. In Korea, the insignificant effect of monetary policy shocks during the boom regime would be attributed in part to the smaller number of samples identified as the boom regime with house price gap alone. Different impulse responses found in Norway from the baseline model, However, show that leveraged housing booms have distinct characteristics that house price gap alone cannot represent.

Figure 1. 12. Impulse Responses to Monetary Policy Shocks across the Regimes, when the Regimes are Identified only by the House Price Gap



Panel (a): Norway



Panel (b): Korea

Notes: In each panel, impulse responses associated with the boom regime (solid line with the shaded area for the median and 68% band estimates) and with the normal regime (dashed line with the shaded area) are reported. The x−axis measures quarters.

Moreover, we find that there is no significant difference in the effects of monetary policy shocks between two regimes identified by house price gap alone. This result is consistent with Goodhart and Hofman (2008) who found that although the effects of interest rate shocks became stronger during the housing booms, there were no statistically significant difference in the impulse responses between those with and without housing booms.

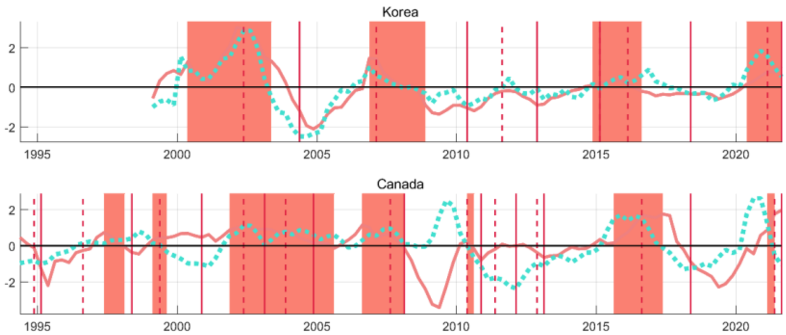
1.4.3. Using the adjusted one–sided HP filter

Our cross–country analysis measures the real house price gap and household credit gap using the two–sided Hodrick–Prescott (HP) filter, which uses future information beyond period t to construct the current trend. The two–sided HP filter is well–known for the end–point problem, in which increasing the sample size causes the trend to be revised by updating new observations. Thus, for prediction tasks or policymaking, the one–sided or real–time filter, which employs only information prior to period t, is often used. In this subsection, we use the one–sided HP filter to construct the real house price gap and household credit gap and estimate the T–SVAR model with these variables. Since the conventional one–side HP filter fails to remove low–frequency fluctuations to the same extent as the two–sided HP filter, we use an adjusted one–side HP filter with a smoothing value, λ^* = 650, and scaling factor = 1.1513 proposed by Wolf et al. (2020) to construct the trend.

Figure 1.13 shows the boom regime identified by the adjusted one–sided Hp filtered gap for all countries except Norway, where house price rises in response to a contractionary monetary policy shock for any model specification. The detailed explanation of the model specification is provided in Appendix 1.D. Note that the boom regimes identified with the adjusted one–sided HP filter tends to precede those identified with the two–sided Hp filter.

This tendency is particularly prominent in Korea, but it is also noticeable in a recent sample in Canada. This pattern of the adjusted one–sided HP filter preceding the two–sided HP filter appears to be due to the one–sided HP filter's backward–looking nature, which may capture inflection points on the cycle earlier than the two–sided HP filter. This finding implies that the adjusted one–sided HP filter is more likely to identify the boom regime as being in the early–mid stage of the boom bust cycle, whereas the two–sided HP filter is more likely to identify the boom regime as being in the mid–late stage of the boom bust cycle.

Figure 1. 13. House Price Gap, Household Credit Gap and the Timing of Boom Regime, when the One−sided HP Filter is Used for the House Price Gap and Household Credit Gap

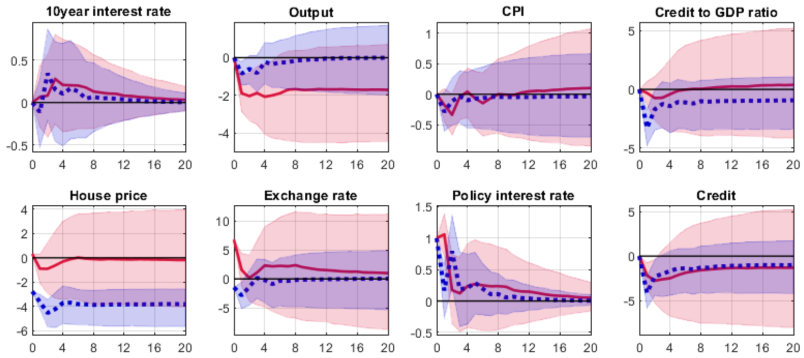


Notes: In each panel, the dashed line indicates the household credit gap. In each panel, the shaded areas indicate the timing of the boom regime identified from the TVAR model based on the house price gap and household credit gap both of which are detrended by the one−sided HP filter, while the dashed and solid vertical lines indicate the starting and ending dates of the boom regime, respectively, identified from the TVAR model based on the house price gap and household credit gap both of which are detrended by the two−sided HP filter.

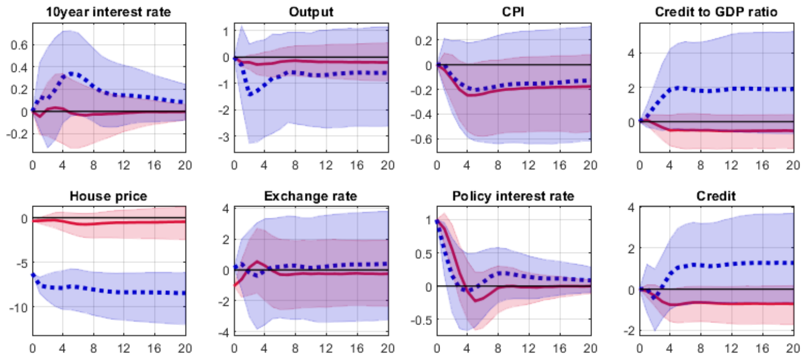
The state−dependent impulse responses estimated with the adjusted one−sided HP filtered gap are described in Figure 1.14. The response of real house prices to tightening monetary policy shocks is significantly weaker in Korea and Canada during the boom regime, contrary to the results of the two−sided Hp filter. The earlier detection of the adjusted one−sided HP filter, which primarily identifies the boom regime as the early−mid stage of the boom bust cycle, might account for the less pronounced response of real house prices during the boom regime.

The response of output, However, is more significant during the boom regime, which is consistent with the results of the two−sided HP filter. For Korea, output decreases significantly by 2 percent during the boom regime, whereas by 1 percent during the normal regime. For Canada, albeit in the short run and less pronounced, output declines significantly only during the boom regime. Overall, the adjusted one−sided HP filter tends to catch the boom regime earlier than the two−sided HP filter, causing weaker response of real house prices to monetary policy shocks during the boom

Figure 1. 14. Impulse Responses to Monetary Policy Shocks across the Regimes, when the One-sided HP Filter is Used for the House Price Gap and Household Credit Gap



Panel (a): Korea



Panel (b): Canada

Notes: In each panel, impulse responses associated with the boom regime (solid line with the shaded area for the median and 68% band estimates) and with the normal regime (dashed line with the shaded area) are reported. The x-axis measures quarters.

regime. However, our finding that the response of output to monetary policy shocks is more significant during the boom regime is well documented even with the adjusted one-sided HP filter.

1.5. Conclusion

Leveraged housing booms can impose a greater danger to financial and macroeconomic stability. Then, how do leveraged asset price booms affect the central bank capabilities to lean against the wind? In this paper, we estimate Threshold SVAR model for three small open economies – Norway, Korea, and Canada – to analyze the effects of monetary policy shocks on output and real housing prices during the boom regime identified by minimum of the standardized real house price gap and the household credit gap as a threshold variable. For all countries, we find that the effects of monetary policy shocks on real house prices more significant and stronger during the boom regime. The responses of output to monetary policy shocks during the boom regime are also more significant and stronger, albeit in the short-run in Korea and Canada. This finding suggests that the pairing of household credit and house price booms could have a distinct impact on the monetary policy transmission mechanism via housing markets than either one alone.

It is worth noting that the baseline results have some country-specific heterogeneities. For example, in Norway, where the total level of household credit falls most sharply after monetary policy shocks, the response of output during the boom period is the most significant and persistent. On the other hand, in Canada, where the majority of mortgage loans are fixed-rate, total household credit responses during the boom regime are insignificant, and there is little difference between the regimes in terms of output responses. We also estimate two extended models with private consumption and fixed capital in order to identify the primary channels of monetary policy transmission mechanism via housing markets. We discover that the main channels during the boom regime vary depending on the country: the balance sheet effect in Norway, the combination of the wealth effect and the balance sheet effect in Korea, and the Tobin's Q effect in Canada.

From a policy point of view, our finding indicates that monetary policy can be effective in collapsing housing prices, and the effectiveness would be determined not just by the quantity of household credit, but also by the extent of housing booms. Contrary to precedented literature suggesting that monetary policy interventions would become ineffective during the high debt period identified by the credit gap alone, the findings of our paper support the criticism of the classic hand–offs approach, which argue that central banks should ignore booms and clean–up after the mass. However, there may be a trade–off in the form of a more significant decrease in output in the short run, the extent of which depends on the mortgage finance characteristics of each economy. And the main transmission channels of monetary policy shocks during the boom regime would vary by country. As a result, before approaching booms in housing markets, policymakers must consider the interaction of housing booms and credit booms, as well as the unique characteristics of the mortgage markets of the economy.

Finally, in terms of central bank capabilities, our findings highlight the importance of distinguishing between a credit boom that is accompanied by an asset boom and a credit boom that is not. We find that when the boom regime is identified jointly by the house price gap and household credit gap, output and real house prices respond more significantly to monetary policy shocks, whereas the response becomes weaker and insignificant during the boom regime identified only by the household credit gap, consistent with previous researches. This finding suggests that depending on whether or not housing booms are accompanied, the high level of leverage might have a distinct impact on monetary policy transmission mechanisms.

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Appendix 1.A. Data and Sources

Variable Name	Sources	Notes
Market Yield on U.S. Treasury Securities at 10−Year Constant Maturity,	Federal Reserve Economic Data	Percent, Quarterly, Not Seasonally Adjusted
Gross domestic product − expenditure approach	OECD Statistics	National currency, volume estimates, OECD reference year, annual levels, seasonally adjusted
Private final consumption expenditure	OECD Statistics	National currency, volume estimates, OECD reference year, annual levels, seasonally adjusted
Gross fixed capital formation	OECD Statistics	National currency, volume estimates, OECD reference year, annual levels, seasonally adjusted
CPI: All items non−food non−energy	OECD Statistics	Index (2015=100)
Real house price index	OECD Statistics	Index (2015=100), the ratio of the nominal house price index to the consumers' expenditure deflator in each country, seasonally adjusted
BIS effective exchange rate	BIS Statistics Warehouse	Real (CPI−based), Broad Indices
Short−term interest rates	OECD Statistics	Percent per annum
Credit to Households and NPISHs from All sectors at Market value	BIS Statistics Warehouse	Percentage of GDP, adjusted for breaks

Appendix 1.B. Model Specification for section 1.4.1 (Household credit gap only)

	Threshold value	Boom regime			Normal regime		
		Number of obs	Mean		Number of obs	Mean	
			Ph gap	Hdebt gap		Ph gap	Hdebt gap
Norway	0.97	32	0.04%	1.96%	57	−0.04%	−1.11%
Korea	0.43	32	0.86%	0.84%	57	−0.43%	−0.42%
Canada	0.01	38	0.12%	1.00%	70	−0.10%	−0.54%

Source: Author calculations.

Appendix 1.C. Model Specification for section 1.4.2 (House price gap only)

	Threshold value	Boom regime			Normal regime		
		Number of obs	Mean		Number of obs	Mean	
			Ph gap	Hdebt gap		Ph gap	Hdebt gap
Norway	0.01	31	0.98%	−0.39%	58	−0.51%	0.18%
Korea	0.01	26	1.11%	0.78%	63	−0.41%	−0.28%

Source: Author calculations.

Appendix 1.D. Model Specification for section 1.4.3 (One−sided HP filter)

	Threshold value	Boom regime			Normal regime		
		Number of obs	Mean		Number of obs	Mean	
			Ph gap	Hdebt gap		Ph gap	Hdebt gap
Korea	−0.17	36	0.81%	0.76%	53	−0.54%	−0.48%
Canada	−0.08	43	0.66%	0.4%	65	−0.45%	−0.34%

Source: Author calculations.

Chapter 2. Homeownership Channels, Rent Stickiness, and Monetary Policy Transmission Mechanisms

2.1. Introduction

How do rental housing markets affect the monetary policy transmission mechanism via housing sector? Since the Great Recession, the role of housing sector in the transmission mechanism of monetary policy has gained much more attention in the macroeconomic literature (Iacoviello, 2005; Del Negro and Otrok, 2007; Iacoviello and Neri, 2010). However, a few studies focused on the role of rental housing markets. In fact, a large proportion of households have rented houses rather than owned houses with a mortgage. According to OECD housing tenure distribution, average 17.9 percent of households had lived in private rented accommodation from 2010 to 2020. Furthermore, actual and imputed rents for housing accounted for 26.3 percent of the core CPI index in 10 OECD economies[②] in 2021 on average, which might influence the responsiveness of inflation to monetary policy shocks.

In particular, rental housing has several distinct features from owner-occupied housing with mortgage debt, so household decisions between rental housing and mortgaged housing after monetary policy shocks could affect the monetary policy transmission mechanism, which is known as "homeownership channels" (Dias and Duarte, 2022). First, contrary to homeownership, rented houses are unable to be employed as collaterals for mortgage loans. Moreover, households could avoid the interest cost of debt by substituting mortgaged housing with rental housing. Thus, after tightening (expansionary) monetary

② Canada, Colombia, Denmark, Israel, Korea, Norway, Sweden, Switzerland, United Kingdom, and United States.

policy shokcs, the “deleveraging (leveraging) effects” will occur via homeownership decision channels, reinforcing the effects of aggregate demand shocks (Eggertson and Krugman, 2012).

Second, rental housing is less expensive than homeownership on averages[3]. This could increase (decrease) household liquidity in the short-run when households substitute mortgaged (rental) housing with rental (mortgaged) housing after monetary policy shocks. Albeit in the short-run, these “household liquidity enhancement effects” via homeownership channels can curtail the effects of monetary policies.

It is also interesting that house prices and housing rents show distinct responses to monetary policy shocks. While house prices decrease significantly following monetary policy shocks, housing rents show sticky responses (Corsetti et al., 2018; Dias and Duarte, 2019; Dias and Duarte, 2022). Theoretically, housing rents are determined by landlords who purchase houses and supply them as rental housing (Ortega et al., 2011; Rubio, 2019). Under this assumption, decreases in housing prices following monetary policy shocks lead to increased supply of rental housing, resulting in declines in housing rents, which contradicts empirical findings. The weak response of housing rents to monetary policy shocks are partly due to the fact that the majority of rental housing has been under fixed term agreements, which predetermine the fixed date for tenancy and do not allow landlords to raise housing rents without tenant agreement during those periods.

I conjecture that the rent stickiness could affect homeownership decision channels. In the presence of a rental housing market, households could shift from mortgaged to rental housing after a contractionary monetary policy shock. The substitution of mortgaged housing with rental housing increases household liquidity in the short-run, but it forces into more rapid deleveraging which would amplify the long-term effects of monetary policy shocks. In this context, when rent stickiness is

[3] According to Zillow, the U.S. average house price to rent ratio is 11.4.

prevalent, rented houses are insufficiently supplied despite of growing demands for rental housing, leading to less substitution between mortgaged and rental housing. Thus, the effects of homeownership channels on monetary policies would be curtailed in the short-run under the housing rent rigidity assumption.

The objective of this paper is summarized as threefold. First, I provide international evidences on the distinct responses of housing rents and house price to monetary policy shocks by estimating panel VARX model on 10 OECD economies. Second, I propose a two-agent New Keynesian model with rental housing market to replicate the estimated responses of housing sector to monetary policies. Third, I analyze the effects of monetary policies with and without homeownership decision channels and investigate the role of homeownership decision channels and rent stickiness.

The main finding of this paper is as follows. First, I find that after one standard deviation of monetary policy shocks, real house prices decline significantly by 0.86 percent, whereas real housing rents decrease insignificantly by 0.06 percent, which support distinct responses of housing rents and house prices using international panel data. Second, the calibrated models suggest that homeownership channels have asymmetric short- and long-term effects on monetary policies. On the one hand, substitution of mortgaged housing with rental housing after interest rate hikes result in weaker short-term effects of monetary policies on output as the household liquidity enhancement effects dominate the deleveraging effects in the short-run. On the other hand, long-term effects of monetary policy shocks on output are more persistent with homeownership channels as the presence of rental housing leads to more rapid deleveraging. Third, I find that rent stickiness plays a key role in replicating the empirical responses of housing rents to monetary policy shocks, but it has restricted and temporary effects on monetary policies at the aggregate level.

The contribution of this paper is summarized as twofold. First, to the best of my knowledge, this paper is the first to investigate the effects of homeownership channels in the monetary policy mechanism concentrating on substitution between mortgaged housing and rental housing after monetary policy shocks. In particular, I find asymmetric short-term and long-term effects of homeownership channels on the effectiveness of monetary policies at the aggregate levels, where the effects of monetary policy shocks are weaker in the short-run, but more persistent in the long-run. Second, this paper is the first to document that the macroeconomy effects of rent stickiness on homeownership channels. I find that rent stickiness has a transient and insignificant quantitative effect on monetary policy shocks, which would justify a theoretical model that does not consider the presence of rent stickiness.

The empirical analysis of this paper is directly related to literature analyzing the diversified effects of monetary policy shocks on housing prices and housing rents. Corsetti et al. (2018) estimated a factor model with high-frequency identification on 11 euro countries to find that house prices decreased by 0.4 percent to a 25bps contractionary monetary policy shock, whereas housing rents increased only by 0.05 percent. Dias and Duarte (2019) and Dias and Duarte (2022) found the similar results, where house prices fell significantly by 0.3 percent and housing rents rose insignificantly by 0.05 percent after a 25 bps contractionary monetary policy shock. I contribute to these strands of the literature by providing international evidences from 10 OECD economies.

