University of New South Wales
School of Economics

Honours Thesis

# The Effects of Monetary Policy on the Australian Housing Market: A New Approach

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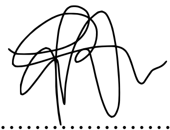
# Declaration

**I declare that this Honours thesis is my own work and that, to the best of my knowledge, it contains no material which has been written by another person or persons, except where acknowledgement has been made. This thesis has not been submitted for the award of any degree or diploma at the University of New South Wales Sydney, or at any other institute of higher education.**



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Carol Kong
22nd November, 2019

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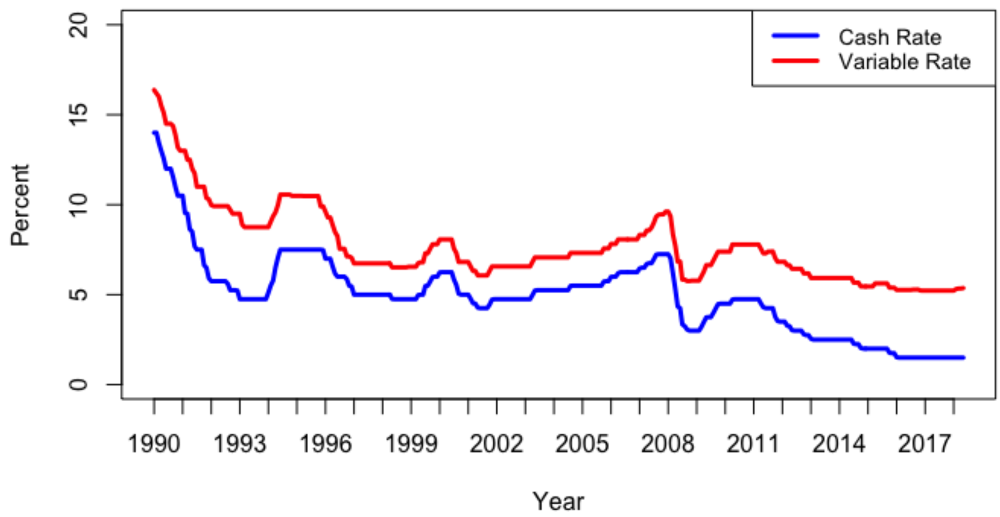
# Abstract

Monetary policy has long been viewed as an important determinant of house prices in Australia. Nonetheless, the effects of monetary policy inactions are less understood. In this paper, I seek to empirically estimate the effects of monetary policy on house prices in Australia, including the effects when the policy interest rate is unchanged. Specifically, I achieve this by adopting a recently proposed identification approach which identifies a monetary policy shock as a shift in the entire yield curve. Using city-level panel data spanning from 1999 to 2018, I further examine the extent to which house price responses to a monetary policy shock vary across the eight capital cities in Australia. Results confirm the presence of city heterogeneity in the effects of a change in the short-term interest rates —one of the three dimensions of an identified monetary policy shock. Interestingly, ceteris paribus, a cut in the cash rate does *not* always lead to a rise in house prices. I find that the response ultimately depends on how each policy announcement is perceived by market participants, which is reflected in the change in the shape of the yield curve. Furthermore, I explore the possibility that the state heterogeneity in the level of building approvals per capita can explain the differential monetary policy effects on city house prices. Impulse responses show that a city with a higher level of state building approvals per capita and hence a relatively more elastic housing supply tends to have a stronger negative house price response to a positive monetary policy shock. While this finding on city heterogeneity is somewhat puzzling, I conjecture that this may be explained by the differences in interest rate demand elasticity. Taken together, my results are valid to the extent that a positive shock in short-term interest rates leads to a sizeable negative impact on house price growth rates in Australia.

# Chapter 1
# Introduction

The severity of the 2008-2009 Global Financial Crisis (GFC) surely has awakened policymakers to how devastating the impacts of downturns in housing markets can be on the broader economy. It has also triggered a synchronised reduction in global interest rates. Since then, extensive attention has been drawn to the relationship between monetary policy and house prices. Understanding the effects of monetary policy on house prices is particularly important to Australia for two main reasons. Firstly, housing comprises a significant fraction of our nation's wealth. This implies that house price fluctuations can have huge impacts on aggregate consumption which is what the Reserve Bank of Australia (RBA) ultimately cares about (Dvornak and Kohler, 2007). Secondly, as seen in Figure 1.1, the policy interest rate has historically been closely linked to the standard variable mortgage rate. This strong positive relationship suggests that any changes in interest rate should pass through nearly one-for-one to the mortgage lending rate. With mortgages in Australia being dominated by variable rates, it is sensible to expect that house prices would experience stronger monetary policy transmission effects (Debelle, 2004; Mishkin, 2007).

Figure 1.1: Cash Rate and Standard Variable Mortgage Rate




Source: Reserve Bank of Australia

Despite the importance, more attention has been drawn to studying either the time series behaviour or determinants of house prices (Abelson et al., 2005; Otto, 2007). Relatively fewer studies are focused on directly examining the causal relationship between monetary policy and house prices in Australia. In this regard, this paper asks: what are the effects of monetary policy on house prices in Australia?

Answering this question, however, is not easy given the recent policy rate behaviour. It can be observed in Figure 1.1 that there have been extended periods of a constant cash rate, most notably nearly the end of sample period. This makes identifying a monetary policy shock more difficult. Moreover, as I will demonstrate, using traditional identification approaches might miss some important effects that a monetary policy announcement has on house prices. In light of this challenge, I ask a more specific question: what are the effects of monetary policy inactions on house prices in Australia? Put differently, how would house prices respond to a monetary policy announcement?

The first half of this paper focuses on the application of a new identification approach proposed by Inoue and Rossi (2019) (hereafter I&R). They define a monetary policy shock as a shift in the entire yield curve on a day of the monetary policy announcement and it is this measurement of a monetary policy shock that allows me to estimate the effects of monetary policy even when the cash rate remains constant. Using a dynamic Nelson and Siegel (1987) framework, I decompose one-day shifts in yields around all policy announcements over 1999-2018 into changes in three time-varying yield curve factors, with each capturing a different dimension of a monetary policy shock. Impulse responses estimated using Jordà's (2005) local projection methodology demonstrate that a positive monetary policy shock leads to a deterioration in nationwide house price growth, albeit limited.