The empirical literature investigating the presence of rent rigidity is also in line with this paper. Hoffmann and Kim (2006) showed that housing rents had the smallest frequency of price change per month (about 2 percent) among COCIP items using German individual consumer price data. Shimizu et al. (2010) found that 89 percent of rent observations had no changes in rent per year by using a micro price data set of 720,000 Japanese individuals. For the United States, Gallin and Vergrugge (2017) provided empirical evidences from Bureau of Labor Statistics (BLS) microdata that about 67 percent of rent observations remained unchanged during a 6-month period. I refer to this literature to calibrate rent stickiness in my model.

The theoretical analysis of this paper is related to the literature on the monetary policy transmission mechanism via housing sector. Iacoviello (2005) and Iacoviello and Neri (2010) proposed a standard dynamic stochastic general equilibrium framework of a New Keynesian model with housing to investigate a role of housing sector in the monetary policy transmission mechanism, but they did not consider homeownership channels. In this context, a strand of the literature tried to extend Iacoviello (2005) with homeownership decisions between owning and renting. Ortega et al.(2011) and Rubio (2019) proposed Iacoviello-type model to include rental housing markets using a CES aggregate utility function. I adopt their assumption of a CES aggregate utility function between owning and renting to analyze the effects of homeownership channels. However, Ortega et al.(2011) and Rubio (2019) predicted that housing rents would decline equally or stronger than housing prices after tightening monetary policy shocks, contradicting with the empirical results. Moreover, they analyzed homeownership channels in monetary policy transmission mechanisms focusing on the size of rental housing markets in steady state, rather than substitution between rental housing and mortgage housing as in this paper.

One of the most related to this paper is Dias and Duarte (2022), who developed a two-agent New Keynesian model with homeownership channels to replicate the empirical responses of housing rents and compared the redistributive effects of monetary policy shocks with and without homeownership channels. I adopt the assumption of a segmented housing market with nominal rent rigidities of Dias and Duarte (2022) to reproduce the estimated response of housing rents. The key difference between this paper and Dias and Duarte (2022) is that I analyze the role of homeownership channels and rent stickiness in the monetary policy mechanism focusing on substitution between mortgaged housing and rental housing. In particular, I find that borrowers can enhance liquidity by replacing mortgaged housing with rental housing after a contractionary monetary policy shocks, which Dias and Duarte (2022) did not consider.

The rest of this paper is organized as follows. Section 2.2 describes the data used, empirical methodologies and shows empirical results. Section 2.3 provides the theoretical model and calibration to replicate the empirical responses of housing sector to monetary policy shocks. Section 2.4 compares the effectiveness of monetary policy shocks with and without homeownership decision channels in terms of rental stickiness through impulse responses for monetary policy shocks. Then I conclude in section 2.5.

2.2. Data and Empirical Methodologies

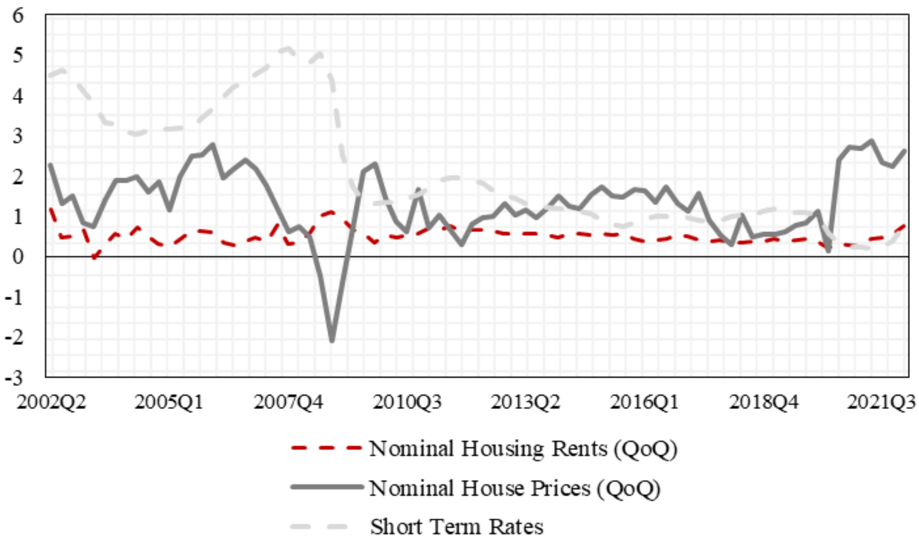
2.2.1. Data

In this section, I estimate responses of housing rents and house prices to monetary policy shocks using panel VAR models on 10 OECD economies. In particular, I focus on whether distinct responses of housing rents and house prices following monetary policy shocks, which were observed in the U.S. (Dias and Duarte, 2019, Dias and Duarte, 2022) and EU (Corsetti et al., 2018), still hold for international data. The empirical analysis of this paper covers 10 OECD economies : Canada, Colombia, Denmark, Israel, Korea, Norway, Sweden, Switzerland, United Kingdom, and United States.[④] I construct strongly balanced panels using quarterly data span from 2002:Q1 to 2022:Q1, which is the longest data available.

Figure 2.1 provides the average of the short-term rates and the average growth rate (QoQ) of house prices and housing rents in nominal terms[⑤] across the sample economies. The figure clearly

④ I exclude EU member states from samples as their monetary policies are integrated, which can violate cross-sectional independent assumption of panel VAR models.
⑤ Seasonally adjusted nominal house price index and Consumer Price Indices (CPIs) for Actual rentals for housing (COICOP 04.1) are employed to measure nominal house prices and nominal housing rents, respectively.

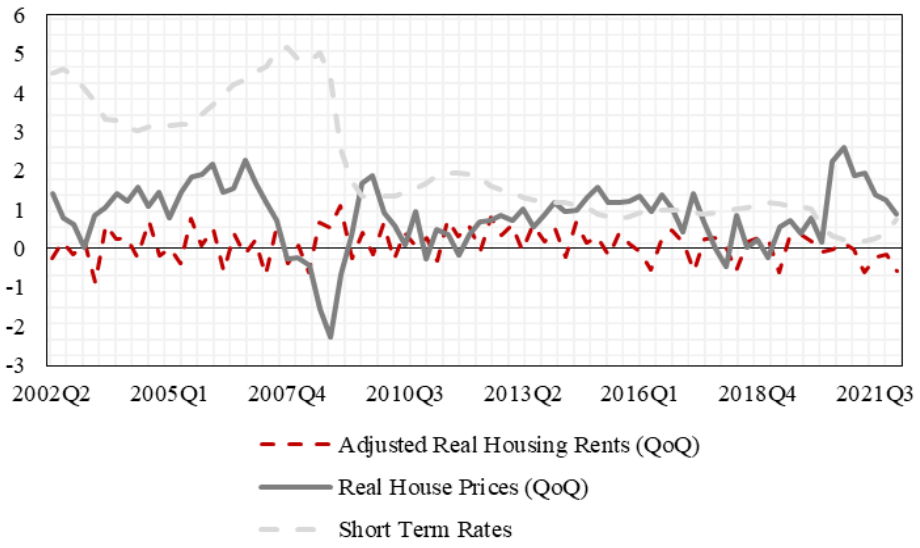
Figure 2. 1. Short term rate and growth rate of
nominal housing rents and house prices



shows that, despite volatile fluctuations in short−term rates, housing rents were more stable than housing prices. For example, the growth rate of nominal housing prices fell by up to −2.1 percent in 2008:Q4, then rose up to 2.88 percent in 2021:Q2, whereas that of nominal housing rents remained around 0.5 percent. Indeed, the standard deviation for the growth rate of nominal housing rents is 0.19, which is barely one−fourth of that for nominal housing prices of 0.83.

I also check the historical data of housing rents and housing prices in real terms. While OECD statistics provides seasonally adjusted real house price index given by the ratio of seasonally adjusted nominal house prices to the seasonally adjusted consumers’ expenditure deflator in each country, that for real housing rents is unavailable. Thus, I approximate the growth rate (QoQ) of real housing rents by subtracting the growth rate of the core CPI from the growth rate of nominal housing rent index. Since the core CPI includes rental costs, I deal with the deduplication issues by correcting them using the core CPI weight of actual and imputed rentals (denoted by $\omega_{i,t}$) as follows.

Figure 2. 2. Short term rate and growth rate of
real housing rents and house prices



$$\Delta \ln realrent_{i,t} = \frac{1}{1-\omega_{i,t}} \left(\Delta \ln rent_{i,t} - \Delta \ln coreCPI_{i,t}\right) \qquad (2.1)$$

Figure 2.2 demonstrates the average of the short-term rates and the average growth rate of house prices and housing rents in real terms across the sample economies. As in Figure 2.1, the growth rate of real housing rents showed less volatility compared to that of real housing prices. The standard deviation for the growth rate of real housing rents turn out to be 0.42, which is higher than that for nominal housing rents (0.19) but still less than that for real house prices (0.82). In both nominal and real terms, housing rents were more stable than house prices.

Table 2.1 summarizes housing tenure distribution in 2020 for sample countries except for Israel where the data is unavailable. The average homeownership rate among sample countries was 57.59 percent, which included both outright owners (without mortgages) and owners with mortgages. The average share of households who rent was 39.6 percent, which included both private and subsidized tenants. It is noteworthy that the detailed structure

Table 2. 1. Housing tenure distribution of sample OECE economies in 2020

	Own outright	Owner with mortgage	Rent (private)	Rent (subsidized)	Other, unknown
Korea	58.7%	..	32.3%	5.2%	3.8%
United Kingdom	39.3%	28.0%	11.1%	20.0%	1.6%
Colombia	31.8%	5.2%	50.6%	..	12.4%
Canada	29.9%	39.3%	30.8%	..	..
United States	25.7%	39.7%	32.7%	..	1.8%
Norway	22.2%	50.8%	23.4%	1.1%	2.5%
Denmark	15.0%	36.2%	48.6%	..	0.1%
Sweden	14.9%	43.6%	39.4%	..	2.2%
Switzerland	4.5%	33.5%	55.5%	5.7%	0.9%

Source: OECD

of the housing tenure distribution differed significantly across the sample countries. For instance, homeownership rate was the highest in Norway at 73.0 percent, 69.57 percent of which were owners with mortgages. On the other hand, the share of households who rent was the highest in Switzerland at 61.2 percent, 90.68 percent of which lived in private rented houses. In order to deal with these cross-country heterogeneity issues, I use panel VAR methods with fixed effects for empirical analysis in the next subsection.

2.2.2. Panel VARX (PVARX) model

This paper investigates the dynamics of housing rents and housing prices in responses to a contractionary monetary policy shock using international data. To control cross-country heterogeneity in 10 OECD economies, I employ panel-data vector autoregression with exogenous variables (Panel VARX) methodology in a generalized method of moments (GMM) framework following Abrigo and Love (2016). A reduced-form panel VARX model is specified as follows.

$$Y_{i,t} = Y_{i,t-1}A_1 + \cdots + Y_{i,t-p}A_p + X_t B + u_i + e_{i,t} \tag{2.2}$$

where $Y_{i,t}$ is the vector of endogenous variables for country i and period t, X_t is the vector of exogenous variables to control global factors (identical across the country), u_i are vectors of country-specific panel fixed effects, and $e_{i,t}$ are idiosyncratic standard errors obtained by the White estimator with $\mathrm{E}(e_{i,t}) = 0, \mathrm{E}(e_{i,t}' e_{i,t}) = \Sigma$, and $\mathrm{E}(e_{i,t}' e_{i,s}) = 0$ for t > s. $A_1, \ldots, A_p$ and B are reduced-form parameters to be estimated, which are assumed to be common across countries.

The specification of country-specific fixed effects with lagged dependent variables cause Nickell bias in a panel estimation (Nickell, 1981). In order to control the country-specific fixed effects, GMM estimation is employed based on forward orthogonal deviation (FOD) transformation proposed by Arellano and Bover (1995), which is given by

$$\tilde{Y}_{i,t} = (Y_{i,t} - \overline{Y_{i,t}})\sqrt{T_{i,t}/(T_{i,t}+1)} \tag{2.3}$$
$$\tilde{e}_{i,t} = (e_{i,t} - \overline{e_{i,t}})\sqrt{T_{i,t}/(T_{i,t}+1)} \tag{2.4}$$

where $T_{i,t}$ is the number of available future observations for country i at time t, and $\overline{Y_{i,t}}$ denotes the average of all available future observations for country i at the time.

The lag order of the panel estimation is set to be two. Although the optimal lag order is three according to the coefficient of determination, there is a kink in the responses of inflation in the model with three lag order due to curse of dimensionality. The responses of house price to rent ratio, housing prices, and housing rents, However, are robust in any lag order specification. The first and second lag of each endogenous variables are employed as instrument variables for FOD GMM estimation with the assumption

$\mathrm{E}(Y_{i,t-1}\check{e}_{i,t}) = \mathrm{E}(Y_{i,t-2}\check{e}_{i,t}) = 0.$

I employ seven endogenous variables to estimate the effects of monetary policy shocks on housing rents and housing prices more robustly: (i) log−differenced output ($\triangle lny_t$), (ii) first−differenced inflation ($\triangle \pi_t$), (iii) first−differenced household credit to GDP ratio ($\triangle hdebt_t$), (iv) log−differenced effective real exchange rates ($\triangle lne_t$), (v) log−differenced house prices ($\triangle \ln hp_t$), (vi) log−differenced housing rents ($\triangle lnrent_t$), and (vii) short−term rate (i_t). I estimate the models twice with nominal and real terms of house prices and housing rents. Meanwhile, the 10 minus 2 year treasury yield spread of the U.S. ($spread_t$) is employed as an exogenous variable to control the effects of unconventional monetary policies of the U.S. since the Great Recession. Appendix 2.A details the sources of the data.

All the endogenous variables except for short−term rate in my PVAR model are transformed into log−difference or first−difference to ensure stationarity. When the series has unit root, the GMM estimators suffer from the weak instruments problem (Blundell and Bond, 1998). Thus, I conduct a unit root test to guarantee that the estimation is free from unit−root problem. Table 2.2 provides the result of panel unit root tests of LLC (Levin, Lin et al. 2002). The results show that the null hypothesis of a unit root is rejected for all variables at 1 percent confidence levels.

The identification of monetary policy shocks in my panel VAR model is achieved by assuming a recursive structure on contemporaneous relationships of the endogenous variables. Specifically, I adopt commonly used Cholesky decomposition with the ordering $\{\triangle lny_t, \triangle \pi_t, \triangle hdebt_t, \triangle lne_t, \triangle lnrent, i_t, \triangle lnhp_t\}$, which imposes lagged responses of other endogenous variables except for house prices to monetary policy shocks. House prices are ordered last to assume that house prices can contemporaneously react to monetary policy shocks as in Bjørnland and Jacobsen (2010)[⑥]. On

⑥ However, I do not consider a contemporaneous effect of house prices to short−term rates as in Bjørnland and Jacobsen (2010). This is because Iacoviello (2005) pointed out that there was a little welfare gain when

Table 2. 2. The results of LLC panel unit root test

series	LLC	p−value
$\triangle\, lny_t$	−18.7741***	0.000
$\triangle\, \pi_t$	−25.1880***	0.000
$\triangle\, hdebt_t$	−4.8830***	0.000
$\triangle\, lne_t$	−15.1551***	0.000
$\triangle\, lnhp_t$ (real)	−6.5134***	0.000
$\triangle\, lnrent_t$ (real)	−11.5250***	0.000
$\triangle\, lnhp_t$ (nominal)	−5.1755***	0.000
$\triangle\, lnrent_t$ (nominal)	−4.3734***	0.000
i_t	−4.5775***	0.000

the other hand, housing rents are ordered before short−term rates as the sticky response of housing rents was identified in a strand of the literature (Hoffmann and Kim, 2006, Shimizu et al., 2010, Gallin and Vergrugge, 2017).

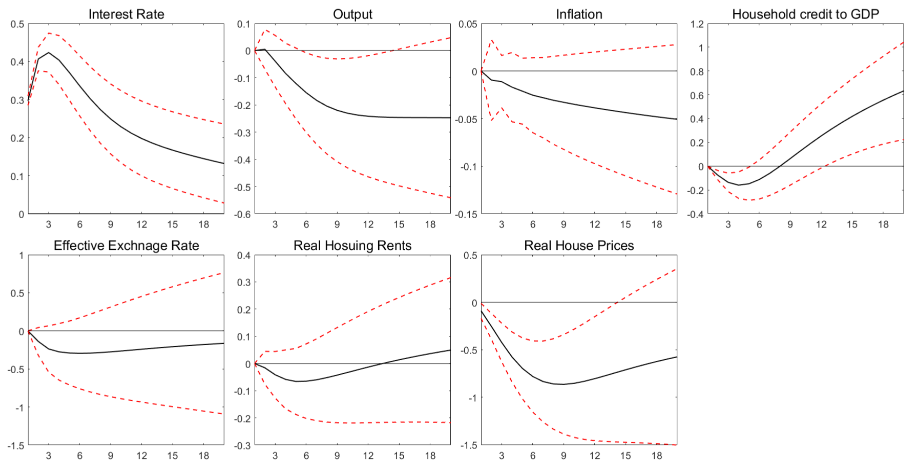
2.2.3. Responses of Real House Prices and Housing Rents After Monetary Policy Shocks

Figure 2.3 shows the impulse response functions to one standard deviation monetary policy tightening shocks for the panel VARX model including real house prices and real housing rents as a variable of interest. We report 90 percent confidence intervals calculated by Monte Carlo estimated standard errors obtained from the White estimator with 1,000 repetitions. Cumulative responses are reported for all variables except for short term rate which is not first or log−differenced for stationarity.

I focus on whether housing rents and house prices show distinct responses to monetary policy shocks. On the one hand, real house prices fall significantly by 0.86 percent after tightening monetary policy shocks. On the other hand, real housing rents show insignificant decline by 0.06 percent. It is noteworthy that the

central banks responded to asset prices. Still, the different responses of house prices and housing rents to monetary policy shocks hold for an alternative specification with the ordering $\{\triangle\, lny_t, \triangle\, \pi_t,\ \triangle\, hdebt_t, \triangle\, lne_t, \triangle\, lnhp_t, \triangle\, lnrent, i_t\}$, the results from which are provided in Appendix 2.B.

Figure 2. 3. Impulse Responses to Monetary Policy Shocks with Real Housing Rents and House Prices



Notes: Dashed lines show 90% confidence intervals calculated by Monte Carlo estimated standard errors obtained from the White estimator with 1,000 repetitions. The x-axis measures quarters.

degree of response of real housing rents in terms of point estimates is about tenth of that of real housing prices. The estimated responses of housing rents and housing prices from 10 OECD economies are in line with Corsetti et al. (2018), Dias and Duarte (2019), and Dias and Duarte (2022) who found that housing rents showed much smaller responses compared to real estate prices after monetary policy shocks using the EU and U.S. data, respectively.