In the second half of the paper, I examine another important issue that is less addressed in the literature: do the effects of monetary policy on house prices vary across cities in Australia? To this end, I model the house price growth rates for eight capital cities in Australia using the monthly hedonic home value indexes. It is found that amongst the three dimensions of monetary policy, a shock in short-term interest rates produces significant heterogeneous effects on city house prices. Furthermore, I investigate if the variation in the number of state building approvals per capita can be a plausible determinant of the city heterogeneity in the effects of monetary policy. I find that incorporating a supply-side influence into my baseline specification produces a much more sizeable decline in house price growth rates across Australia. Nonetheless, the finding that a city with a higher number of building approvals per capita has a stronger house price response to a conventional monetary policy shock, appears to contradict what a standard demand and supply framework would suggest. To this end, I conjecture that this puzzle may be caused by the differences in interest rate demand elasticity.

Additionally, I explore the links between changes in the yields and house price responses arising from a monetary policy shock. Using six representative monetary policy announcements, my results suggest that changes in the short-end of the yield curve may be able to cause a larger impact on house price growth. More research, however, is required to systematically examine the role of the changes in the longer-term interest rates. Overall, using I&R's (2019) identification approach, I find that irrespective of whether there is a change in the policy rate, each monetary policy announcement can produce a distinct effect on house price growth. Importantly, the effect will depend on how each policy decision is perceived by the market participants which is reflected in the yield curve changes. My results show that a cut (rise) in cash rate may not necessarily lead to an increase (a decrease) in house price growth in Australia.

This paper proceeds as follows. Chapter 2 reviews two related strands of literature, namely the effects of monetary policy shock on the housing market and its identification. Chapter 3 describes the data used and some features of the Australian housing market are highlighted in Chapter 4. Chapter 5 then explains the identification framework and the methodology employed. Chapter 6 contains the empirical results. It first presents the estimated monetary policy shocks before considering the preliminary analyses and results for the heterogeneous effects of monetary policy on house prices. Chapter 7 discusses the results established while robustness checks are given in Chapter 8. Chapter 9 concludes.

# Chapter 2
# Literature Review

## 2.1 Monetary Policy and Housing Market

The effects of monetary policy are well documented (for US examples, see Christiano et al., 1996b; Inoue and Rossi, 2019; Sims, 1992); (for an Australian example, see Haug et al., 2005). Nonetheless, relatively less attention has been given to examining the responses of housing prices to monetary policy shocks. The relationship between interest rates and house prices is typically observed through the lens of a user cost model, which in the housing context, represents the annual cost of homeownership (Himmelberg et al., 2005). One of the key components of the user cost is mortgage interest rate. As Elbourne (2008) explains, higher interest rates will translate into a higher cost of borrowing, thereby dampening the demand for housing and house prices. The causality going from short-term rates, through mortgage lending, to house prices is also empirically confirmed (Jordà et al., 2015).

There is ample evidence of the effects of monetary policy on house prices from the US and a number of other developed countries. For example, Jarocinski and Smets (2008) find that a 50 basis point rise in the federal funds rate results in a delayed 1 percent decline in real house prices in the US. On the other hand, Giuliodori (2004) finds that the dynamics and significance of the real house price responses to monetary policy shocks vary across the eight European countries used in his sample, with the largest and most significant declines being observed in the UK, Sweden and Finland. Both studies further examine the important role of house prices in affecting consumption. This relates to another significant branch of the housing market literature, which investigates the housing-related channels of transmission mechanism.[1] However, this paper will solely focus on the first stage of the transmission mechanism, which goes from monetary policy to house prices.

Over the past two decades, house prices in capital cities across Australia have experienced rapid growth. National house prices over the period 2012-2017 alone

[1] See Giuliodori (2004) and Mishkin (2007) for a detailed theoretical discussion of the effects of house price changes in the transmission mechanism. Evidence from Australia can be found in Dvornak and Kohler (2007), whose results show a significant effect of changes in housing wealth on consumption expenditure.

have grown 40.3%, predominately driven by Sydney and Melbourne. As a result, the Australian housing market has received considerable attention from media commentators and researchers. Within the Australian housing market literature, a large body of work is focused on studying the time series behaviour and determinants of house prices (Abelson et al., 2005; Otto, 2007). On the other hand, some papers investigate the issue of overvaluation and the associated risk of a house price bubble (Dvornak and Kohler, 2007; Fry et al., 2010).

There is, however, a small but growing body of literature on studying explicitly the effects of monetary policy on house prices in Australia. Wadud et al. (2012) are amongst the few who empirically analyse the responses of housing market variables to monetary policy in the Australian context. Their results suggest that a contractionary monetary policy shock is followed by significant reduction in housing activity in Australia. While they do not find a significant negative effect on real house prices, they confirm the important role of short-term interest rate for explaining the house prices in Australia. This is consistent with Otto (2007) who shows that mortgage rate is a major determinant in the growth rate of house prices.

My work is also related to the literature on heterogenous effects of monetary policy. While this paper does not directly examine the non-linear relationship between interest rates and house prices, I show how house price changes respond to different monetary policy actions. A recent study by Lim and Tsiaplias (2016) finds that the response of house prices in Australia depends nonlinearly on interest rate levels. They also estimate a threshold of interest rate below which house prices become unstable, thereby offering implications for the setting of mortgage lending rates. Nonetheless, they employ regional house price to income ratios, whose ability in accurately assessing the state of house prices is criticised by Himmelberg et al. (2005).

Another paper that similarly looks at the asymmetric relationship between interest rates and house prices in Australia is Valadkhani et al. (2019), who demonstrate that a rise in interest rates causes a bigger fall in house prices than the increase caused by a cut of an equivalent magnitude. They also illustrate that the effects of changes in mortgage rates on house prices vary across cities, with greater effects observed in Sydney and Melbourne. That said, their work is the most closely related to mine. Similar to Valadkhani et al. (2019), I use monthly city-level house price data in Australia to test for the presence of city heterogeneity in the effects of interest rates. However, my paper is distinct in several aspects. First, I explicitly model the effects of monetary policy via changes in the entire term structure of interest rates, as opposed to their modelling of standard variable mortgage rate. In doing so, I am able to examine how changes in the overnight policy rate affect the longer-term interest rates, which then have broader impacts on house prices. Second, the way I identify a monetary policy shock allows me to estimate the effects of any specific monetary policy announcements, including those of a monetary policy inaction. This is an important feature highlighted in my analysis and is ignored by Valadkhani et al. (2019) as they only allow for two regimes in their model —rising and falling mortgage rates. Third, considering how persistent house prices are, I capture the dynamic responses over a much longer horizon, thus enabling reactions to unfold. Lastly, I attempt to model a possible mechanism underlying the heterogeneous responses of city house prices. This is a limitation of their study as their paper abstracts from a detailed discussion on why they observe different responses in cities.

Compared to the aggregate impact, the heterogenous effects of monetary policy on city house prices have been much less explored for Australia.[2] This is surprising given the considerable differences between regional housing markets. In light of the limited evidence, one of the goals of this paper is to fill this gap in the Australian literature. Specifically, I investigate whether heterogeneity in housing supply elasticities can explain why monetary policy affects house prices differently across cities in Australia.[3]

## 2.2 Identification of Monetary Policy Shock

To estimate the effects of monetary policy on house prices, one needs to first identify the exogenous component of a monetary policy shock. This has always been challenging due to the endogenous response of monetary policy to a variety of factors. For example, in the housing market context, the joint determination between interest rates and house prices has made establishing a causal effect more difficult (Jordà et al., 2015). The central bank might adjust the monetary policy in response to movements in house prices, and hence the change in the policy interest rate is likely to be endogenous. Indeed, the question of which identifying assumptions to be imposed remains a subject of debate among researchers.

In order to overcome this identification challenge, several approaches have been developed over time. Amongst various approaches, vector autoregressions (VARs)

[2] Apart from Valadkhani et al. (2019), Otto (2007) also looks at the differential effects by estimating a separate equation that includes variable mortgage rate as an independent variable for the eight capital cities in Australia.

[3] Aastveit and Anundsen (2018) also examine if the house price responses in the US to a monetary policy shock depend on the elasticities of housing supply.

first introduced by Sims (1980) are no doubt the most widely employed tool in the analysis of monetary policy effects. Despite their widespread use, these empirical models have drawn criticisms, including those from Evans and Kuttner (1998). They criticise that VAR models contain only limited amount of information among other problems, and therefore are skeptical of the technique's accuracy. An alternative strategy has been the narrative approach used by Romer and Romer (1989), who utilise FOMC minutes to identify exogenous monetary policy shocks. Yet, Christiano et al. (1996a) argue that one major deficiency is that the identifying assumption underlying this strategy —that most, if not all, of a large monetary policy action is exogenous —lacks motivation. Christiano et al. (1996a, 1999) make a significant contribution to understanding the monetary policy shock literature by summarising the different approaches taken and implications in analysing monetary policy effects.

Traditionally, a monetary policy shock is identified as an exogenous change in short-term interest rates (Christiano et al., 1999). However, one limitation lies in its inability to estimate the effects of monetary policy when there is no change in short-term interest rates. This is similar to the issue faced by Romer and Romer (1989) approach, which generates only a small number of episodes involving large policy actions and hence shocks. Indeed, the US has experienced a period of zero lower bound where the Federal Reserve's ability to stimulate the economy through cutting interest rates is constrained, in turn prompting them to adopt unconventional monetary measures. In cases like this, zero shift in the policy rates would be considered as zero monetary policy shock according to the traditional literature, thus the narrow focus on short-term interest rates will miss many dimensions of monetary policy shocks. In the face of this problem, Inoue and Rossi (2019) propose a new procedure to identify monetary policy shocks, which they define as shifts in the whole term structure of government bond yields on a day of monetary policy announcement. Measuring the shock as a function, as opposed to a scalar thus allows them to generate many 'episodes', one for each change in the shape of the entire function of yields on a monetary policy announcement day. Using this new approach, they are able to find that both conventional and unconventional easing monetary policies lead to an increase in output and inflation in the US.[4]

As discussed in the Introduction, there have been more instances recently where the policy rate in Australia is kept constant. This undoubtedly presents difficulties for researchers to estimate the effects of monetary policy shocks. Further, the RBA governor Philip Lowe indicated in a recent report that the Bank has considered the

[4] In their model, they divide the data into two sub-samples, one being the conventional monetary policy period (1995:1-2008:10) and another being the zero lower bound period (2008:11-2016:6).

use of unconventional monetary policy tools, reinforcing the possibility of Australia being constrained by the zero lower bound in the future. Thus, this provides a rationale for adopting this novel identification approach proposed by Inoue and Rossi (2019) which can capture the effects of monetary policy during both conventional and unconventional times in a unified framework.

This paper contributes in the following ways. First, to the best of my knowledge, this paper is the first to apply this recently developed shock identification procedure to an Australian dataset. This allows me to reassess the causal relationship running from monetary policy to house prices in Australia, thereby contributing an additional evidence to the literature. Second, using city-level panel data, I explore whether there exists any city heterogeneity in the effects of monetary policy in Australia —an issue that has received less attention. Lastly, I use monthly data that covers city house prices until the end of 2018. Most studies on the Australian housing market only employ data up to late-2000s and data on a quarterly basis. Consequently, they fail to account for the period 2012-2017, over which the most interesting house price dynamics occurred. By employing a much richer and updated dataset, this study can yield additional insights and render implications that are more relevant to the contemporary housing markets in Australia.

# Chapter 3
# Data

This Chapter introduces the data. I first describe the index selected to represent house prices and the advantages associated with this choice. This is then followed by the details regarding the data on yields and monetary policy announcements. The sample period is 01/1999-12/2018, which is the common time frame for the data series.

## 3.1 House Prices

Data on the house prices are sourced from CoreLogic, which covers 98% of the Australian housing market. Specifically, monthly hedonic home value indexes for all dwellings (which include both houses and units) on the eight capital cities in Australia are used to represent house prices. In that respect, I use 'house price' to denote home value index throughout the paper. Instead of using the level of house price, I model the monthly change in house price ($\mathrm{HP_{it}}$), calculated as,

$$\Delta \mathrm{HP_{it}} = (\ln \frac{\mathrm{HP_{it}}}{\mathrm{HP_{it-1}}}) \times 100 \tag{3.1}$$

and hence is expressed in a percentage. After first-differencing, this amounts to 239 observations for each city series. There are two reasons why the city house price series are first differenced. First, as discussed shortly, house prices in Australia display time trends. First differencing the series thus reduces the trends and helps address non-stationarity. Second and more importantly, $\mathrm{HP_{it}}$ is an index which simply serves as a means of summarising the underlying values of properties, and so computing changes can provide a more meaningful economic interpretation than using levels (Himmelberg et al., 2005).