Furthermore, it is noteworthy that other macro variables, including output, inflation, and the household credit to GDP ratio, show significant responses that are in consistent with standard macroeconomic theories[⑦]. In particular, output declines significantly by 0.25 percent, inflation rate decreases by 0.05 percentage points, and household credit to GDP ratio falls significantly by 0.16 percentage points in the short run. Although responses of inflation and real effective exchange rate are insignificant at a 90 percent

⑦ The long-run increase of household credit to GDP ratio after contractionary monetary policy shocks is due to decline of output over the same period as we measure household credit as a percentage of GDP.

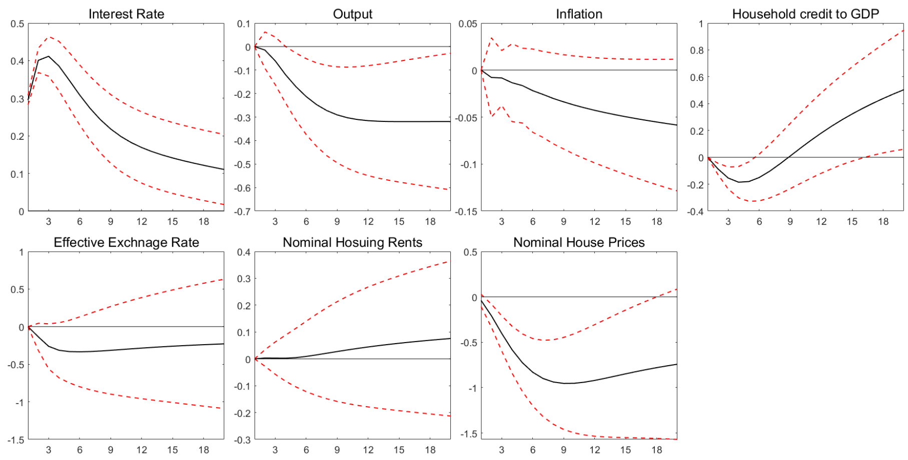
confidence level, the consistent responses of other macro variables would strengthen the credibility of my impulse response analysis.

2.2.4. Responses of Nominal House Prices and Housing Rents After Monetary Policy Shocks

Since housing rents are included in the core CPI index with a large weight, I calculate real housing rents using the core CPI weight of actual and imputed rentals. However, because my estimates of real housing rents are not based on micro CPI data, it may be subject to measurement errors. To supplement this, I estimate another panel VAR model with nominal housing rents without adjustment. Figure 2.4 shows the impulse response functions to one standard deviation monetary policy tightening shocks for the panel VAR model including the nominal house prices and nominal housing rents. 90 percent confidence intervals calculated by Monte Carlo estimated standard errors obtained from the White estimator with 1,000 repetitions and cumulative responses except for short−term rates are reported.

I find that the diversified responses of house prices and housing rents are also observed in those with nominal terms. In particular, nominal house prices decrease significantly by 0.95 percent after a monetary policy shock, whereas nominal housing rents increase insignificantly by 0.07 percent. Also, it is noted that the responses of other macro variables are consistent with stnadrad macroeconomic theories as in the previous subsection except for insignificant responses of real effective exchange rates. Using international panel data from 10 OECD economies, I find different responses of house prices and housing rents to monetary policy in both real and nominal terms, which would provide empirical evidences on rent stickiness as in micro−empirical evidences (Hoffmann and Kim, 2006; Shimizu et al., 2010; Gallin and Vergrugge, 2017).

Figure 2. 4. Impulse Responses to Tightening Monetary Policy Shocks with Nominal House Price and Housing Rents



Notes: Dashed lines show 90% confidence intervals calculated by Monte Carlo estimated standard errors obtained from the White estimator with 1,000 repetitions. The x−axis measures quarters.

2.3. The Model and Calibration

Homeownership channels, where households could decide between homeownership with mortgages and rental housing, can affect the monetary policy transmission mechanism in two ways. On the one hand, when more constrained households substitute mortgaged (rental) housing with rental (mortgaged) housing after monetary policy shocks, there will be faster household deleveraging (leveraging), amplifying the effects of monetary policy shocks. In particular, households can avoid (bear) the burden of interest costs by replacing mortgaged (rental) housing with rental (mortgaged) housing, which leads to the stronger deleveraging effects.

On the other hand, substitution of mortgaged (rental) housing with rental (mortgaged) housing may enhance (decrease) the liquidity of households' wealth in the short run as housing rents are less expensive then house prices on averages. This liquidity enhancement effects may lead to curtailed effects of tightening monetary policy shocks, buy only in the short−run.

Demand−side and supply−side dynamics in rental housing markets determine the aggregate effects of homeownership channels on the monetary policy transmission mechanism. On the demand side, the higher the substitution elasticity between rental and mortgaged housing, the weaker the short-term effects of monetary policy shocks, but the stronger the long-term effects. In terms of supply sides, rent stickiness can suppress substitution between rental housing and mortgaged housing. In particular, if only a small proportion of landlords could adjust their optimal housing rents as well as the supply of rental housing, there would be insufficient rental housing supplies to fulfill the increasing (decreasing) demand for rental housing following a contractionary (expansionary) monetary policy shock in the short run. Thus, rent stickiness would lead to less substitution between mortgaged and rental housing after monetary policy shocks, resulting in stronger short-term effects but weaker long-term effects of monetary policies.

In this context, I develop a New Keynesian dynamic stochastic general equilibrium (DSGE) model with homeownership channels to replicate the estimated responses of housing rents and housing prices in the previous section. In particular, I extend the standard framework of Iacoviello (2005) with two additional assumptions : (i) constant elasticity of substitution (CES) aggregate utility function between mortgaged housing and rental housing as in Ortega et al.(2011) and Rubio (2019), and (ii) segmented housing supplies with nominal rigidity in housing rents as in Dias and Duarte (2022).

2.3.1. Unconstrained Households

Unconstrained (patient) households, who would be savers around the steady state, maximize their lifetime utility function given by

$$E_0\sum_{t=0}^{\infty}\beta^t\left(lnc'_t + jlnh'_{1,t} - \frac{N'^{\eta}_t}{\eta}\right) \tag{2.5}$$

where E_0 is the expectation operator, $\beta \in (0,1)$ is the discount factor for unconstrained households, $\eta > 0$ and $j > 0$ is the relative weight of housing in the utility function. Unconstrained households choose c'_t, $h'_{1,t}$, and N'_t, which represent consumption, homeownership, and working hours at time t respectively. It is worth noting that I simplify the analysis by assuming that unconstrained households could only consume housing services in the form of homeownership as in Rubio (2019) and Dias and Duarte (2022).

The budget constraint for unconstrained households is given by

$$c'_t + q_{1,t}(h'_{1,t} - h'_{1,t-1}) + \frac{R_{t-1}}{\pi_t} b'_{t-1} = b'_t + w'_t N'_t + F_t \tag{2.6}$$

where $q_{1,t} = \frac{Q_{1,t}}{P_t}$ denotes the real housing price, R_{t-1} is the nominal interest rate on loans between t−1 and t, $w'_t = \frac{W'_t}{P_t}$ denotes the real wage for constrained households, and $\pi_t = \frac{P_t}{P_{t-1}}$ denotes the gross inflation rate. Note that unconstrained households receive lump−sum profit F_t from retailors, landlords, and real estate brokers.

Solving the problem gives the first−order conditions as follows

$$\frac{1}{c'_t} = \beta E_t\left(\frac{R_t}{\pi_{t+1} c'_{t+1}}\right) \tag{2.7}$$

$$\frac{j}{h'_{1,t}} = \left[\frac{q_{1,t}}{c'_t} - \beta E_t\left(\frac{q_{1,t+1}}{c'_{t+1}}\right)\right] \tag{2.8}$$

$$w'_t = N'^{\eta-1}_t c'_t \tag{2.9}$$

Equation (2.7) is the Euler equation for intertemporal consumption, equation (2.8) represents the intertemporal condition for house prices, and equation (2.9) shows the labor−supply condition for constrained households.

2.3.2. Constrained Households

Constrained (impatient) households, who would be borrowers around the steady state, maximize their lifetime utility function given by

$$E_0 \sum_{t=0}^{\infty} \beta''^{t} \left(lnc''_t + jlnh''_t - \frac{N''^{\eta}_t}{\eta} \right) \tag{2.10}$$

where $\beta'' \in (0, \beta)$ is the discount factor for constrained households, which is lower than that for unconstrained households. Constrained households choose c''_t, $h''_{,t}$, and N''_t, which represent consumption, housing services, and working hours at time t respectively. Note that unlike unconstrained households who could only consume housing services via homeownership[8], constrained households could choose both mortgaged housing and rental housing (Ortega et al., 2011; Rubio, 2019; Dias and Duarte, 2022). h''_t is a CES aggregation of homeownership ($h''_{1,t}$) and rental services ($h''_{2,t}$) given by

$$h''_t = \left[\omega_h^{1/\varepsilon_h} h''^{(\varepsilon_h - 1)/\varepsilon_h}_{1,t} + (1 - \omega_h)^{1/\varepsilon_h} h''^{(\varepsilon_h - 1)/\varepsilon_h}_{2,t} \right]^{\varepsilon_h/(\varepsilon_h - 1)} \tag{2.11}$$

where ω_h denotes the preference of constrained households for homeownership and ε_h is the elasticity of substitution between homeownership and rental housing. As Rubio (2019) pointed out, the assumption does not imply that each borrower lives in both a mortgaged house and a rented house at the same time; rather, it reflects the representative preferences of constrained households consisting of a continuum of members.

Constrained households are subject to the budget constraint and the borrowing constraint given by

$$c''_t + q_{1,t}(h''_{1,t} - h''_{1,t-1}) + q_{2,t}h''_{2,t} + \frac{R_{t-1}}{\pi_t} b''_{t-1} = b''_t + w''_t N''_t \tag{2.12}$$

$$b''_t \leq E_t \left(\frac{1}{R_t} m'' q_{1,t+1} h''_{1,t} \pi_{t+1} \right) \tag{2.13}$$

where b''_t denotes borrowing of constrained households at time t and m'' represents a loan to value ratio in terms of the borrowing constraints proposed by Kiyotaki and Moore (1997). As Iacoviello

[8] While unconstrained households own houses outright, constrained households own houses with mortgages. So we call homeownership of constrained households as mortgaged housing.

(2005) pointed out, the assumption $\beta'' < \beta$ guarantees that the borrowing constraint holds with equality. It is also important to note that constrained households cannot benefit from selling their rented houses in the past, and rented houses are not included in the borrowing constraint because they cannot be used as collateral.

The first-order conditions for constrained households are summarized as follows

$$\frac{1}{c''_t} = \beta'' E_t\left(\frac{R_t}{\pi_{t+1}c'_{t+1}}\right) + \lambda''_t R_t \tag{2.14}$$

$$\mathrm{j}\frac{\omega_h^{1/\varepsilon_h} h''^{-1/\varepsilon_h}_{1,t}}{h''^{(\varepsilon_h-1)/\varepsilon_h}_t} = \left[\frac{q_{1,t}}{c'_t} - \beta E_t\left(\frac{q_{1,t+1}}{c'_{t+1}}\right) + \lambda''_t m'' q_{1,t+1}\pi_{t+1}\right] \tag{2.15}$$

$$\mathrm{j}\frac{(1-\omega_h)^{1/\varepsilon_h} h''^{-1/\varepsilon_h}_{2,t}}{h''^{(\varepsilon_h-1)/\varepsilon_h}_t} = \frac{q_{2,t}}{c'_t} \tag{2.16}$$

$$w''_t = N''^{\eta-1}_t c''_t \tag{2.17}$$

where λ''_t is the Largrange multiplier on the borrowing constraint, which increase with the life time utility from borrowing R_t dollars. We can compare the difference between the constrained households' demands of mortgaged housing and rental housing by using the equation (2.15) and (2.16). The demand for homeownership is affected by relative price differential with hosing rents, future house price expectations ($\beta E_t\left(\frac{q_{1,t+1}}{c'_{t+1}}\right)$), and the value as the collateral ($\lambda''_t m'' q_{1,t+1}\pi_{t+1}$). On the other hand, the demand for rental housing is determined mostly by relative price differential with mortgaged housing.

2.3.3. Entrepreneurs and Retailors

In my model, entrepreneurs are assumed following Rubio (2019) and Dias and Duarte (2022), who assumed that entrepreneurs did not use real estate as inputs as well as collaterals. Although Kiyotaki and Moore (1997) and Iacoviello (2005) incorporated housing in a production function as well as a borrowing constraint of entrepreneurs, I do not consider this in order to concentrate on homeownership decision channels among households.

Entrepreneurs use a Cobb−Douglas constant returns−to−scale technology using capital and labor as inputs in order to produce an intermediate good Y_t as follows

$$Y_t = A_t K_t^{\mu} N'^{\alpha(1-\mu)}_t N''^{(1-\alpha)(1-\mu)}_t \tag{2.18}$$

where A_t is the technology parameter and N'_t and N''_t are labor input of unconstrained households and constrained households, respectively. μ and $\alpha \in (0,1)$ would be income shares of capitals and unconstrained households labor in the steady state.

Entrepreneurs optimize their lifetime utility function given by

$$E_0 \sum_{t=0}^{\infty} \gamma^t ln\, c_t \tag{2.19}$$

where $\gamma \in (0,\beta)$ is the discount factor for entrepreneurs lower than that for unconstrained households in order to guarantee that the borrowing constraint holds with equality. Entrepreneurs solve their maximize problems subject to the technology constraint in equation (2.18), the flows of funds, and the borrowing constraint given by

$$Y_t/X_t + b_t = c_t + \frac{R_{t-1}}{\pi_t} b'_{t-1} + w'_t N'_t + w''_t N''_t + I_t + \xi_{K,t} \tag{2.20}$$

$$b_t \leq 0 \tag{2.21}$$

where I_t denotes capital investment defined as $I_t = K_t - (1-\delta)K_{t-1}$ with the capital adjustment costs $\xi_{K,t} = \frac{\psi}{2\delta}(I_t/K_{t-1} - \delta)^2 K_t$, and b_t denotes borrowing of entrepreneurs at time t. It is noted that entrepreneurs are neither savers nor borrowers in this model as we assume zero loan to value ratio for mortgage lending of entrepreneurs as in equation (2.21).

Solving the entrepreneurs’ problem provides the first order conditions given by

$$\frac{1}{c_t} = \gamma E_t\left(\frac{R_t}{\pi_{t+1} c_{t+1}}\right) + \lambda_t R_t \tag{2.22}$$

$$\frac{1}{c_t}\left(1+\frac{\psi}{\delta}\left(\frac{I_t}{K_{t-1}}-\delta\right)\right)$$
$$= E_t\left[\frac{\gamma}{c_{t+1}}\left(\frac{\mu Y_{t+1}}{X_{t+1}K_t}+1-\delta+\frac{\psi}{\delta}\left(\frac{I_{t+1}}{K_t}-\delta\right)\left(\frac{1}{2}\left(\frac{I_{t+1}}{K_t}-\delta\right)+1-\delta\right)\right)\right] \tag{2.23}$$
$$w'_t = \frac{\alpha(1-\mu)Y_{t+1}}{X_{t+1}N'_t} \tag{2.24}$$
$$w''_t = \frac{(1-\alpha)(1-\mu)Y_{t+1}}{X_{t+1}N''_t} \tag{2.25}$$

where λ_t denotes the shadow value of the borrowing constraint for entrepreneurs at the time t.

As in Iacoviello (2005), I follow the retailer's problem from Bernanke, Gerlter and Gilchrist (1999) in order to motivate sticky prices. Without any modification, the aggregate price level evolution is given by

$$P_t = (\theta P_t^{1-\epsilon} + (1-\theta)P_t^{*1-\epsilon})^{1/(1-\epsilon)} \tag{2.26}$$

where θ denotes the probability of fixed prices in Calvo (1983) and $P_t^* = \left(\int_0^1 P_t^*(z)^{1-\epsilon} dZ\right)^{1/(1-\epsilon)}$ with reset price $P_t^*(z)$ for retailor $z \in [0,1]$ given by $\max_{P_t^*(z)} \sum_{k=0}^{\infty} E_t\left\{\beta^k \frac{c'_t}{c'_{t+k}}\left(\frac{P_t^*(z)}{P_{t+k}} - \frac{X}{X_{t+k}}\right)\frac{P_t^*(z)}{P_{t+k}}^{-\epsilon} Y_{t+k}\right\} = 0$. Note that these assumptions yield a forward−looking Phillips curve in the linearized system.

2.3.4. Housing Supply

Housing supply in my model is assumed following Dias and Duarte (2022). The fixity of housing supply in the aggregate is assumed as in Iacoviello (2005). The total stock of real estate, $\bar{H}$, is segmented into a part for homeownership (outright ownership and mortgage), $H_{1,t}$, and that for rental housing, $H_{2,t}$, summarized as $\bar{H} = H_{1,t} + H_{2,t}$ for every time t. Although the total supply of housing is fixed in the aggregate, the landlords can adjust the current composition of homeownership and rental housing by renting (selling) houses available for owning (renting) via real estate brokers. Adjustment costs for landlords of Dias and Duarte (2022) are not assumed in this paper as we focus on reproducing less volatile responses of housing rents to monetary policy shocks.

There is a competitive unit mass of landlords who maximize their lifetime profits given by

$$\max_{I_{2,t}H_{2,t}} E_0 \sum_{t=0}^{\infty} \Lambda_t \left(\frac{q_{2,t}}{X_{2,t}} H_{2,t} - q_{1,t} I_{2,t} \right) \tag{2.27}$$

where $\Lambda_t = \prod_{\tau=0}^{t} \frac{\pi_\tau}{R_{\tau-1}}$, the saver' s relevant discount factor, $X_{2,t}$ is the real estate brokers' markup which would be described below, $H_{2,t}$ is the landlords' supply of rental housing at time t, and $I_{2,t}$ denotes the landlords' investment of rental housing stock as in the equation (2.28).

$$H_{2,t} = I_{2,t} + H_{2,t-1} \tag{2.28}$$

Equation (2.27) and (2.28) shows that landlords adjust the supply of rental housing by buying new housing stocks $I_{2,t}$ at $q_{1,t}$, converting them into those for rental housing, $H_{2,t}$, and renting them via real estate brokers at $\frac{q_{2,t}}{X_{2,t}}$.