Three important advantages of this dataset are worth noting here. First, the indexes are constructed by a hedonic regression methodology that takes into account the changes in observable property attributes such as the number of bedrooms, land area and suburb. This implies that a lengthy list of controls capturing the characteristics is not necessary. Second, the residential property values are updated on a daily basis and a monthly constant quality index is compiled on the last calendar day of each month, with results released with a one-day delay. Thus, they offer a more timely and accurate measure than the quarterly house price statistics published by the Australian Bureau of Statistics (ABS) with a 2.5 month delay. Lastly, city-level data can be used to construct a panel dataset, thus allowing one to test for any city heterogeneity in the effects of monetary policy that potentially exists.

### 3.2 Australian Yields

I&R's (2019) identification approach relies on the changes in the entire yield curve, it is therefore necessary to estimate the yield curve for Australia before any analyses. Since no zero-coupon bonds are issued in the Australian market, I obtain the daily term structure data from the zero-coupon interest rate analytical series published by the RBA.[1] As will be discussed, the daily frequency is critical for identifying the exogenous monetary policy shocks surrounding every announcement from the Bank. These hypothetical yields are inferred from the observed coupon-bearing Australian Government bonds using a modified Merrill Lynch Exponential Spline methodology and are measured as a percentage.[2] The yield curves are estimated at maturities in quarterly increments out to 10 years, thereby giving a total of 41 maturities.

The inferred yields from 1992 through 2018 are contained in Appendix A. Two features are worth noting. Firstly, as mentioned in the Introduction, there have been several extended periods of a constant cash rate. This is particularly evident in the year 2014 and the period starting from late 2016, during which relatively flat medium- to long-term yields are observed. Secondly, the yields near the end of the sample period were at very low levels by historical standards. In fact, from August 2016 through to December 2018, the overnight cash rate has remained at 1.5%, a level that is very close to the zero lower bound.[3]These features thus reinforce the use of I&R's (2019) identification approach, details of which are provided in the next section.

Using inferred yields instead of actual zero coupon yields may be problematic. Given these yields are estimated, they are subject to estimation error. As Swanson (2007) criticises, different modelling techniques and hence different underlying assumptions would produce different estimates. Since the monetary policy shocks in this paper are constructed based on the inferred yields, any measurement error of the yields will be incorporated into the measure of the shocks, which can result in imprecise

[1] This dataset is available since 1992.

[2] See Finlay and Chambers (2009) for the complete estimation methodology.

[3] In October 2019, the cash rate has been cut to a new record low of 0.75%.

estimates of the house price effects. While the use of inferred yields is not ideal, I believe this is the second-best option given the unavailability of such bonds in the market. Furthermore, this gives me data of a daily frequency to work with, which is important for reasons that will be discussed in the following section. Thus, there is scope for future research to reassess the application of this identification methodology to Australian data when such bonds become available.

## 3.3 Monetary Policy Announcements

Data on monetary policy changes and announcement dates are also sourced from the RBA. Commencing February 2008, any monetary policy change takes effect on the day following the Reserve Bank Board Meeting and therefore the announcement date is obtained from the Statement published on the Bank website. For period prior to December 2007, the announcement dates and effective dates were the same. Among the total of 220 RBA meetings during the sample period,[4] there are 24 announcements of interest rate cut and another 24 of interest rate rise. On the other hand, there are a total of 172 monetary policy inactions (that is, keeping the cash rate unchanged), of which 27 times occur consecutively from October 2016 to December 2018.[5]

## 3.4 Other Variables

Data on some additional variables are used in the models and to support the interpretation of results established in this paper. Their details are provided in Table 3.1.

Table 3.1: Data Description

| **Variable** | **Unit** | **Source** | **Details** |
|---|---|---|---|
| **Standard variable lending rate** | Percentage | RBA series ID: FILRHLBVS | Lending rate on housing loans (Owner-occupier) |
| **Building approvals** | Number | ABS Cat No: 8731.0 | Total number of dwelling units (states) |
| **Average housing loan** | Value | ABS Cat No: 5609.0 | Total value of commitments divided by total number of commitments (states) |
| **Population** | Number | ABS Cat No: 3101.0 | Estimated Resident Population (states) |

[4]No meetings are held in January, except for 1990 and 1992.

[5]As of writing, the period of cash rate on hold is extended to 32 months, making 10/2016-05/2019 the longest period in Australia's history without a change in cash rate.

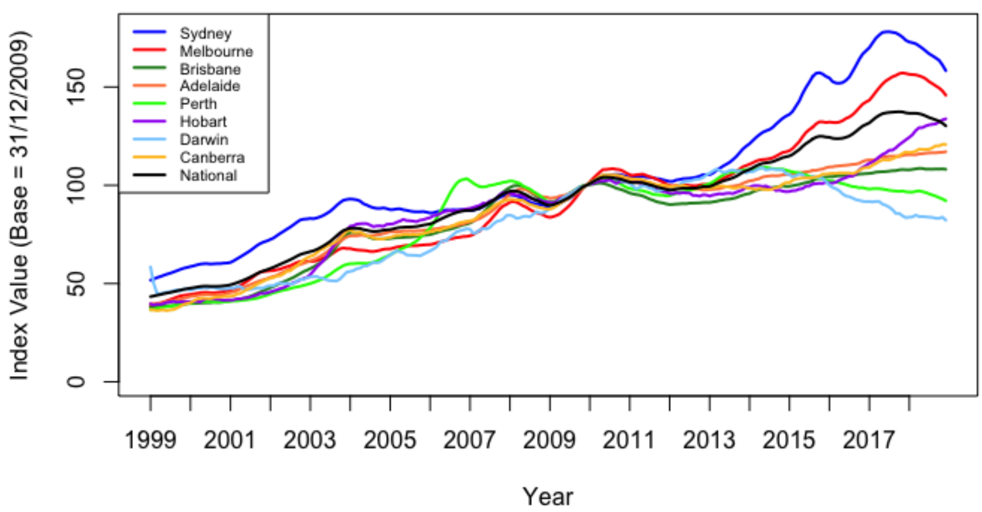
# Chapter 4
# The Australian Housing Market

This Chapter documents some of the key features observed in the Australian housing market, which need to be reflected in the models employed.

## 4.1 Spatial Variation in House Prices

Figure 4.1 shows the monthly hedonic home value indexes for the eight capital cities alongside a national measure from 1999 to 2018. One feature that stands out in the plot is the massive growth that house prices in Australia have experienced, despite occasional busts. Using the national home value index, house prices have risen by 200% over the period being studied. Several factors have contributed to this rapid growth, including demographic changes (e.g. migration), strong economic growth and low interest rates (Yates, 2011). The presence of this upward trend over time reinforces why the house prices need to be log-differenced.