The first−order conditions for the landlords are given by

$$\frac{q_{2,t}}{X_{2,t}} = q_{1,t} - E_t\left(\frac{\pi_{t+1}}{R_t} q_{1,t+1}\right) \tag{2.29}$$

Equation (2.29) shows that housing rents ($q_{2,t}$) at time t are determined by current house prices ($q_{1,t}$), expected value of house prices at time t+1 ($E_t\left(\frac{\pi_{t+1}}{R_t} q_{1,t+1}\right)$), and real estate brokers' mark up at time t ($X_{2,t}$). It is noteworthy that housing prices in the steady state are determined by the current value of a future sum of housing rents ($\frac{q_2}{X_2(1-\beta)} = q_1$), where real estate broker' s mark up (X_2) can affect the house price to rent ratio. As Dias and Duarte (2022) pointed out, However, rent stickiness assumption would generate some deviations from this systematic relationship between housing rents and house prices.

Rent stickiness is assumed as in Dias and Duarte (2022), where only a small fraction of rental contracts could change their housing rents every period. In order to motivate rent stickiness, I assume that there are monopolistic competitive real estate brokers who serve as intermediaries of rental housing transaction between landlords and households. In particular, real estate brokers buy the housing stock for renting from landlords and rent them at a markup to households. However, a fraction $1-\theta_2$ of real estate brokers can set optimal housing rents each period, whereas a fraction θ_2 cannot. The real estate brokers' problem is solved following Bernanke, Gerlter and Gilchrist (1999), delivering housing-rental Phillips curve given by

$$\hat{\pi}_t + \hat{q}_{2,t} - \hat{q}_{2,t-1} = \beta E_t\left(\hat{\pi}_{t+1} + \hat{q}_{2,t+1} - \hat{q}_{2,t}\right) - \frac{(1-\theta_2)(1-\beta\theta_2)}{\theta_2}\hat{X}_{2,t} \quad (2.30)$$

2.3.5. Interest Rate Rule

In order to replicate the estimated responses of short-term rates following AR (2) process, the monetary policy rule specification follows Coibion and Gorodnichenko (2011) which allowed for interest smoothing of order two[⑨].

$$R_t = R_{t-1}^{r_R} R_{t-2}^{r_{2,R}} \left(\pi_{t-1}^{1+r_\pi} (Y_{t-1}/Y)^{r_Y} \overline{rr}\right)^{1-r_R-r_{2,R}} e_{R,t} \quad (2.31)$$

where $\overline{rr}$ and Y denotes steady-state real interest rate and output respectively. Monetary policy is assumed to respond systematically to past inflation and past output, which is consistent with the recursive ordering assumption in section II. $e_{R,t}$ denotes a white noise shock process with zero mean and variance σ_R^2.

2.3.6. Equilibrium

The equilibrium consists of $\{h'_{1,t}, h''_{1,t}, h''_{2,t}, H_{2,t}, X_t, X_{2,t}\, N'_t, N''_t, Y_t, I_t,$

---

⑨ As including output growth term generates implausible impulse responses in our models, we do not include output growth term in the rule. Output gap term is also excluded as we do not have productivity shocks in our model.

$K_t, c_t, c'_t, c''_t, b_t, b'_t, b''_t\}_{t=0}^{\infty}$ with the sequence of values $\{w'_t, w''_t, R_t, P_t, P_t^*,$ $X_t, \lambda_t, \lambda''_t, q_{1,t}, q_{2,t}\}_{t=0}^{\infty}$ satisfying equations (2.5) to (2.31) given the previous conditions $\{h'_{1,t}, h''_{1,t-2}, h''_{2,t}, H_{2,t-1}, K_{t-1}, R_{t-1}, b'_{t-1}, b''_{t-1}, P_{t-1}\}$, the monetary shocks $\{e_{R,t}\}$, and the following market clearing conditions

Goods market : $c_t + c'_t + c''_t + I_t = Y_t$ (2.32)

Rental housing market : $h''_{2,t} = H_{2,t}$ (2.33)

Homeownership housing market : $h'_{1,t} + h''_{1,t} = \bar{H} - H_{2,t}$ (2.34)

Loan market : $b'_t + b''_t = 0$ (2.35)

2.3.7. Calibration

In this subsection, I calibrate the model using long-run averages or parameter values from the literature. Table 2.3 presents the calibrated parameter values of the model based on long-run averages or from the literature. For preference parameters, I set discount factors for unconstrained households (β) to be 0.99, which is a standard value allowing average annual rate of return to be around 4 percent. I set discount factors for entrepreneurs (γ) to be 0.98 so that firm' s annual rate of return is twice as big as equilibrium rate as in Iacoviello (2005). I set discount factors for constrained households (β'') to be 0.95, which is consistent with Lawrance (1991), Caroll and Samwick (1997), and Samwick (1998). I set the housing preference weight across households (j) to 0.1 as in Iacoviello (2005), resulting in the real estate value to GDP ratio of 1.36 in steady states, which is in line with the standard value in macroeconomy with housing sector literature. I set $\eta = 1.276$ so that the inverse Frisch elasticity is 0.276, following Gertler and Paradi (2011). Preference share parameters for mortgaged housing (ω_h) is calibrated to be 0.545 in the CES aggregate utility function between mortgaged housing and rental housing for constrained households, so that the stock for mortgaged housing to rental housing ratio ($\frac{h''_1}{h''_2}$) in the steady state is 1.2, which is the average ratio of households living in mortgaged housing and rental housing in the United States from 2010 to 2020.

Table 2. 3. Calibrated Parameter Values

Parameter	Value	Description	Sources
Parameters for preferences			
β	0.99	Saver discount factors	Standard
γ	0.98	Entrepreneur discount factors	Iacoviello (2005)
β''	0.95	Borrower discount factors	Iacoviello (2005)
j	0.1	Preference weight on housing across households	Housing value to GDP ratio of 1.36, Iacoviello (2005)
$\eta - 1$	0.28	Inverse of Frisch elasticity	Gertler and Paradi (2011)
ω_h	0.5455	Preference share parameter of homeownership for borrowers	U.S. average mortgaged housing to rental housing ratio, $\frac{h''_1}{h''_2} = 1.2$
ε_h	2.5	Elasticity of substitution between owning and renting	Minimize the distance between the empirical impulse responses
Parameters for final goods sectors			
μ	0.3	Share of variable capital in the output	Iacoviello (2005)
δ	0.03	Variable capital depreciation rate	Iacoviello (2005)
X	1.05	Steady−state gross markup	Iacoviello (2005)
ψ	2	Variable capital adjustment cost	Iacoviello (2005)
α	0.64	Savers labor income share	Iacoviello (2005)
θ	0.84	Probability of fixed prices	Iacoviello and Neri (2010)
Parameters for housing sectors			
X_2	2.2	Real estate brokers markup	U.S. average house price−to−rent ratio in Dias and Duarte (2022) : $\frac{q_1}{4*q_2} = 11.4$
θ_2	0.83	Rent stickiness	Gallin and Verbrugge (2019)
m''	0.7	LTV limits for constrained households	U.S. long−run average loan to−value ratio of mortgage holders in Gelain et al. (2012)
Parameters for monetary policies			
r_R	(1.45, 0.73)	Taylor rule smoothing of order 1	Panel OLS regression results for short−term rates, Iacoviello (2005)
$r_{R,2}$	(−0.49, 0)	Taylor rule smoothing of order 2	Panel OLS regression results for short−term rates, Iacoviello (2005)
r_y	(0.44, 0.13)	Taylor rule (lagged output)	Coibion and Gorodnichenko (2011), Iacoviello (2005)
r_π	(0.58, 0.27)	Taylor rule (lagged inflation)	Coibion and Gorodnichenko (2011), Iacoviello (2005)
σ_R^2	0.3	Standard deviation of monetary policy shocks	Iacoviello (2005)

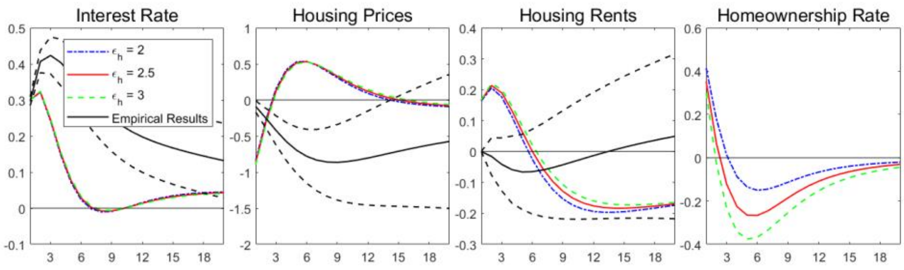
In terms of final good sectors, I choose standard values in macroeconomic literature so that $\mu = 0.3$, $\delta = 0.03$, and $X = 1.05$, following Iacoviello (2005). The inverse elasticity of investment to the capital shadow price (ψ) is set to 2 based on King and Wolman (1996) and Iacoviello (2005). The labor income share of unconstrained households (α) is calibrated to 0.64, which is the estimate of Iacoviello (2005). I set the Calvo parameter for final goods (θ) to 0.84 as in Iacoviello and Neri (2010).

For the parameters related to housing sectors, I set real estate brokers' mark up (X_2) to 2.2 as in Dias and Duarte (2022), which guarantees house price to rent ratio in the steady state ($\frac{q_1}{q_2} = \frac{1}{(1-\beta)X_2}$) to 45.6 percent, which coincides with the U.S. average house price to rent ratio. The Calvo parameter for rental housing (θ_2), which determines rental stickiness, is calibrated to 0.83 based on empirical findings using micro data of Gallin and Verbrugge (2019). I set the loan-to-value ratio for constrained households $m'' = 0.7$ to match the long-run average loan to-value ratio of the U.S. residential mortgage holders, following Gelain et al (2013).

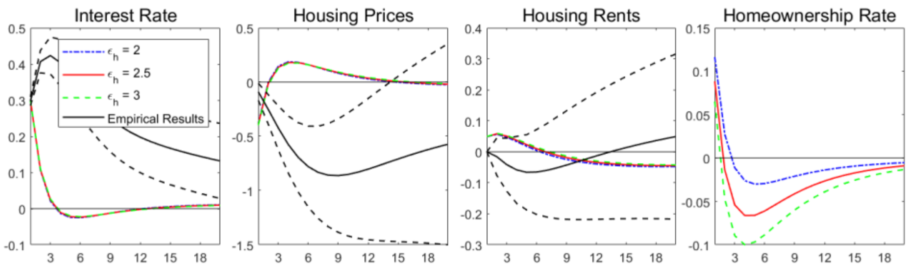
For the monetary policy rule, I use post-Volcker estimates in Coibion and Gorodnichenko (2011) regarding Taylor rule coefficients for output and inflation. Based on my panel OLS regression results on 10 OECD economies, I set Taylor rule smoothing parameters $r_R = 1.45$ and $r_{R,2} = -0.49$ instead of 1.12 and -0.18 as in Coibion and Gorodnichenko (2011) to replicate the responses of short-term rate after monetary policy shocks more precisely. As the assumption of interest smoothing of order two in Coibion and Gorodnichenko (2011) might affect the dynamics of the model, I also consider a standard Taylor rule with interest rate smoothing of order one following Iacoviello (2005), which calibrates $r_R = 0.7$, $r_{R,2} = 0$, $r_y = 0.13$, and $r_\pi = 0.27$.

Finally, I calibrate the elasticity of substitution between rental housing and mortgaged housing (ε_h) based on the empirical results in housing sector in section II. Although Ortega et al. (2011), Rubio (2019), and Rubaszek and Rubio (2020) set $\varepsilon_h = 2$, their calibration

Figure 2. 5. Impulse Responses of Housing Sector to Monetary Policy Shocks with Different Calibration for the Elasticity of Substitution between owing and renting



Panel (a): Taylor rule of following Coibion and Gorodnichenko (2011)



Panel (b): Taylor rule of following Iacoviello (2015)

Notes: Blue dotted lines denote impulse responses from the model with $\varepsilon_h = 2$, red lines denote those from the model without $\varepsilon_h = 2.5$, green dashed lines denote those from the model without $\varepsilon_h = 3$, and black dashed lines show 90% confidence intervals calculated by Monte Carlo estimated standard errors obtained from the White estimator with 1,000 repetitions. The x−axis measures quarters.

was not based on any empirical evidences due to lack of related studies. Figure 2.5 compares the impulse response functions from the calibrated model based on different values of the elasticity of substitution to the empirical results from panel VAR models in Figure 2.3. Although I do not estimate the responses of homeownership rates [10] to monetary policy shocks as homeownership rates for OECD economies are only available for annual data, I refer to Dias and Duarte (2022) that homeownership

[10] I measure homeownership rates as a percentage of housing units occupied by owners, rather than a percentage of households who own houses as in Dias and Duarte (2022). Indeed, the US Census Bureau calculates the homeownership rate by dividing the number of owner-occupied housing units by the number of occupied housing units.

rates declined significantly by 0.8 percent ten quarters after one standard deviation monetary policy shocks for the U.S. data. I set $\varepsilon_h = 2.5$[11] as it better replicates the empirical responses of housing rents as well as homeownership rates across different specification of Taylor rules than other values.

2.4. Simulation Results

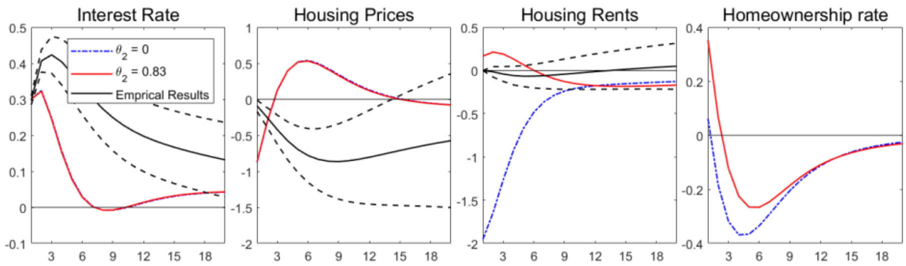
In this section, I evaluate the impulse response functions from the calibrated model by comparing responses of housing sector to monetary policy shocks with the empirical impulse response functions from the panel VAR model in section II. In particular, I focus on how rent stickiness affects diversified responses of housing rents and house prices as in the empirical results. Furthermore, I investigate the effects of homeownership channels on the monetary policy mechanism by comparing the impulse response functions between the model with and without homeownership channels. I also analyze a role of rent stickiness in homeownership channels by comparing the impulse response functions with and without rent stickiness at the aggregate level.

2.4.1. The Role of Rent Stickiness in Different Responses of Housing Rents and House Prices

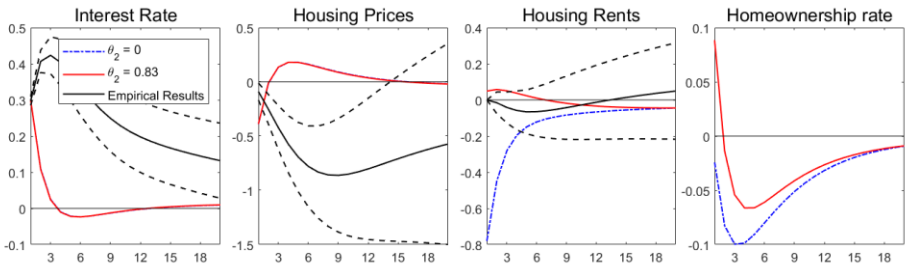
I model rent stickiness by assuming that $1-\theta_2$ of real estate brokers could adjust their housing rents after monetary policy shocks. Thus, I could analyze the role of rent stickiness in replicating the empirical responses of housing rents to monetary policy shocks by comparing the impulse response functions from the models with two different parameterizations for $\theta_2 =$ 0.83 and 0. Figure 2.6 presents the impulse response of short-term rates, real housing rents, real house prices, and homeownership rates after one standard deviation monetary policy shocks with and without

[11] The conclusion of this paper is robust to other calibration choices of the elasticity of substitution unless it has positive values.

Figure 2. 6. Impulse Responses of Housing sector
after a Monetary Policy shock with and without rent stickiness



Panel (a): Taylor rule of following Coibion and Gorodnichenko (2011)



Panel (b): Taylor rule of following Iacoviello (2015)

Notes: Red lines denote impulse responses from the baseline model with rent stickiness ($\theta_2 = 0.83$), blue dotted lines denote those from the model without rent stickiness ($\theta_2 = 0$), and black dashed lines show 90% confidence intervals calculated by Monte Carlo estimated standard errors obtained from the White estimator with 1,000 repetitions. The x−axis measures quarters.

rent stickiness. The empirical responses of short−term rates, real housing rents, and real house prices from the panel VAR models in section 2.2.3 are also displayed in black lines in order to evaluate the calibrated model.

It can be easily noticed that the baseline models with rent stickiness are better to replicate the empirical responses of housing rents to monetary policy shocks than those without rent stickiness across different Taylor rule specifications. In particular, housing rents fall stronger than house prices to monetary policy shocks without rent stickiness as in Rubio (2019). On the other hand, the responses of house prices from the calibrated models are almost identical regardless of rent stickiness features[12]. It is also

[12] Although the empirical housing price responses to monetary policy

interesting that homeownership rates increase shortly right after monetary policy shocks in the model with rent stickiness, whereas they fall immediately and more severely in the model without rent stickiness.

I can conjecture the mechanism how rent stickiness would induce the diversified reactions of housing rents and prices to monetary policy shocks. When only a small fraction of real estate brokers can adjust housing rents and supplies of rented houses, housing rents fall less then house prices after contractionary monetary policy shocks. This deviation leads to increase of homeownership rates in the short-run, leading to less substitution of mortgaged housing and rental housing in the presence of rent stickiness. I analyze how these changes in homeownership decision affect the monetary policy transmission mechanism in the next subsection.

2.4.2. The Role of Homeownership channels in the Monetary Policy Transmission Mechanism

In order to examine how homeownership channels affects the effects of monetary policies, I calculate the impulse response functions from the model without homeownership channels, where constrained households could not substitute rental (mortgaged) housing with mortgaged (rental) housing, but they consume both types of housing at a fixed ratio (i.e. $\varepsilon_h = 0$). Figure 2.7 shows the impulse response functions to monetary policy shocks from the baseline models with and without homeownership channels under the monetary policy rules following Coibion and Gorodnichenko (2011).