Figure 4.1: Monthly House Prices



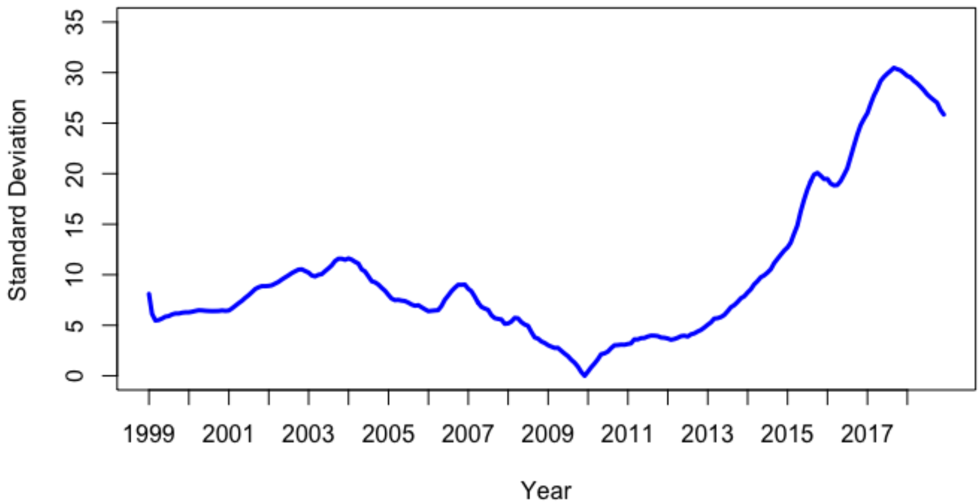

Source: CoreLogic

Another feature that is perhaps not visible in Figure 4.1 yet needs to be captured in the models is that house prices exhibit substantial persistence. It is known that house prices display substantial inertia, meaning that it takes time for prices to respond to new information and adjust to new levels.

Nationwide price increases, however, constitute only one part of the story because of the considerable differences in regional housing markets. One characteristic of the Australian housing market is its high geographical concentration, with a majority of housing demand and supply focusing on the eight major capital cities. In this regard, differences in sub-national housing markets will be analysed at a city level throughout this paper. The substantial spatial variation is evident in Figure 4.1. As is clear, the level of Sydney house prices has consistently been the strongest in the country, except for the period around 2006-2009 when Perth house prices surpassed those in the other cities most likely driven by the mining boom. House prices started to diverge dramatically in around 2012, as evidenced by the much wider gaps between different house price series.

It is also worth noting that house prices in Sydney and Melbourne —two most populous cities in Australia —have experienced much larger rises and falls than the others, in turn driving the national measure. During the period 2012-2017 alone, the national measure of house prices has grown 40.3%. Sydney and Melbourne witnessed a massive 70.4% and 57.1% increase in house prices respectively for the same period.[1]

Figure 4.2: Cross-City Standard Deviation in House Prices



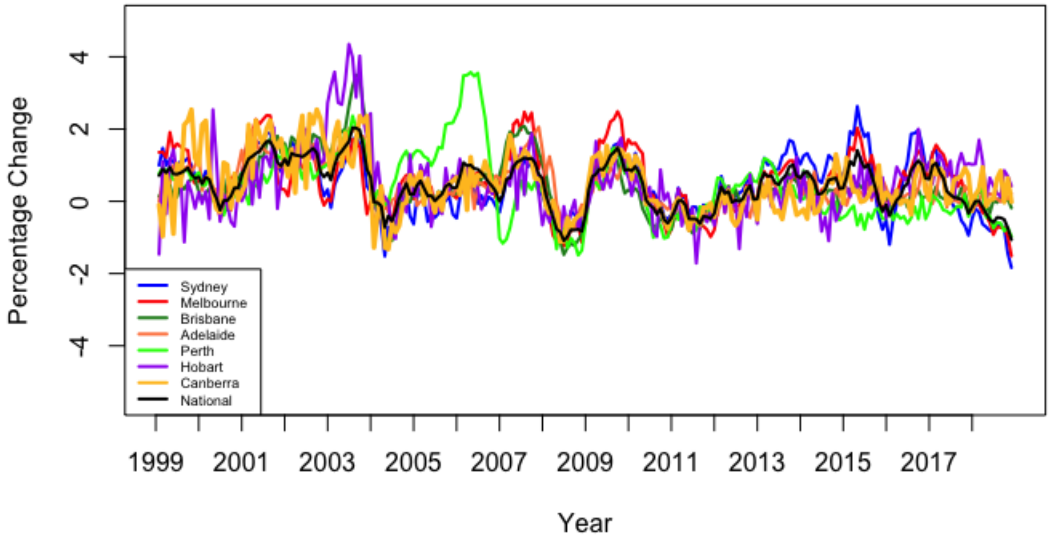

Source: CoreLogic; based on my calculations

This feature of spatial variations is also supported in Figure 4.2 which displays the

[1]Calculations are based on monthly hedonic home value indexes.

cross-city standard deviation in the level of house prices across time. Clearly, there has been a rise in the regional price dispersion in Australia over time. It escalated in the second half of the sample period, with the largest house price differential occurring in late 2017 before declining. This can be explained by the recent sharp falls in Sydney and Melbourne house prices, whereas some other cities such as Hobart and Canberra experienced modest growth in house prices, thereby reducing the price differential. In addition to the level, the growth rate of house prices also varies considerably across cities and across time, as is evident in Figure 4.3.

Figure 4.3: Monthly Changes in House Prices




Notes: Changes in house prices are calculated as the log difference based on monthly hedonic value indexed from CoreLogic. Due to the huge negative changes in Darwin house prices at the start of sample period, the series is excluded for better comparison. The full graph showing house price changes in all capital cities is plotted in Figure B.1.)

Table 4.1 shows the descriptive statistics for the growth rate of house prices for eight capital cities and a national measure, which is what I model in this paper. This table is comparable to the one in Otto (2007). Compared to his estimates, the averages in Table 4.1 are generally lower. This is partly due to the fact that I am working with monthly data whereas Otto (2007) employs quarterly house price indexes. It can also be explained by the different sample period used. Using data over the period 01/1999-12/2018, I am able to capture some of the biggest downturns in Australian house prices that occurred in the past two decades, including the post-GFC trough and the more recent housing market correction. Otto (2007) on the other hand uses the period 03/1986-02/2005 which is characterised by rapid rises in house prices, particularly in Sydney and Brisbane.

Despite using different index series for house prices, the averages in both tables show regional variation in Australia, which is consistent with what Figures 4.1 and 4.2 suggest. Furthermore, with the updated data, it can be seen that city-specific house prices become more correlated with one another. This indicates that housing markets between cities are increasingly linked, so that any strength or weakness in one market tends to feed into other markets. However, there are two exceptions. The first comes from Perth: the start of mining boom in around 2005 has led Perth to experience the much stronger house price growth than all other cities. A peak in around 2014 was subsequently followed by a significant slowdown in mining activities, thereby causing Perth house prices to fall. As other city housing markets were still booming at the time, the falling house prices in Perth have driven down its correlation with all cities over time. The second exception is seen in Darwin, where house price growth has become less correlated with all other cities but Perth. The same statistics computed using the level of house prices are reported in Table B.1.

The above observations on the house price data reveal two important implications. First, it is crucial for the models in this paper to capture these properties of the house price data. For example, a long lag length will be used in the models to capture the high persistence of house prices. Second, the considerable spatial differences in house prices imply that the effects of monetary policy are likely not uniform across cities in Australia as a result of the differential housing demand and supply responses.

Table 4.1: Descriptive Statistics for the Growth Rates of House Prices

| | **Sydney** | **Melbourne** | **Brisbane** | **Adelaide** | **Perth** | **Hobart** | **Darwin** | **Canberra** | **National** |
|---|---|---|---|---|---|---|---|---|---|
| **Mean** | 0.468 | 0.559 | 0.436 | 0.452 | 0.384 | 0.509 | 0.143 | 0.499 | 0.460 |
| **Std** | 0.819 | 0.851 | 0.818 | 0.615 | 0.929 | 0.949 | 1.638 | 0.831 | 0.640 |
| **Max** | 2.628 | 2.485 | 3.495 | 2.080 | 3.571 | 4.348 | 3.214 | 2.536 | 2.039 |
| **Min** | -1.835 | -1.513 | -1.477 | -1.005 | -1.492 | -1.713 | -15.129 | -1.325 | -1.090 |
| | | | | **Contemporaneous Correlations** | | | | | |
| **Sydney** | 1.000 | | | | | | | | |
| **Melbourne** | 0.815 | 1.000 | | | | | | | |
| **Brisbane** | 0.566 | 0.608 | 1.000 | | | | | | |
| **Adelaide** | 0.517 | 0.614 | 0.818 | 1.000 | | | | | |
| **Perth** | 0.239 | 0.331 | 0.478 | 0.397 | 1.000 | | | | |
| **Hobart** | 0.343 | 0.364 | 0.659 | 0.575 | 0.350 | 1.000 | | | |
| **Darwin** | -0.003 | 0.005 | 0.163 | 0.128 | 0.219 | 0.103 | 1.000 | | |
| **Canberra** | 0.486 | 0.559 | 0.696 | 0.647 | 0.364 | 0.476 | 0.205 | 1.000 | |
| **National** | 0.877 | 0.863 | 0.821 | 0.753 | 0.566 | 0.564 | 0.105 | 0.675 | 1.000 |

# Chapter 5
# Methodology

Details on the methodology employed in this paper are provided in this Chapter. Before estimating the responses of house price growth, I first need to estimate the yield curve for Australia in order to identify a monetary policy shock. As with I&R (2019), I apply the Nelson and Siegel (1987) framework to model yields as a function of their maturity. This parametric approach is presented below.

## 5.1 Identification of Monetary Policy Shocks

I start by providing some intuition about the importance of the yield curve. According to the expectation hypothesis, the long-term interest rates are the risk-adjusted expected future short-term interest rates (Piazzesi, 2010). It is now known that monetary policy has the potential to move both the short-term rates via its policy instrument, and the future expectations of short-term rates. Thus, any change in the yield curve can be used as a measure of a change in market expectations of future path of monetary policy (Gürkaynak and Wright, 2012). It is this information contained in the yield curve that I&R (2019) exploit in their identification approach. They measure the surprise component of monetary policy by looking at the effects of announcements on yields. Before turning to the details of their approach, I explain how I model the yield curve for Australia.

To do this, I follow the functional form of Nelson and Siegel (1987). The basic idea of the their model is that the yield curve can be decomposed into three time-varying factors, namely level, slope and curvature components. Hence, it can be estimated as:

$$y_t(\tau) = \beta_{1,t} + \beta_{2,t}(\frac{1-e^{-\lambda\tau}}{\lambda\tau}) + \beta_{3,t}(\frac{1-e^{-\lambda\tau}}{\lambda\tau} - e^{-\lambda\tau}) \tag{5.1}$$

where $y_t(\tau)$ is the yield to maturity at time t; $\beta_{1,t}$, $\beta_{2,t}$ and $\beta_{3,t}$ are parameters capturing the level, slope and curvature of yield curve respectively; $\tau$ is the maturity of the bond expressed in years and $\lambda$ is the tuning parameter controlling the speed of exponential decay. Notice that, in principle, zero-coupon bond yields are required to estimate the Nelson and Siegel (1987) model. However, due to the unavailability of such bonds in Australia, yields inferred from the observed coupon bonds are used instead.

Each of the yield curve factors extracted carries important economic interpretation. From equation (5.1), it can be seen that the level factor ($\beta_{1,t}$) corresponds to the monetary policy effects that cause a simultaneous shift in interest rates at all maturities. Changes in the slope factor ($\beta_{2,t}$) act to rotate the yield curve. This factor captures the effects that affect the short-term interest rates, thus it also represents the conventional measure of a monetary policy shock. The third factor ($\beta_{3,t}$) influences the curvature of the term structure, and it corresponds to the medium-term monetary policy effects.

Linear combinations of the factors are also useful in understanding the yield curve movements. The sum of the level and slope factors, ($\beta_{1,t} + \beta_{2,t}$), gives the instantaneous yield or equivalently, the yield obtained at maturity zero. Whereas the combination ($\beta_{3,t} - \beta_{1,t}$) represents the long-term expectations about the future course of monetary policy which result in a disproportionate shift in yields. According to I&R (2019), this is the additional dimension of monetary policy that cannot be captured by changes in short-term interest rates.[1] By means of a parametric model, I can therefore understand how monetary policy affects the yields at different maturities (via changes in the three yield curve factors), which in turn allows me to disentangle different aspects of monetary policy.