It is worth noting that the responses of the housing sectors to monetary policy shocks differ significantly between models with and without homeownership channels. In the model without homeownership channels, mortgaged housing and rental housing of

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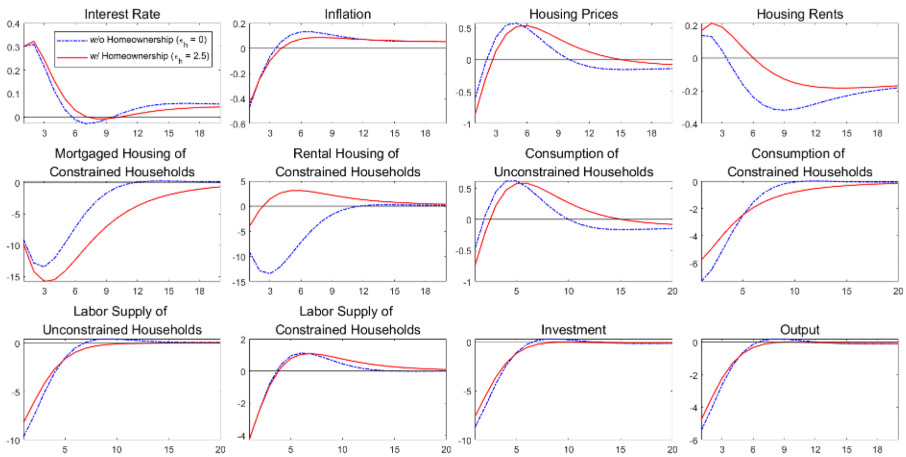
shocks are stickier than those from calibrated models, the magnitude of minimum responses is similar.

Figure 2. 7. Impulse Responses Functions after a Monetary Policy shock with and without homeownership channels



Notes: Red lines denote impulse responses from the baseline model with homeownership channels ($\varepsilon_h = 2.5$), whereas blue dotted lines denote those from the model without homeownership channel ($\varepsilon_h = 0$)

constrained households fall equally after monetary policy shocks. In the model with homeownership channels, However, mortgaged housing of constrained households falls stronger, whereas rented housing of constrained households falls weaker. These disparities in mortgaged and rental housing responses to monetary policy shocks imply that borrowers substitute mortgaged housing with rental housing after interest rate hikes.

Note that homeownership rates are calculated as a percentage of houses occupied by homeowners, not as a percentage of households who own houses as in Dias and Duarte (2022). This is how the United States Census Bureau calculates homeownership rates. As a result, even in the model without homeownership channels, where the ratio of renters to owners in borrowers keeps constant, homeownership rates might change after monetary policy shocks due to changes in income distribution between borrowers and savers. In particular, homeownership rates rise in the model without homeownership channels because savers increase the number of owner-occupied housing units as their income rises as a result of interest rate hikes. In the model with homeownership channels, on the other hand, homeownership rates rise briefly due to rent rigidity and subsequently fall as constrained households replace mortgaged housing with rental housing.

It is also interesting that there is a pronounced difference in aggregate economy responses to monetary policy shocks between the economy with and without homeownership channels. In order to investigate the effects of homeownership channels in better detail, I calculate the difference in impulse response functions between the models with and without homeownership channels as follows

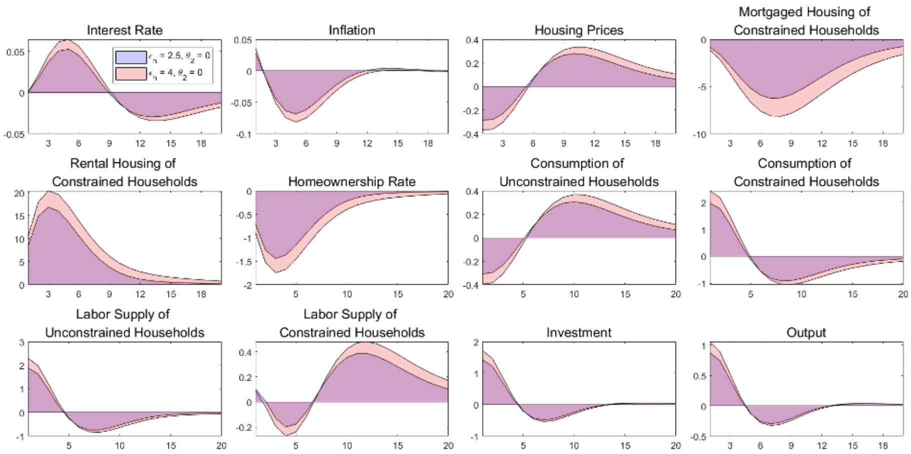
$$diff_{k,t} = irf_{k,t}^{w/\ homeowership} - irf_{k,t}^{w/o\ homeowership} \qquad (2.36)$$

Note for contractionary monetary policy shocks, positive (negative) values of the differences imply that the effects of monetary policies would be weaker (stronger) in the presence of homeownership channels.

First, I examine the impacts of homeownership channels on monetary policies without the rent rigidity assumption in order to figure out how substitution between mortgaged and rental housing after monetary policy shocks affects the monetary policy transmission mechanism. In particular, I compare the differences from the model with distinct values of the elasticity of substitution between owning and renting (ε_h). Figure 2.8 depicts the difference in impulse response functions due to homeownership channels across the models with different values of the elasticity of substitution ($\varepsilon_h = \mathbf{2.5}$ and 4) under the flexible housing rent assumption ($\theta_2 = \mathbf{0}$).

I focus on the changes in the impulse responses with and without homeownership channels when the elasticity of substitution between renting and owning is 2.5, which is the benchmark for my calibration. The most pronounced difference in real sectors is found in the impulse responses for consumption of constrained household. Concretely, the difference is around 2 percent for the short−run, then becomes negative at 0.89 percent after 5 quarters, implying the weaker short−term effects and stronger long−term effects of

Figure 2. 8. Difference in impulse response functions between the models with and without homeownership channels in the absence of rent stickiness



Notes: Blue areas denote the differences in impulse responses with and without homeownership channels where $\varepsilon_h = 2.5$ and $\theta_2 = 0$, and red areas denotes those where $\varepsilon_h = 4$ and $\theta_2 = 0$. The impulse responses without homeownership channels are calculated without rent rigidity (i.e. $\theta_2 = 0$) for ease of comparison.

monetary policies with homeownership channels.

How does the homeownership channel lead to curtailed responses of borrowers' consumption to monetary policy shocks? This is because constrained households can not only reduce interest costs, but also enhance liquidity by substituting homeownership with less expensive rental housing. However, this liquidity enhancement effects persist transiently. Meanwhile, replacing mortgaged housing with rental housing forces into more rapid deleveraging, resulting in stronger long-term effects of monetary policy shocks.

The contrasting effects of homeownership channels on the short-term and long-term impulse responses to monetary policy shocks are more pronounced when the elasticity of substitution between renting and owning increases to 4. With a higher elasticity of substitution, more constrained households replace mortgaged housing with rental housing after interest rate hikes. In particular, the difference in mortgaged housing of constrained households with and without homeownership channels is − 8.17 percent with $\varepsilon_h = 4$, stronger than − 6.25 percent with $\varepsilon_h = 2.5$. At the aggregate level, this increased substitution between rental housing and mortgaged housing causes weaker declines of output in the short−run, but more persistent decrease in the long−run.

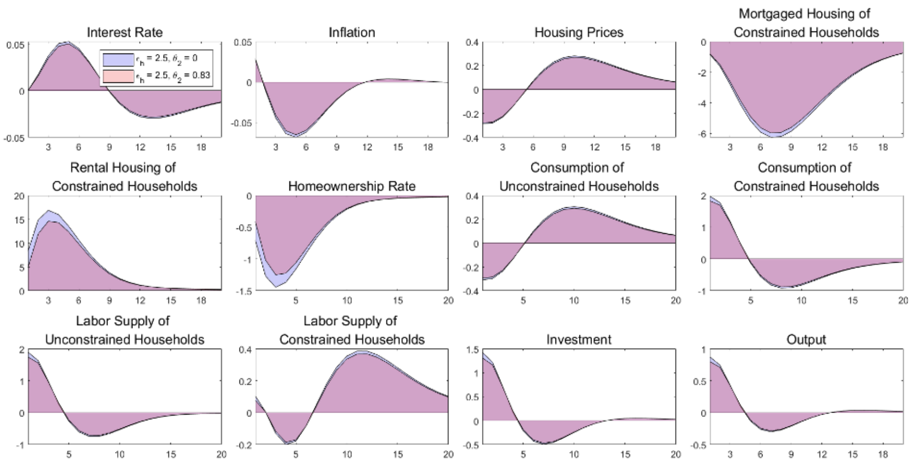
In short, I find that homeownership channels curtail the short−term effects of monetary policy shocks by enhancing household liquidity temporarily. However, as more households substitute mortgaged housing with rental housing after contractionary monetary policy shocks, there is more rapid deleveraging via homeownership channels, which amplifies the long−term effects of monetary policy shocks.

2.4.3. The Role of Rent Stickiness in Homeownership Channels

Finally, I analyze how the rent rigidity assumption affects homeownership channels in the monetary policy transmission mechanism. In the previous section, rent stickiness plays a role in replicating the estimated responses of housing rents and house prices. Figure 2.9 shows the difference in impulse response functions due to homeownership channels between the models with and without rent stickiness ($\theta_2 = 0.83$ and 0). The major difference between the two scenarios is that in the lack of rent stickiness, housing rents and homeownership rates fall sharply and immediately after monetary policy shocks, whereas in the presence of rent stickiness, housing rents remain stable and supplies of rented houses increase more slowly.

I find that rent stickiness prevents borrowers from switching mortgaged home for rental housing after a contractionary monetary policy shock. The difference in the responses of borrowers' rental housing units with and without homeownership channels three quarters after the shocks, in particular, is 16.87 percent in the absence of rent stickiness, but 14.59 percent in the presence of rent stickiness. This means that rent stickiness leads to less substitution of mortgaged housing with rental housing after contractionary monetary policy shocks. Therefore, the reverse

Figure 2. 9. Difference in impulse response functions between the models with and without homeownership channels across different values of rent stickiness



Notes: Blue areas denote the differences in impulse responses with and without homeownership channels with $\varepsilon_h = 2.5$ and $\theta_2 = 0.83$, and red areas denotes those with $\varepsilon_h = 2.5$ and $\theta_2 = 0$. The impulse responses without homeownership channels are calculated without rent rigidity (i.e. $\theta_2 = 0$) for ease of comparison.

occurs when I analyze the effects of the increased substitution between mortgaged housing and rental housing via homeownership channels in section 2.4.2.

After monetary policy shocks, the curtailed substitution of mortgaged housing for rental housing, as well as sticky housing rents, leads to diminished constrained household liquidity enhancement in the short run. Therefore, the difference in the responses of output with and without homeownership channels decreases from 0.87 percent to 0.79 percent at the impact of shocks in the presence of the rent rigidity assumption. This suggests that the rent stickiness would weaken the effects of homeownership channels on monetary policies. Those of borrowers' consumption also decreases from 1.97 percent to 1.82 percent. However, only for the first 2-3 quarters following the shocks does rent stickiness influence the effects of homeownership channels in the monetary policy mechanism. Furthermore, compared to homeownership channels, the quantitative effects of rent stickiness at the aggregate levels are quite modest.

In conclusion, I find that the housing rent rigidity assumption prevents borrowers from substituting mortgaged housing with rental housing in the short-run, which leads to stronger short-term effects of monetary policies. It is also important to note that although rent stickiness assumption plays a key role in replicating the empirical responses of housing rents, however, the quantitative effects of rent stickiness in monetary policies at the aggregate levels are modest and transient.

2.5. Concluding Remarks

This paper fills the gap between empirical evidences and theoretical models regarding the role of homeownership decision channels and rent stickiness in the monetary policy mechanism. I show the diversified responses of housing rents and house prices to monetary policy shocks with panel VARX models on 10 OECD economies with different housing tenure distribution. I propose a New Keynesian model with homeownership decision channels assuming rent stickiness in order to replicate these empirical results. Substitution of mortgaged (rental) housing with rental (mortgaged) housing curtails the short-term effects of monetary policy shocks by liquidity enhancement, but leads to more persistent effects in the long-run by forcing into more rapid deleveraging. Although the rent rigidity assumption plays a key role in replicating the empirical facts, it has a limited effect in the monetary policy mechanism, which would justify a theoretical model that does not assume rent rigidity.

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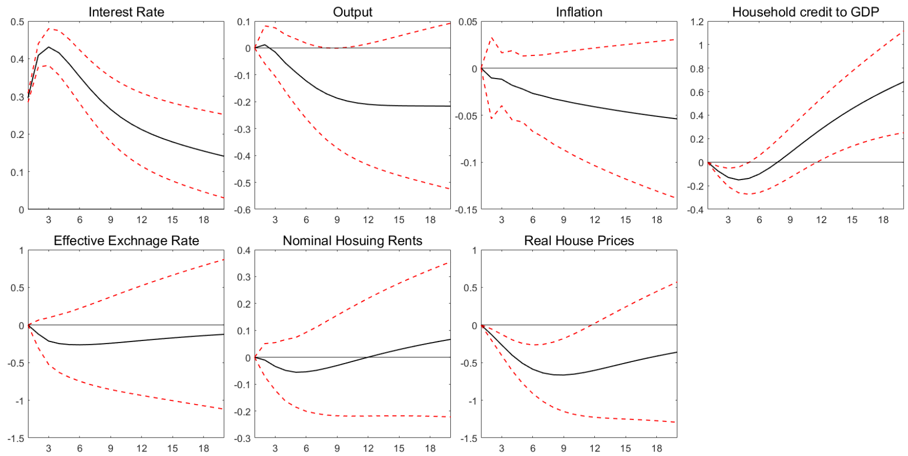
Appendix 2.A. Data and Sources

Variable Name	Sources	Notes
10−Year Treasury Constant Maturity Minus 2−Year Treasury Constant Maturity	Federal Reserve Economic Data	Percent, Quarterly, Not Seasonally Adjusted
Gross domestic product − expenditure approach	OECD Statistics	National currency, volume estimates, OECD reference year, annual levels, seasonally adjusted
CPI: All items non−food non−energy	OECD Statistics	Index (2015=100)
Real house price index	OECD Statistics	Index (2015=100), the ratio of the nominal house price index to the consumers' expenditure deflator in each country, seasonally adjusted
Nominal house price index	OECD Statistics	Index (2015=100), seasonally adjusted
Rent prices index	OECD Statistics	Index (2015=100), seasonally adjusted
CPI weights: All items non−food non−energy	OECD Statistics	Per thousand of the National CPI Total
CPI weights: Actual and Imputed rentals for housing	OECD Statistics	Per thousand of the National CPI Total
BIS effective exchange rate	BIS Statistics Warehouse	Real (CPI−based), Broad Indices
Short−term interest rates	OECD Statistics	Percent per annum
Credit to Households and NPISHs from All sectors at Market value	BIS Statistics Warehouse	Percentage of GDP, adjusted for breaks

Appendix 2.B. Responses of real housing rents and housing prices under the alternative specification



Notes: Estimation based on the Cholesky ordering $\{\triangle lny_t, \triangle \pi_t, \triangle hdebt_t, \triangle lne_t, \triangle lnrent, \triangle lnhp_t, i_t\}$. Dashed lines show 90% confidence intervals calculated by Monte Carlo estimated standard errors obtained from the White estimator with 1,000 repetitions. The x-axis measures quarters.

Chapter 3. International Capital Flow Shocks and Economic Crisis in East Asian Countries

3.1. Introduction

At the end of the 1990s, during the Asian Financial Crisis (AFC), several east Asian economies experienced significant output contraction, currency depreciation, and rapid capital outflows as a result of international capital flow shocks. After a decade, the 2008 Global Financial Crisis (GFC), precipitated by the United States, was expected to have similar effects on those east Asian countries. However, capital outflows and their impacts on east Asian economies were relatively modest during the GFC in comparison to the AFC (Jeasakul et al., 2014, Genberg, 2017).

Our research questions originate from this. That is, did the capital controls that had been strengthened after the AFC explain the modest effects of international capital flow shocks on east Asian economies during the GFC period? Indeed, there has been some debate over the policy effectiveness of post−AFC reforms of AFC economies in terms of financial stability. In particular, capital account liberalization policy could increase financial flexibility against international capital flow shocks, but also increase volatility through risk sharing via synchronized financial markets (Aghion et al., 2004, Obstfeld et al., 2009, Villafuerte et al., 2015).

To answer this question, we compare the cross−country and cross−period variation in the effects of net capital outflow shocks on ASEAN+3 countries with and without the AFC experience between two sub−period samples (AFC episodes and GFC episodes). In particular, we define "AFC economies" as ASEAN+3 economies that had experienced domestic currency depreciation more than 50% against the US dollar during the AFC. Korea, Indonesia, the Philippines, and Thailand[13] are classified as

[13] Though satisfying the criterion, Malaysia and Lao PDR are excluded from

AFC economies under this criterion. Meanwhile, Vietnam and Japan[14], members of ASEAN+3 unbelonging to AFC economies, are classified as "Non-AFC economies". Notably, Singapore and Hong Kong, which meet Non-AFC criterion, are not included in Non-AFC economies since their capital inflows were much larger and more volatile than those of the other ASEAN+3 economies, which would have distorted the results. Lao PDR, Malaysia, China, Cambodia, Brunei, and Myanmar are also excluded from our analysis due to insufficient data for VAR analysis. For sub-period samples, we define the "AFC episodes" from 1993:Q4[15] to 2004:Q4 and the "GFC episodes" from 2001:Q1 to 2015:Q4.

We estimate a Bayesian panel VAR model for three sub groups : (i) AFC economies during the AFC episodes, (ii) AFC economies during the GFC episodes, and (iii) Non-AFC economies during the GFC episodes. First, we identify if there is cross-period difference in the effects of net capital outflow shocks on real GDP growth rate for AFC economies between the AFC and GFC episodes. Next, we compare the cross-country difference in the effects of net capital outflow shocks between AFC economies and Non-AFC economies during the GFC episodes. Due to lack of available data series, we cannot identify cross-period difference for Non-AFC economies between the AFC and GFC episodes.

Our findings are summarized as follows. First, we find that the negative effects of net capital outflow shocks on real GDP growth rate for AFC economies are curtailed from the AFC period to the GFC period. On the other hand, output contracts more persistent during the GFC period for Non-AFC economies than for AFC economies. It is noteworthy that AFC economies had reinforced capital control policies after the AFC, whereas Non-AFC economies had expanded capital openness for recent two decades.

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our analysis due to lack of quarterly data.

[14] Though satisfying the criterion, China, Cambodia, Brunei, and Myanmar are excluded from our analysis due to lack of quarterly data.

[15] This is the earliest date available to construct strongly balanced panel data for AFC economies.

This finding suggests that post-AFC reforms of AFC economies were effective to raise their resilience to net capital outflow shocks during the GFC period.

However, there might be another source of this cross-country and cross-period difference in the effects of capital flow shocks such as compositional changes in the main driver of capital inflow reversals from the AFC to the GFC episodes. In order to identify the effects of capital inflow shocks by component, we extend Bayesian panel VAR models by breaking down capital inflows into their component parts. We find that shocks to individual components of capital inflows have different effects on both AFC economies and Non-AFC economies. For AFC economies, capital inflow shocks led by equity in portfolio investments have the most significant negative effects on real GDP growth rate. For Non-AFC economies, capital inflow shocks driven by direct investments decrease real GDP growth rate most significantly.

Note that equity in portfolio investments was not the main source of capital inflow reversals for AFC economies during the AFC episodes, but it became the second largest sources during the GFC episodes. Although the main drivers of capital flow shocks had changed in a way of strengthening their negative effects on real GDP growth rate during the GFC episodes, AFC economies recieved less negative effects from capital flow shocks. On the other hand, direct investment, which has the largest negative effects of real GDP growth rate for Non-AFC economies, was not the main sources of capital flow reverals for them during the GFC episodes. Nonetheless, output contracts more persistent in Non-AFC economies than in AFC economies during the GFC episodes. As a result, even when considering compositional changes in capital flow shocks, we can still conclude that post-AFC reforms are beneficial in preventing another AFC.

The contributions of this paper to existing literatures are summarized as follows. To begin, this is the first study to examine the time-varying effects of international capital flow shocks with a particular emphasis on the AFC and GFC episodes. While the effects of the GFC on ASEAN+3 countries are frequently studied, only a few studies link the AFC episode to the GFC episode. Jeasakul et al. (2014) attempted to examine the AFC economies' resilience to the GFC, but their analysis was based entirely on simple OLS regression, which could suffer from omitted variable or endogeneity issues.

Additionally, these exercises can be used to infer the effectiveness of capital control policies adopted by AFC economies following the AFC. Although international capital flow shocks are the main sources of financial volatilities for small open economies, many countries have suffered from setting optimal policy responses to a surge in capital flows. The results of this paper provides empirical evidences to support the effectiveness of post-AFC reforms, which would serve as a reference point for other small open economies.

Third, we provide empirical evidences that each component of capital inflows has distinct effects on output for AFC and Non-AFC economies. In particular, equity in portfolio investments has the most significant negative effects on output for AFC economies, whereas direct investments have the most significant negative effects on output for Non-AFC economies. This finding shows that capital control policies would be more efficient if they focused on the component of capital flows that has the significant detrimental impact on output rather than strengthening control over all components.

The remainder of the paper is structured in the following manner. Section 3.2 summarizes the relevant literature. Section 3.3 summarizes the data's characteristics. Section 3.4 describes the empirical methods used and summarizes the findings. The paper concludes with Section 3.5.

3.2. Related Literature

International capital flow shocks are the primary source of business and financial volatility in small open economies. As a result, numerous studies have been conducted to examine the effects of capital flow shocks on small open economies. Aghion et al. (2004) demonstrated through a DSGE model that international capital flow shocks could be a source of financial instability in small open economies, particularly those in the early stages of capital account liberalization. Bonciani and Ricci (2020) used a local projection model to determine the statistically significant negative effects of global financial uncertainty shocks on 40 small open economies' macroeconomic variables. Tomura (2010) demonstrated that the degree to which an economy was open to international capital flows was critical in determining how the boom-bust cycle in housing markets is formed. Kim and Kim (2013) estimated a structural VAR model to find a statistically significant correlation between boom-bust cycles and capital inflows in a sample of Asia-Pacific countries since the 1990s.

Meanwhile, a body of research has examined the effects of capital control policies implemented in response to increased financial distress during the AFC or GFC. Obstfeld (2009) summarized relevant empirical studies and concluded that there was scant evidence that financial liberalization would improve east Asian countries' economic performance. Fratzscher (2012) developed a factor model to account for the divergent capital flow patterns observed in each economy during the GFC and demonstrated that differences in the quality of domestic institutions might be a significant source of this heterogeneity. Forbes (2012) examined capital flow trends in Asia's economies and concluded that strengthening the domestic financial system rather than directly reducing total capital flows might be more appropriate.

The most relevant research for our paper is Jeasakul et al. (2014), which compared the effects of international capital flow shocks on east Asian economies during the AFC and the GFC. While Jeasakul et al. (2014) provided cumulative analysis of capital flow patterns in east Asian economies, the statistical method used was simple OLS, which might have issues with omitted variables and endogeneity. The key difference with our paper is that we address this endogeneity problems by using a Bayesian panel VAR model with a recursive identification. Moreover, our paper compare the effects of net capital outflow shocks for east Asian countries with and without the AFC experience, which would contribute to analyze how policy responses to capital flow shocks would lead to time-varying effects of international capital flow shocks among east Asian countries.

3.3. Data and summary statistics

In this section, we briefly overview the characteristics of international capital flows, macro-economic variables, and capital account openness measures by Chinn and Ito (2008) for ASEAN+3 economies from 1990 to 2018, the sample periods including the AFC and GFC episodes. Appendix 3.A details the sources of data used. To analyze cross-country variation, we classify ASEAN+3 economies into AFC economies (with a depreciation rate of domestic currency against the U.S. dollars greater than 50%) and Non-AFC economies (with a depreciation rate less than 50%) in 1998. Hong Kong and Singapore are excluded from Non-AFC economies due to their extreme volatility in capital inflows (as a percentage of trend GDP), but are included in the average of the ASEAN+3 economies. Lao PDR and Malaysia are excluded from AFC economies, and Myanmar, China, Cambodia, and Brunei are ruled out from Non-AFC economies as they have insufficient quarterly data for VAR analysis in the next chapter. Table 3.1 summarizes the country classification of ASEAN+3 economies.

The simple averages of net capital inflows (capital inflows minus capital outflows) as a percentage of trend GDP[16] for AFC, Non-AFC, and ASEAN+3 economies from 1990 to 2018 are shown in Figure 3.1[17]. Note that blue shaded area represents the outbreak

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[16] Since financial account is measured in current US dollars, the trend GDP is calculated by applying a Hodrick–Prescott filter to the nominal GDP (current US dollars) with a value of 100 (for yearly data).

[17] Capital flows for Non-AFC economies begin from 1996, which is the

Table 3. 1. Classification of ASEAN+3 Economies

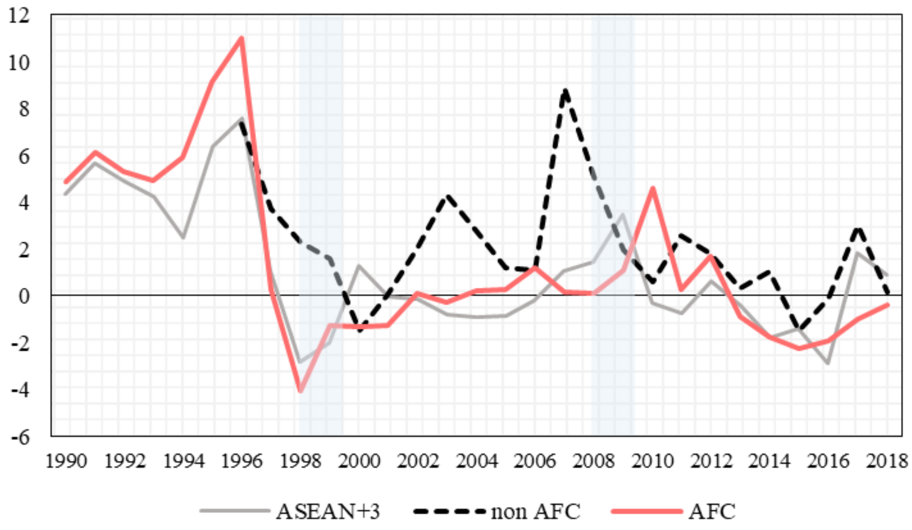
Group	Countries	Notes
AFC	Korea, Indonesia, Philippines, Thailand,	Economies that experienced more than 50% depreciation of their currency against the U.S. dollars during the Asian Financial Crisis, except for Lao PDR and Malysia
Non-AFC	Vietnam, Japan,	ASEAN+3 economies unbelonging to the AFC, except for Hong Kong, Singapore, Mynammar, China, Cambodia, and Brunei
ASEAN+3	AFC + Non-AFC + Malaysia, Lao PDR, Myanmar, China, Cambodia, Brunei, Singapore, Hong Kong	Includes Hong Kong, Singapore, and those belonging to AFC economies (Lao PDR, Malaysia) and Non-AFC economies (Mynamar, China, Cambodia, Brunei) that do not have sufficient quarterly data for VAR analysis.

of AFC and GFC. Three significant features are pronounced in the dynamics of net capital inflows in the sample countries. First, all countries experienced net capital inflow reversals during the AFC and GFC. In particular, the degree of net capital inflow reversal during the AFC episodes was about 15 percentage points in AFC economies, which was the greatest among the country groups.

Second, for AFC economies, the degree of net capital inflow reversals during the GFC episodes was significantly smaller than those during the AFC episodes. Specifically, the ratio of net capital inflows to trend GDP in AFC economies fell from 11.0 percent to −4.05 percent during the AFC episodes (from 1996 to 1998), whereas the same ratio remained around 0.14 percent to 0.09 percent during the GFC episodes (from 2007 to 2008). In particular, net capital inflows to trend GDP ratio had been almost zero percent for AFC economies from 2002 to 2008, which was partly due to tightening capital controls of AFC economies after the AFC.

earliest data available.

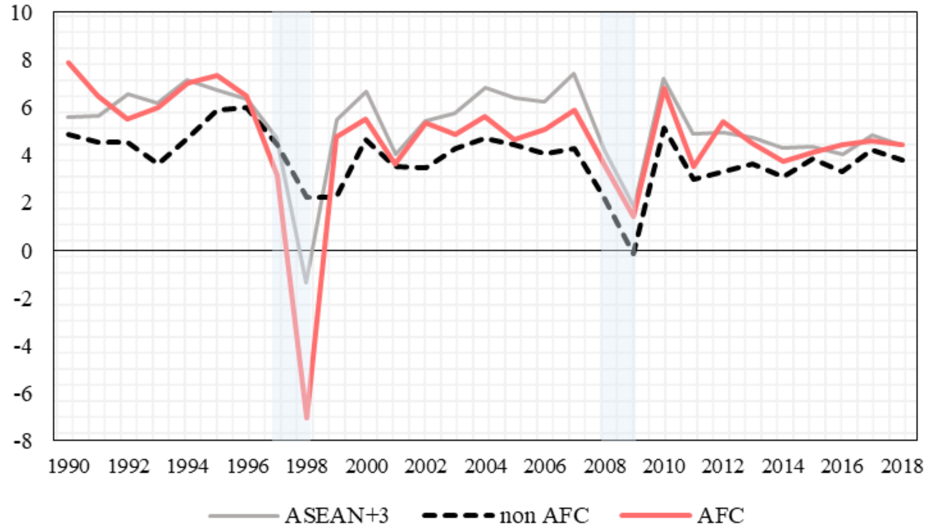
Figure 3. 1. Net Capital Inflows (as a Percentage of Trend GDP)



Third, contrary to AFC economies which showed a stable net capital inflows from 2007 to 2008, Non-AFC economies experienced a high degree of capital inflow reversals throughout the GFC episodes. Especially, the net capital inflows to trend GDP ratio fell from 8.89 percent in 2007 to 0.69 percent in 2010. Although net capital inflows in Non-AFC economies did not turn into negative during the GFC episodes, this rapid decrease in net capital inflows could be a source of financial instabilities in these countries.

The simple averages of real GDP growth rate for AFC, Non-AFC, and ASEAN+3 economies from 1990 to 2018 are shown in Figure 3.2. Those patterns found in net capital flows were replicated in the real GDP growth rate. Real GDP growth rate fell in all country groups during the AFC and GFC, but the degree of decline was greatest in AFC economies during the AFC, from 3.15 percent in 1997 to -7.05 percent in 1998. It is interesting that AFC economies experienced weaker output declines during the GFC than those during the AFC. In particular, GDP growth rates declined from 5.87 percent in 2007 to 1.44 percent in 2009. The similar patterns of capital flows and outputs observed in AFC economies during the

Figure 3. 2. Real GDP Growth Rate (Annual Percentage)



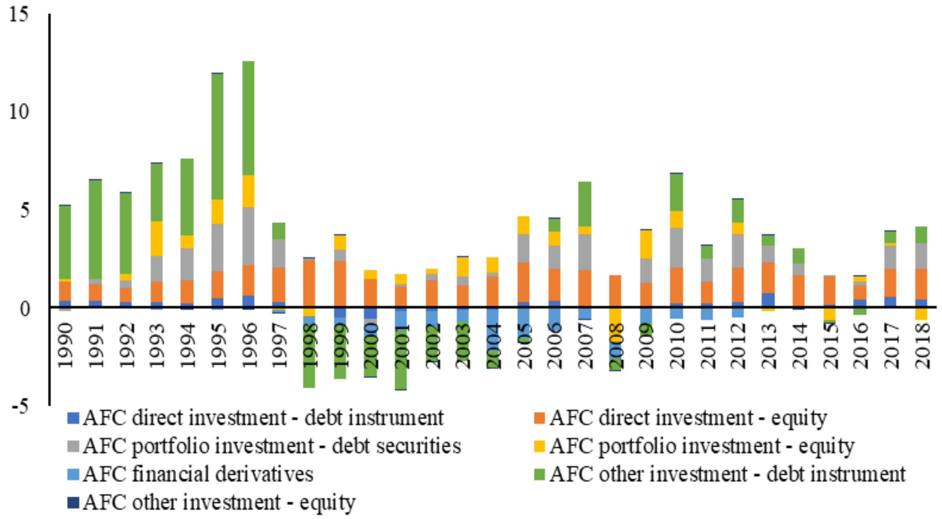
two financial crisis episodes suggest that the effects of capital flow shocks on output may have become weaker for AFC economies from the AFC to the GFC episodes.

However, for non-AFC economies, the decrease in real GDP growth rate became more pronounced from the AFC to the GFC. In particular, real GDP growth rate during the AFC fell from 4.45 percent in 1997 to 2.23 percent in 1998, but then fell more during the GFC, from 4.26 percent in 2007 to −0.15 percent in 2009.

Capital inflow reversals may have different macroeconomic effects depending on which component is the primary source. Thus, the composition of capital inflows into AFC and Non-AFC economies is also analyzed. The components of capital inflows (as a percentage of trend GDP) for AFC economies from 1990 to 2018 are depicted in Figure 3.3. We focus on the which component of capital inflows was the main driver of capital flow reversals during the AFC and GFC episode.

During the AFC, the largest component of capital inflow reversals in AFC economies was debt instruments in other

Figure 3. 3. Components of Capital Inflows in AFC economies
(as a Percentage of Trend GDP)

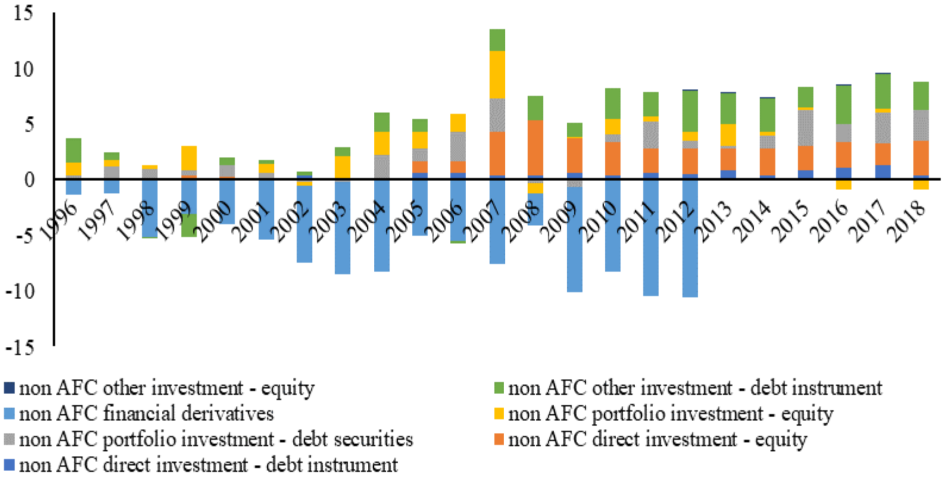


investments[18], decreasing by 4.15 percentage points from 1997 to 1998. And the second largest component was debt securities in portfolio investments, declining by 1.34 percentage points from 1997 to 1998. During the GFC episode, the primary component of capital inflow reversals for AFC economies was still debt instruments in other investments, decreasing by 2.85 percentage points from 2007 to 2008. However, equity in portfolio investments, which was the fifth largest component of capital inflow reversals during the AFC, became the second largest component of capital inflow reversals during the GFC, decreasing by 2.09 percentage points from 2007 to 2008.

Figure 3.4 shows the components of capital inflows (as a percentage of trend GDP) for Non-AFC economies from 1990 to 2018. Contrary to AFC economies, the main sources of capital inflow reversals in Non-AFC economies were financial derivatives for the AFC, decreasing by 3.98 percentage points from 1997 to 1998. Those for the GFC episodes were equity in portfolio

[18] Capital flows from other investment capture those from deposit and lending transactions of banks (Koepke, 2020).