Essentially, for each day in the sample period, one set of $\{\beta_{1,t}, \beta_{2,t}, \beta_{3,t}\}$ is estimated. After extracting the factors and the corresponding yields, I then estimate the monetary policy shocks as changes in the entire term structure on the day of a monetary policy announcement with the equation:

$$\varepsilon_t^f(\tau) \equiv \Delta y_t(\tau) \cdot d_t \tag{5.2}$$

where $\Delta y_t(\tau) \equiv y_t(\tau) - y_{t-1}(\tau)$ is the change in the yield curve as a function of maturity $\tau$ on any day t; $d_t$ is the dummy variable which takes a value of one on a day of monetary policy announcement and zero otherwise. Notice that the observations for January, in which no meetings are held, are also included as part of the shock series and are considered as zero monetary policy shock. Hence, a monetary policy shock is a type of what I&R (2019) call, a 'functional shock', in the sense that it measures the shock as a function of maturity. Specifically, it is the difference between the yield curve on the day before a monetary policy announcement and that at the end of the announcement day. Note that equation (5.2) can be equivalently written

[1] See I&R (2019) for more details on the interpretation of factors.

as:

$$\varepsilon_t^f(\tau) = \Delta\tilde{\beta}_{1,t} + \Delta\tilde{\beta}_{2,t}\left(\frac{1-e^{-\lambda\tau}}{\lambda\tau}\right) + \Delta\tilde{\beta}_{3,t}\left(\frac{1-e^{-\lambda\tau}}{\lambda\tau} - e^{-\lambda\tau}\right) \tag{5.3}$$

where $\Delta\tilde{\beta}_{j,t} \equiv \Delta\beta_{j,t} \cdot d_t$.

As with I&R (2019), the following identification conditions are assumed to ensure the monetary policy shocks are exogenous:

1. **Shock identification condition:** House prices are not contemporaneously affected by yield curve shocks.
2. **Relevance condition:** A change in the yield curve on an announcement day is only caused by the monetary policy shock.
3. **Exogeneity condition:** The change in yield curve after an announcement day in the sampling period is not due to the monetary policy shock.

Assumption (1) is also commonly known as the recursiveness assumption in the literature, coined by Christiano et al. (1999). In other words, it means that house prices are not allowed to respond to changes in the yield curve within the period. This should be easy to satisfy as house prices are relatively slow-moving: it typically takes time to purchase a house, and so prices are unlikely to move at the same time as interest rates.

Assumption (2) is a relatively strong assumption, in that it allows domestic monetary policy to be the only (or at least the major) driver of any shifts in yield curve on an announcement day.[2] One might be concerned that movements in yields are driven by other domestic or international influences such as economic data releases, apart from the monetary policy announcement in Australia. Here the use of a one-day window around a monetary policy announcement is critical, as it is fairly unlikely that other potential market moving data are released within the same window. One of the macroeconomic news to which the yields are sensitive is the Australian Consumer Price Index (CPI), for which data are consistently released towards the end of month every quarter. This obviously will not affect contemporaneously the monetary policy decision which is always announced at the beginning of the month. Therefore, this allows me to at least rule out the possibility of CPI contaminating my measure of monetary policy shocks. To the extent to which both release dates of major macroeconomic data and monetary policy announcement coincide, the number of times that this occurs should not be large enough to cause any significant

[2] In other words, it is assumed that any other shocks on an announcement day have only minor, if not zero, effects on the yields.

contamination. This is plausible considering the general low frequency of data release in Australia and the relatively large number of observations in my sample. The validity of Assumption (2) is also supported by Gurkaynak et al. (2004), who show that a one-day window is sufficiently narrow to identify a monetary policy shock. Nonetheless, I acknowledge that the narrower the window is, the less likely the measurement of shock will be contaminated by other events and the more plausible it will be.

## 5.2 Basic Model for House Price Growth

The baseline specification in evaluating the effects of monetary policy shocks on house prices is as follows:

$$\Delta HP_{it} = \Gamma_0 + \Gamma_1(L)\Delta\tilde{\beta}_{1,t} + \Gamma_2(L)\Delta\tilde{\beta}_{2,t} + \Gamma_3(L)\Delta\tilde{\beta}_{3,t} + \gamma(L)\Delta HP_{it-1} + \delta_i + \lambda_t + u_{it} \tag{5.4}$$

The dependent variable —the monthly percentage change in house prices in each city i ($\Delta HP_{it}$) —is regressed on the current and lagged measures of each of the three dimensions of a monetary policy shock $\{\Delta\tilde{\beta}_{1,t}, \Delta\tilde{\beta}_{2,t}, \Delta\tilde{\beta}_{3,t}\}$ as well as lagged house price changes. $\Gamma_j(L)$ for $j = 1, 2, 3$ and $\gamma(L)$ are both polynomials of order 24.[3] This choice of lag length is determined by ensuring that the sum of coefficients on lagged house price changes is below 1, as any values above 1 indicate a non-stationary system and hence unstable dynamics.[4] However, it comes at a cost of generating a long lag length and thus a large number of regressors.[5] $\delta_i$ captures the city-specific trends that are invariant over time, whereas $\lambda_t$ captures the year fixed effects that are common to all cities. Finally, $u_{it}$ is the error term for city i in month t. The specification in equation (5.4) assumes that the effects of each beta shock on house price changes are constant across all cities. I will extend this model in Section 6.2.4 to test for any city heterogeneity in the effects of monetary policy shocks.

Recall that $\Delta\beta_{2,t}$ represents the monetary policy effects on short-term interest rates, it is therefore expected that the sum of $\Gamma_2(L) < 0$, meaning that ceteris paribus, a rise in short-term interest rates results in a fall in city house price growth in the long run. This can be explained by the close link between the cash rate and variable mortgage rate: a rise in policy rate translates into a similar rise in mortgage rate

[3] $\Gamma_j(L) = \sum_{s=0}^{p} \Gamma_{j,s} L^s$ and $\gamma(L) = \sum_{s=0}^{p} \gamma_s L^s$, where L is the lag operator.

[4] In principle, the number of lags on $\{\Delta\tilde{\beta}_{1,t}, \Delta\tilde{\beta}_{2,t}, \Delta\tilde{\beta}_{3,t}\}$ can be different to that on house price changes. However, I decide to use the same lag length and assume that a lag of 24 months is sufficient for any monetary policy effects on house prices to unfold.

[5] As discussed later, this becomes problematic when I use these long lags as instruments to correct for a possible bias caused by the dynamic panel structure.

and hence a higher cost of borrowing. This in turn dampens the demand for housing and house prices.

The inclusion of lagged house price changes as regressors helps control for any serial correlation. Moreover, the lag length attached to the variable $\Delta \mathrm{HP}_{\mathrm{it-1}}$ reflects the feature of persistence in house price data. Unlike other asset prices, house prices are a slow-moving variable, and so they are expected to adjust, if any, with a delay to changes in interest rates. This process of adjustment is represented by the current and lagged beta shocks on the right hand side of the equation.