Figure 3. 4. Components of Capital Inflows in Non−AFC economies
(as a Percentage of Trend GDP)

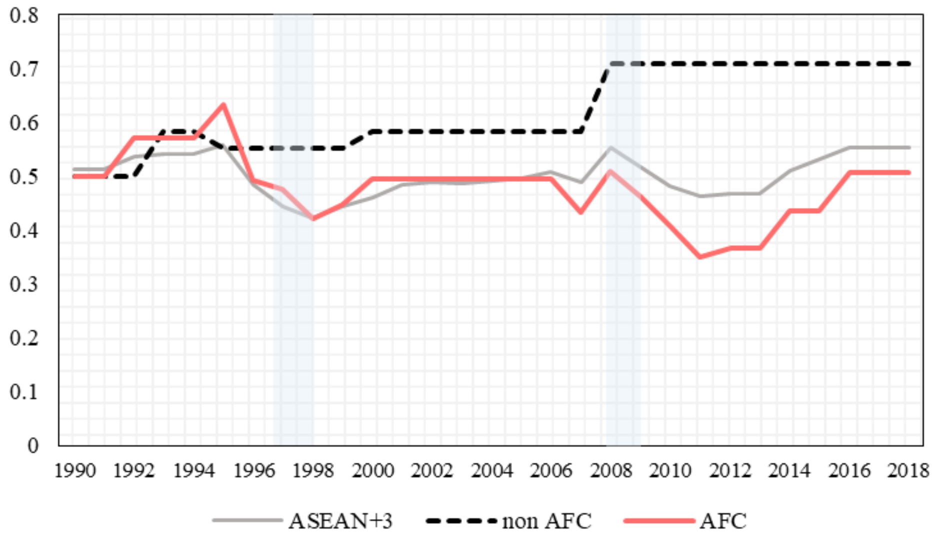


investments, decreasing by 5.24 percentage points from 2007 to 2008. It is interesting that for both AFC and Non−AFC economies, equity and debt securities in portfolio investments, which were not significant during the AFC, became the main sources of capital flow reversals during the GFC. To analyze the implication of these compositional changes to capital flow shocks, the independent effects of shocks to each component of capital inflows on the AFC economies will be examined in greater detail in Section 3.4.

Lastly but most importantly, the Chinn−Ito index is analyzed to examine capital control policies on international capital flows of AESEAN+3 countries. The Chinn−Ito index measures capital account openness, which is primarily determined by capital flow liberalization/management/control policies. Among the various capital account liberalization/control measures, the Chinn−Ito index is available for a relatively long period of time and is constructed using the same standard for each economy, making it easier to compare the degree of capital account openness across economies. The index value is between 0 and 1. A greater value indicates a greater degree of capital account openness.

Figure 3.5 illustrates simple average values of the Chinn−Ito

Figure 3. 5. Chinn–Ito Measure



index for the ASEAN+3, AFC, and Non–AFC economies. Throughout the 1990s, AFC economies had similar capital account openness with Non–AFC economies, but following the AFC episode, capital account openness began to decline, decreasing from 0.63 in 1995 to 0.42 in 1998 and 0.35 in 2011. It can be conjectured that AFC economies had tightened capital controls as a result of the AFC and GFC. In contrast, Non–AFC economies, which did not suffer serious output declines after the AFC, had liberalized capital accounts throughout the sample periods. In particular, Chinn–Ito index of Vietnam increased from 0 in 1990 to 0.42 in 2008. In chapter 4, we would examine whether these capital controls were successful in decreasing the volatility generated by capital flow shocks in AFC economies during the GFC episode.

3.4. Cross–country and cross–period difference in the effects of international capital flow shocks for AFC economies

From the previous section, we find that net capital inflows and real GDP growth rate of AFC economies decreased less during the

GFC than during the AFC. These changes could have resulted from more prudent capital control polices of AFC economies, which the Chinn−Ito index implies. Moreover, the changes in the composition of capital flows during the AFC and GFC episode might have a role in these time varying effects of international capital flow shocks.

To analyze if there are cross−country and cross−period differences in the effects of net capital outflow shocks for AFC and Non−AFC economies from the AFC to the GFC episodes, we estimate Bayesian panel VAR model and compare cross−country and cross−period differences in impulse response functions on three different sub−samples given by

(i) AFC economies around the AFC episode
(from 1993:Q4 to 2004:Q4[19])
(ii) AFC economies around the GFC episode
(from 2001:Q1 to 2015:Q4)
(iii) Non−AFC economies around the GFC episode
(from 2001:Q1 to 2015:Q4)

The sample data for Non−AFC economies begins from 1999:Q4, hence we are unable to estimate the impulse responses for Non−AFC economies around the AFC episode. Our comparative analysis is conducted in two steps. First, we compare the impulse response functions from (i) and (ii) to identify how the effects of net capital outflow shocks for AFC economies changed from the AFC to the GFC episodes. Second, we compare those from (ii) and (iii) to analyze whether AFC economies which conducted post−AFC reforms were better able to cope with capital flow shocks during the GFC episodes than Non−AFC economies.

3.4.1. Methodology

In order to analyze the effect of net capital outflow shocks for

[19] We consider before and after seven years after the crisis for each sample.

AFC and Non−AFC economies, we use the Bayesian pooled estimator[20], which assumes homogeneous dynamic coefficients across units. Note that the Bayesian pooled estimator is the simplest form of Bayesian panel VAR models, which relaxes general properties of panel models such as dynamic interdependencies or static interdependencies. However, our panel data does not include large open economies such as the United States or China which have strong linkages to export−oriented small open economies. Moreover, the objective of our analysis is to estimate the average response in each sub−group with and without AFC experiences to net capital outflow shocks. Thus, the Bayesian pooled estimator is appropriate for our analysis. The Bayesian panel VAR is estimated as follows.

$$\begin{pmatrix} y_{1,t} \\ y_{2,t} \\ \vdots \\ y_{N,t} \end{pmatrix} = \begin{pmatrix} A_1 & 0 & \dots & 0 \\ 0 & A_1 & \dots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \dots & A_1 \end{pmatrix} \begin{pmatrix} y_{1,t-1} \\ y_{2,t-1} \\ \vdots \\ y_{N,t-1} \end{pmatrix} + \cdots + \begin{pmatrix} A_p & 0 & \dots & 0 \\ 0 & A_p & \dots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \dots & A_p \end{pmatrix} \begin{pmatrix} y_{1,t-p} \\ y_{2,t-p} \\ \vdots \\ y_{N,t-p} \end{pmatrix} + \begin{pmatrix} \varepsilon_{1,t} \\ \varepsilon_{2,t} \\ \vdots \\ \varepsilon_{N,t} \end{pmatrix} \tag{3.1}$$

$$\varepsilon_{i,t} \sim \mathrm{N}(0, \textstyle\sum_c), \text{ with } \textstyle\sum_{ii,t} = \mathrm{E}(\varepsilon_{i,t}\varepsilon'_{i,t}) = \textstyle\sum_c \text{ and } \mathrm{E}\left(\varepsilon_{i,t}\varepsilon'_{j,t}\right) = 0 \text{ for } \mathrm{i} \neq \mathrm{j}$$

where $\mathbf{N}$ denotes the number of cross−sectional units and p stands for a maximum number of lag order for the panel VAR model. We use two lags in Bayesian panel VAR model as the number of samples in each sub−group is limited. A standard Litterman/Minnesota prior is used to derive the likelihood function in Bayesian estimation (Dieppe et al., 2016). In particular, a standard set of the hyperparameters are employed following $\lambda_1 = 0.1$ (overall tightness), $\lambda_2 = 0.5$ (the relative cross−variable weight), $\lambda_3 = 1$ (the lag decay), and $\lambda_4 = 100$ (exogenous variable tightness).

[20] We estimate Bayesian panel VAR models using the BEAR toolbox provided by the ECB.

For quarterly data of AFC and Non-AFC economies, three variables are considered: (i) CUR (current account as a ratio to trend GDP), (ii) CAP (net capital outflows as a ratio to trend GDP), and (iii) RGDP (year-over-year growth rate[21] of real GDP). CUR is included to account for endogenous responses of capital flows by current account imbalances. CAP is included to identify the effects of capital flow shocks, which is the primary objective of our analysis. Note that for CAP, net capital outflows (capital outflows minus capital inflows) are employed instead of net capital inflows. Thus, the rapid increase in CAP can be interpreted as a financial crisis episode. The trend GDP is calculated by HP (Hodrick-Prescott) filter with λ=1,600 (quarterly data) on current US dollar GDP. RGDP is included to infer the real economy's response to international capital flow shocks.

As an identification strategy, a recursive structure on contemporaneous structural parameters is assumed, which was suggested by Sims (1980). In particular, Choleksy factorization with the order {CUR, CAP, RGDP} is adopted, where the contemporaneously exogenous variables are ordered first.

The reasons behind the ordering {CUR, CAP, RGDP} are as follows. First, CUR is assumed contemporaneously exogenous to CAP and RGDP, which helps to identify more exogenous components of CAP movements by excluding endogenous movements of CAP caused by current account fluctuations. Current account imbalances are automatically financed by capital flow movements, and we would like to exclude such endogenous responses of CAP. Second, CAP is assumed contemporaneously exogenous to RGDP to infer the effects of shocks to CAP on RGDP, including the contemporaneous effects of shocks to CAP on RGDP within a year.

These identifying assumptions are motivated by the findings of Kim, Kim, and Wang (2004), who examined the effects of capital

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[21] Since Vietnam's quarterly GDP data was only available in the form of year-over-year growth rates, we use year-over-year growth rates of real GDP for output.

flow shocks on a variety of macroeconomic variables in Korea. We test alternative identifications with different orders and variables, but the primary results are qualitatively similar.

3.4.2. Impulse response to net capital outflow shocks on different sub-groups

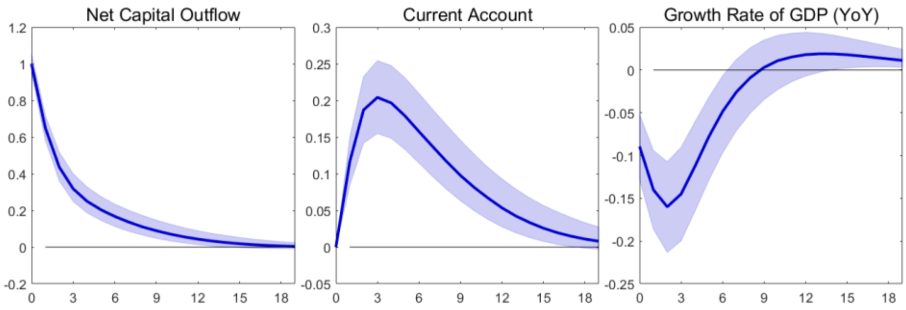
Figure 3.6 shows the impulse responses to a one-percentage-point increase in net capital outflow to trend GDP ratio (CAP) for three sub-groups : (i) AFC economies around the AFC episodes, (ii) AFC economies around the GFC episodes, and (iii) Non-AFC economies around the GFC episodes. Non-cumulative impulse response functions are reported for CUR, CAP, and RGDP. Note that non-cumulative impulse responses for RGDP represents the percentage point changes in the year-over-year growth rate of real GDP. We do not report cumulative impulse response functions for RGDP as there might be seasonality issues.

Across three sub-groups, shocks to net capital outflows have a significant positive effect on current account to trend GDP ratio (CAP), which is consistent with economic theories on international trade that a deficit in the capital account is offset by a surplus in the current account. In particular, CAP increases by up to 0.2 percentage points for AFC economies in the AFC episodes, 0.04 percentage points for AFC economies in the GFC episodes, and 0.13 percentage points for Non-AFC economies in the GFC episodes.

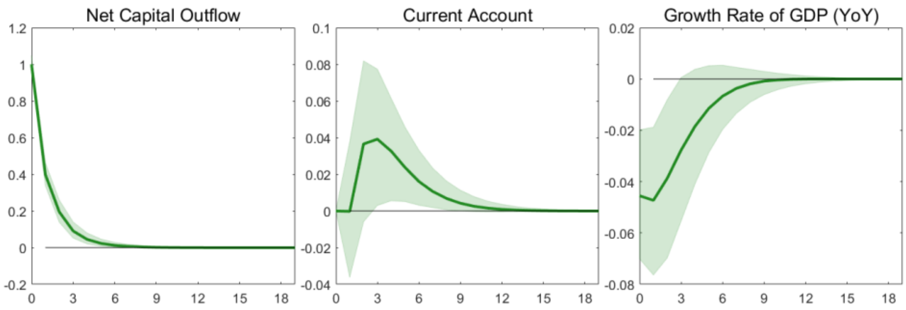
Furthermore, net capital outflow shocks have a significant negative effect on year-over-year growth rate of real GDP (RGDP) for all sub-groups. However, the degree of the decrease is varied by each sub-group. To be more specific, RGDP decreases by up to 0.16 percentage points for AFC economies in the AFC episodes, 0.05 percentage points for AFC economies in the GFC episodes, and 0.05 percentage points for Non-AFC economies in the GFC episodes.

Figure 3.7 presents the impulse response functions to a net capital outflow shocks for two sub-groups in a single figure to

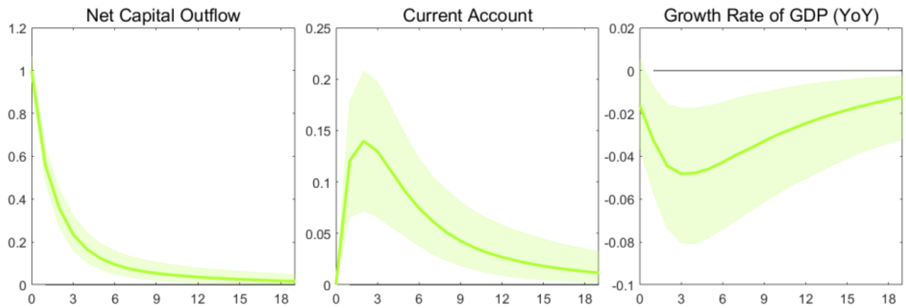
Figure 3. 6. Impulse response functions,
the effect of net capital outflow shocks on macroeconomic variables.



(1) AFC economies – AFC episodes (from 1993:Q4 to 2004:Q4)



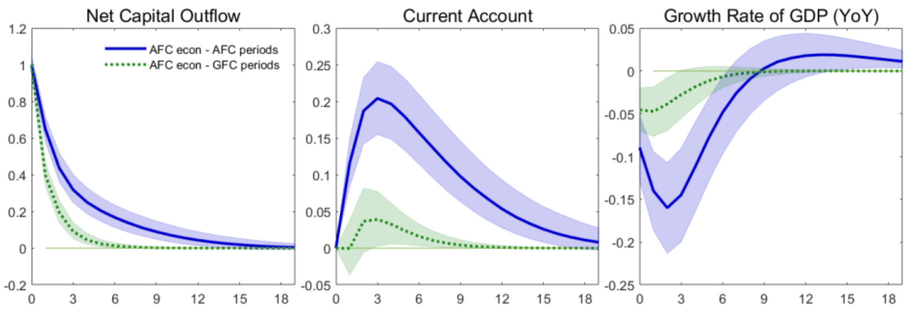
(2) AFC economies – GFC episodes (from 2001:Q1 to 2015:Q4)



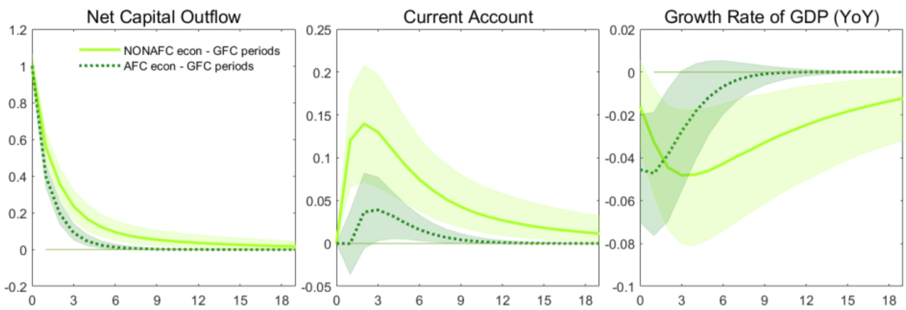
(3) Non–AFC economies – GFC episodes (from 2001:Q1 to 2015:Q4)

Notes : An international capital flow shock is defined as an increase in the percentage ratio of net capital outflow to trend GDP by 100 basis points. The effect is estimated on three sub-groups (1. AFC in AFC, 2. AFC in GFC, 3. Non–AFC in GFC). The shaded area shows 68 percent confidence intervals calculated by Bayesian estimation with 5,000 iterations (with 1,000 as a burn–in) identify the cross−period difference (AFC economies during the AFC and GFC episodes) and the cross−country difference (AFC economies and Non−AFC economies during the GFC episodes). It is easily noticed that the negative impact of a net capital outflow shock on RGDP becomes weaker over time from the AFC to the GFC episodes for AFC economies. Furthermore, the negative effects of net capital outflow shocks on GDP growth rates are less persistent in AFC economies than in non−AFC economies throughout the GFC episodes: significantly negative for the first three quarters in AFC economies but for the entire horizon in Non−AFC economies.

Figure 3. 7. Impulse response functions,
the effect of net capital outflow shocks on macroeconomic variables



(1), (2) AFC economies − AFC and GFC episodes



(2), (3) AFC and Non−AFC economies − GFC episodes

Notes : An international capital flow shock is defined as an increase in the percentage ratio of net capital outflow to trend GDP by 100 basis points. The effect is compared between the periods (1. AFC in AFC, 2. AFC in GFC) and countries (2. AFC in GFC, 3. Non−AFC in GFC). The shaded area shows 68 percent confidence intervals calculated by Bayesian estimation with 5,000 iterations (with 1,000 as a burn−in)

These findings imply that post−AFC reforms would have helped AFC economies build resilience to international capital flow shocks after the Asian financial crisis. However, there might be another source of these cross−country and cross−period changes. In particular, we find that main drivers of capital inflow reversion changed from the AFC episodes to the GFC episode for both economies. So we check the robustness of the results focusing on compositional changes in capital flow shocks in the nest subsection.

3.4.3. Impulse response to negative shocks on each component of capital inflows

As discussed in Section 3.3, the primary component of capital inflow reversals for AFC economies during the AFC and GFC episodes was debt instrument in other investments. Moreover, equity in portfolio investments, which had accounted for small shares of capital inflow reversals during the AFC episodes, became the second largest component in the GFC episodes. Meanwhile, the main sources of capital inflow reversals in non−AFC economies shifted from financial derivatives during the AFC episodes to equity in portfolio investments during the GFC episodes.

To examine how these compositional changes would affect the impacts of capital flow shocks on real GDP growth rate, we extend the three−variable baseline models into the four−variable models. To be more specific, we decompose CAP (net capital outflows as a ratio to trend GDP) into CAPI by component (the minus of capital inflows by component as a ratio to trend GDP) and CAPO (the minus of capital outflows as a ratio to the trend GDP). Note that capital inflows by components and capital outflows are multiplied by a negative sign to maintain consistency with the baseline models. That is, a one−point increase in CAPI corresponds to a one−point decline in capital inflows per component, consistent with a one−point increase in net capital outflows (CAP).

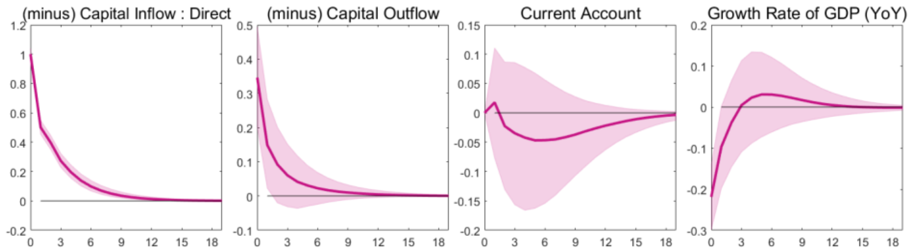
Capital inflows are decomposed into four major components: i) direct investments (debt instrument + equity), ii) other investments (debt instrument + equity), iii) debt securities in portfolio investments, and iv) equities in portfolio investments.[22] Four distinct models are estimated for each component of capital inflows, and the models follow the same recursive structure and lag ordering with our baseline model (i.e. CUR, CAPI by component, CAPO, and RGDP with two lags). The extended models are estimated for AFC economies from 1993:Q4 to 2015:Q4, and Non-AFC economies from 1999:Q4 to 2015:Q4, which are the longest panel data available around the AFC and GFC episodes.

Figure 3.8 shows the impulse responses of each variable to a one-percentage-point increase in CAPI by component for the AFC economies, which corresponds to a one-percentage-point decline in each component of capital inflows. Non-cumulative impulse responses are reported for CAPI, CAPO, CUR and RGDP. Note that non-cumulative impulse responses for RGDP represent the percentage point changes in year-over-year growth rate of real GDP. In all the models, the minus of capital outflows to trend GDP ratio (CAPO) and current account to trend GDP ratio (CUR) shows positive responses to shocks in the minus of capital inflows by component to trend GDP ratio (CAPI by component), which is consistent with standard theory on international trade that a surplus in the current account offsets a deficit in the capital account.

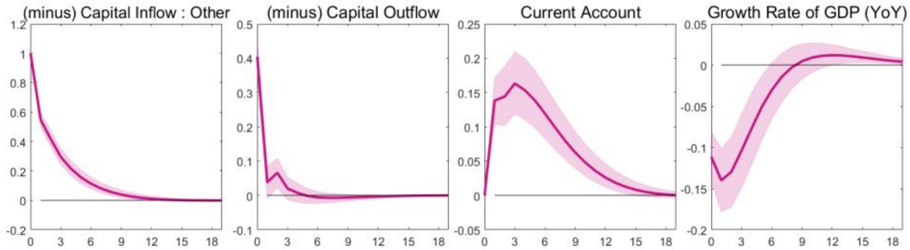
Note that negative shocks to other investments and equity in portfolio investments have a significant negative effect on RGDP in AFC economies, but other shocks do not. In particular, negative shocks to other investments reduce real GDP growth rate by up to 0.14 percentage points, whereas negative shocks to equity in portfolio investments reduce it by up to 0.41 percentage points. The estimation results suggest that capital inflow reversals through

[22] Debt instruments of other investment take up the most part of other investment but equity instruments of other investment take up only a small part of other investment. Thus, there is no need to separate these investments into further details. On the other hand, portfolio investments into two types, equities and debt securities, because these two types of flows often show different trends and magnitude, and these types of flows increase fast over time and have become very important in recent years.

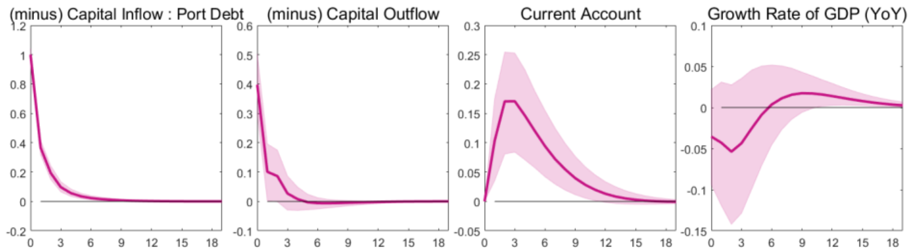
Figure 3. 8. Impulse response functions, the effect of capital inflow shocks by each component on macroeconomic variables for AFC economies



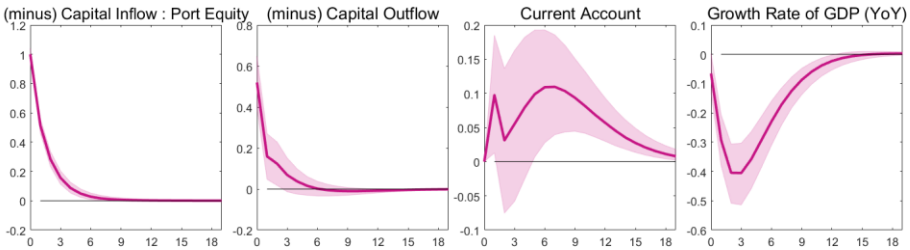
(1) Direct Investment



(2) Other Investment



(3) Portfolio Investment: Debt Securities



(4) Portfolio Investment: Equities

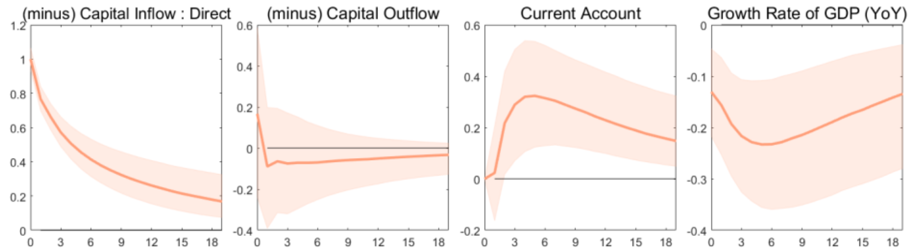
Notes : An capital inflow shock by component is defined as an increase in the percentage ratio of (negative) capital inflow by component to trend GDP by 100 basis points. The effect is estimated on AFC economies for the period from 1993:Q4 to 2015:Q4. The shaded area shows 68 percent confidence intervals calculated by Bayesian estimation with 5,000 iterations (with 1,000 as a burn-in) equity in portfolio investments have the largest negative impact on real GDP growth rate in AFC economies.

It is worth noting that equity in portfolio investments, which was the fifth largest component of capital inflow reversals during the AFC episodes, became the second largest cause of capital inflow reversals during the GFC episodes for AFC economies. We found in the previous section that the negative effects of net capital outflow shocks on real GDP growth rate for AFC economies during the AFC episodes became weaker during the GFC episodes. As a result, we conclude that post-AFC reforms were successful in increasing their resilience to capital flow shocks, even while capital inflow reversals through equity in portfolio investments, which have the largest negative impact on real GDP growth rate for AFC economies, got stronger during the GFC period.

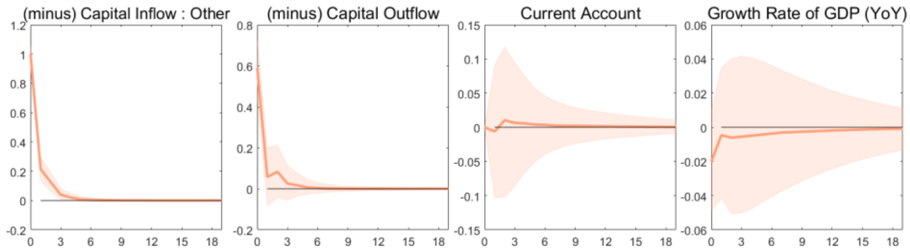
Figure 3.9 shows the impulse responses of each variable to a one-percentage-point increase in CAPI by component for the Non-AFC economies, which is equivalent to a one-percentage-point decline in each component of capital inflows. In contrast to the impulse reactions for AFC economies, which are consistent with standard macroeconomic theories, some puzzling results for Non-AFC economies are observed. In particular, despite negative shock in debt securities in portfolio investments, real GDP growth rate increase significantly. These results, however, are due to an equivalent increase in (negative) capital outflows, which offsets (negative) capital inflows of debt securities in portfolio investments.

Among all the components in capital inflows, negative shocks to direct investments and equity in portfolio investments cause real GDP growth rate to contract significantly for Non-AFC economies. In particular, negative shocks to capital flows in direct investments, which do not have significant negative effects on real GDP growth rate for AFC economies, decrease it the most by 0.23 percentage points for Non-AFC economies. This suggests that the component of capital inflows that has the biggest negative impact on real GDP growth rate may differ among countries. Meanwhile, negative shocks to equity in portfolio investments, which were the main

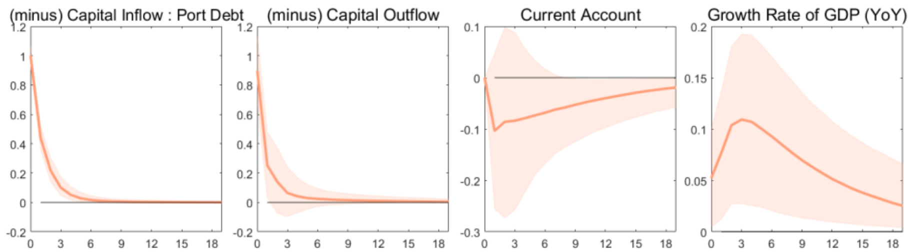
Figure 3. 9. Impulse response functions, the effect of capital inflow shocks by each component on macroeconomic variables for Non−AFC economies



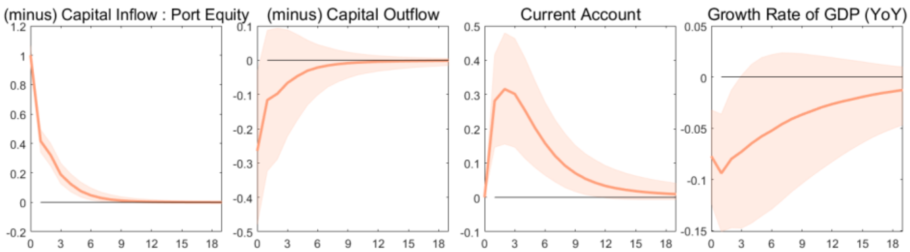
(1) Direct Investment



(2) Other Investment



(3) Portfolio Investment: Debt Securities



(4) Portfolio Investment: Equities

Notes : An capital inflow shock by component is defined as an increase in the percentage rati o of (negative) capital inflow by component to trend GDP by 100 basis points. The effect is e stimated on Non AFC economies for the period from 1999:Q4 to 2015:Q4. The shaded area s hows 68 percent confidence intervals calculated by Bayesian estimation with 5,000 iterations (with 1,000 as a burn−in) drivers of capital inflow reversals during the GFC episodes for Non−AFC economies, reduce real GDP growth rate significantly by 0.09 percentage points.

Note that net capital inflows to trend GDP ratio through debt instruments in direct investments, which have the most significant negative effects on output for Non−AFC economies, declined only by 0.05 percentage points in Non−AFC economies during the GFC (from 2007 to 2008). Even though capital inflow reversals during the GFC episode were not driven by the capital inflow component that has the most significant negative effects on output for Non−AFC economies, output falls more in Non−AFC economies after net capital outflow shocks than in AFC−economies. Even after accounting for changes in the composition of capital inflow reversals, we can conclude that post−AFC reforms in AFC nations are successful at reducing the negative effects of capital flow shocks on real GDP growth rate during GFC episodes.

3.5. Conclusion

The Asian financial crisis was one of the major global financial crises of the late 1990s, destabilizing Asian economies. AFC economics, in particular, the countries that experienced a depreciation of more than 50% of their local currency during the AFC period, suffered from severe economic recession as well as massive capital flight out of the countries. After the outbreak of the crisis, AFC economies had implemented post−AFC reforms, tightening capital controls outside the countries. After a decade, AFC economies appeared to have successfully withstood the global financial crisis, with only a slight decline in output and relatively stable capital outflows.

The purpose of this paper is to analyze whether these post−AFC reforms of AFC economies were effective in reducing the negative effects of international capital flow shocks on output during the GFC episode. Three variable Bayesian panel VAR models are estimated for three sub−groups ; (i) AFC economies during the

AFC episodes, (ii) AFC economies during the GFC episodes, (iii) Non−AFC economies during the GFC episodes.

AFC economies had strengthened capital controls for several years after the AFC whereas Non−AFC economies had liberalized capital accounts continuously. We find that for AFC economies, negative effects of net capital outflow shocks on real GDP growth rate during the AFC episodes became weaker during the GFC episodes. In terms of cross−country differences, we find that the negative effects of net capital outflow shocks on real GDP growth rate during GFC episodes are less severe in AFC economies than in Non−AFC economies. Considering that AFC economies had tightened capital controls whereas Non−AFC economies had increased capital openness, these findings support the effectiveness of post−AFC reforms against the capital flights.

To complement the baseline model results, the extended models augmented with each component of capital inflows are estimated for AFC economies and Non−AFC economies. We find that among various components of capital inflows, negative shocks in other investments and equity in portfolio investments have significantly negative effects on real GDP growth rate for AFC− economies. On the other hand, negative shocks in direct investments and equity in portfolio investments have significantly negative effects on real GDP growth rate for Non−AFC economies.

This finding confirms that the weaker effects of net capital outflow shocks in AFC economies during the GFC episodes are not due to compositional changes in capital flow reversals as equity in portfolio investment was the second largest sources of capital inflow reversals during the GFC episodes. Moreover, the finding that the components of capital inflow reversals that have the biggest negative impact on real GDP growth rate differ by country group has policy implications that selection and concentration on vulnerable sectors are required for an efficient capital control policy.

Overall, the result of this paper supports the effectiveness of post−AFC capital controls of AFC economies with respect to weakening the effects of international capital flow shocks during the

GFC episode. The findings of this article would benefit ASEAN+3 economies as well as other small open economies seeking appropriate policy responses to capital flight outside of their borders.

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Appendix 3.A. Data and Sources

Variable Name	Source	Notes
Trend GDP	World Bank	Uses an HP(Hodrick-Prescott) filter to nominal GDP(current US $) with $\lambda=100$.
Capital Inflows/Outflows/Net Inflows	IMF Financial Statistics : Balance of Payments	Sums each components of financial account (excludes reserves and related items) in terms of asset(outflows) and liabilities(inflows).
Components of Capital Inflows/Outflows	IMF Financial Statistics : Balance of Payments	
Regional Portion of Portfolio Investment Assets of the ASEAN+3	IMF Coordinated Portfolio Investment Survey(CPIS)	Total Portfolio Investment by Economy of Nonresident Issuer, End-of-Period
Chinn-Ito Measure	Chinn and Ito(2006)	
Real GDP Growth Rate	CEIC database, ADB database, Central banks, ASIA Regional Integration Center	
Real Consumption Growth Rate		
Real Investment Growth Rate		
Real Government Spending Growth Rate		
Current Account	IMF Financial Statistics : Balance of Payments	Excludes reserves and related items
Real Exchange Rate	BIS Statistics warehouse	
Nominal Exchange Rate		

Foreign Exchange Reserve Growth Rate	IMF, ADB database, ASIA Regional Integration Center	Excludes gold
Call Rate	International Financial Statistics (IFS)	Uses money market rate for call rates except for Myanmar, China, Lao, Vietnam, Brunei, Cambodia, where the data is not available. For these economies, policy rate data is used instead.

국문초록

통화정책, 부동산 시장 및 국제 자본 흐름에 대한 논문

본 학위논문은 통화정책 및 부동산 시장에 대한 두 개의 소논문과 국제 자본 흐름에 대한 하나의 소논문으로 이루어져 있다. 제1장에서는 부동산 가격과 가계 부채가 모두 상승하는 레버리지 부동산 호황 국면 (boom regime)과 그렇지 않은 국면 (normal regime)에서 통화정책의 효과가 어떤 차이를 갖는지를 비교분석 하였다. 이를 위해 실질 부동산 가격 갭과 가계 부채 갭의 최소값을 문턱 변수 (Threshold variable)로 사용하여 노르웨이, 한국, 캐나다의 3개 소국 개방 경제에 대해 문턱 구조적 벡터자기회귀모형 (Threshold SVAR model)을 추정하였다. 추정 결과 모든 국가에서 레버리지 부동산 호황 국면 동안 실질 부동산 가격 및 생산량에 대한 통화정책의 효과가 더 크고 유의한 것으로 나타났다.

제2장에서는 임대 주택 시장으로 논의를 확장하였다. 즉 가계가 금리 변동 이후 대출을 통한 주택 보유와 주택 임대 중 하나를 선택하는 것이 가능한 주택 소유 결정 채널 (homeownership decision channel)이 통화정책 전달 경로에 미치는 영향을 금리 충격에 경직적으로 반응하는 주택 임대료 (sticky housing rent)를 중심으로 분석하였다. 주택 소유 결정 채널을 포함하도록 확장한 뉴케인지언 모형 (New Keynesian model)을 통한 분석 결과 가계들이 금리 인상 충격 이후 대출을 통해 구입한 주택을 임대 주택으로 대체하는 주택 소유 결정 채널은 통화정책의 단기 효과를 약화시키지만 장기 효과는 강화하는 것으로 나타났다. 반면 비탄력적 주택 임대료는 임대 주택으로의 대체를 단기적으로 억제하여 통화정책의 단기 효과를 강화하였지만, 그 영향은 일시적이고 제한적이었다.

제3장에서는 아시아 금융 위기 경험 국가들의 외환위기 이후 금융개혁(post-AFC reform)이 국제 자본 흐름 충격에 따른 변동성을 완화시키는데 효과적이었는지를 실증 분석하였다. ASEAN+3 국가들을 아시아 금융 위기를 경험한 국가들 (AFC 경제)와 그렇지 않은 국가들 (비 AFC 경제)로 분류한 후, (i) AFC 기간의 AFC 경제, (ii) GFC 기간의 AFC 경제, (iii) GFC 기간의 비 AFC 경제라는 세 개의 소집단에 대해 베이지안 패널 벡터자기회귀 모형 (Bayesian panel VAR model)을

추정하였다. 추정 결과 AFC 경제에서 순자본 유출 충격이 경제성장률에 미치는 부정적인 영향은 AFC 기간에서 GFC 기간으로 가면서 약화된 것으로 나타났다. GFC 기간 동안 순자본 유출 충격의 경제 성장률에 대한 부정적인 영향 역시 AFC 경제에서 비 AFC 경제에 비해 덜 지속적인 것으로 나타났다. 이상의 결과는 외환위기 이후 금융 개혁이 국제 자본 흐름의 변동성을 완화하는데 효과적이었음을 시사한다.

주요어 : 부동산 호황, 가계부채, 임대주택, 통화정책, 국제자본흐름, 금융위기
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