In the previous subsection, I outline reasons why the first step of identifying an exogenous monetary policy shock is plausible. Now, the second step of identifying the responses of house prices also merits a discussion given the one-month gap between a monetary policy announcement and house price data release. While it is conceivable that other factors would have driven house price growth rates during that time interval, I have decided not to include these variables as controls in equation (5.4).

In terms of the effects of macroeconomic data releases, this should be less of a concern. To illustrate why this is true, I consider two cases. The first case is when data releases occur after the house price data release (end of month) but before the monetary policy announcement (in the following month). In this regard, any surprises in macroeconomic news are expected to feed back into monetary policy. Put differently, any new information should be embedded in the next or future cash rate movements. For instance, a positive surprise in Australian CPI will likely translate into an increase in the cash rate. Therefore, the data on monetary policy announcements should have accounted for the effects of macroeconomic news in the first place. The second case concerns the data releases that occur between a monetary policy announcement and house price data release in a given month. House prices are known to display substantial inertia, so they are sluggish in responding to any new information. It is unlikely that the unexpected effects of a data release (if any) will get reflected immediately in the house prices in the same month. On the contrary, the high serial correlation in house prices means any effects may be reflected in a month or so. The long lags on $\Delta \mathrm{HP}_{\mathrm{it}}$ should help me control for the slow response of house prices. With regards to the longer-run determinants of house prices such as population, some of these effects would be picked up by the city and time fixed effects in equation (5.4).

The above discussion is not meant to imply that other factors have no influences on house price growth. Rather, they are excluded to keep my model more parsimonious. It should also be emphasised that the present house price model is limited to capturing one channel through which house prices are affected —the effects of monetary policy shocks on house prices. Ample empirical evidence suggests that a monetary policy shock is often the most important shock to house prices (Elbourne, 2008; Giuliodori, 2004; Iacoviello, 2000). Indeed, I do not expect the inclusion of controls in the model would materially alter the results, especially economic data releases. This is because, unlike the fast-moving financial variables such as exchange rates, house prices are generally less sensitive to macroeconomic news. In sum, the nature of my data and the model design should allay some of the concerns relating to omitted variable bias.

I follow Jordà (2005) and construct impulse responses of city house price growth to a monetary policy shock using local projections. Details on this methodology will be deferred to Section 6.2.2.

# Chapter 6
# Empirical Results

This Chapter presents the main findings of this paper. Section 6.1 first illustrates the identified monetary policy shocks. Some preliminary analyses are then shown in Section 6.2 before turning to the results for heterogenous house price responses in Section 6.3.

## 6.1 Estimated Monetary Policy Shocks

Table 6.1 reports the summary statistics for the actual cash rate changes and estimated shocks $\{\Delta\tilde{\beta}_{1,t},\ \Delta\tilde{\beta}_{2,t},\ \Delta\tilde{\beta}_{3,t}\}$ over the sample period 01/1999-12/2018. Note that the mean of cash rate changes is very close to zero. This reflects both the large number of decisions of keeping the cash rate on hold, as well as the inclusion of all January observations which record a zero monetary policy shock due to the absence of meetings. To put these values in perspective, I normalise the average beta shocks such that a 25 basis point increase in cash rate is, on average, associated with 0.024%, -0.090% and 0.346% change in $\beta_{1,t}$, $\beta_{2,t}$ and $\beta_{3,t}$ respectively. The negative sign on $\beta_{2,t}$ suggests that on average, the RBA raises cash rate less than the market has anticipated.

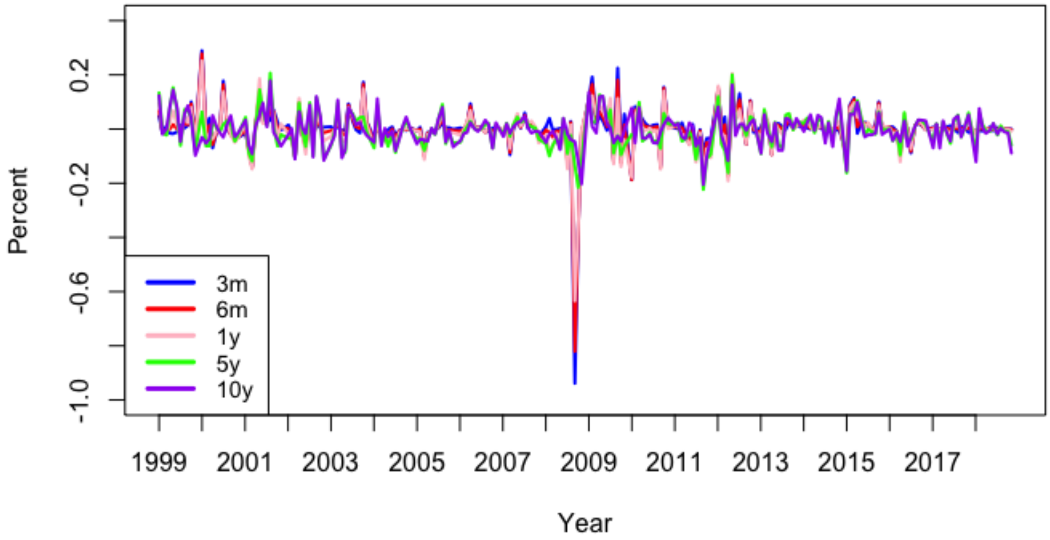
Table 6.1: Summary Statistics for Actual Cash Rate Changes and Estimated Beta Shocks

| | **Change in Cash Rate** | $\Delta\tilde{\beta}_{1,t}$ | $\Delta\tilde{\beta}_{2,t}$ | $\Delta\tilde{\beta}_{3,t}$ |
|---|---|---|---|---|
| **Mean** | -0.01 | -0.001 | 0.005 | -0.019 |
| **Std** | +0.18 | 0.062 | 0.114 | 0.190 |
| **Max** | +0.50 | 0.192 | 0.436 | 0.762 |
| **Min** | -1.00 | -0.305 | -1.110 | -0.800 |
| **Normalised** | +0.25 | 0.024 | -0.090 | 0.346 |

Figure 6.1 plots the identified shocks on some selected maturities over the sample period. The largest surprises are observed during the GFC, suggesting that most of the shocks were largely unexpected. It is unsurprising to see that for shorter-term maturities, the magnitude of the shock is larger. This is intuitive, given that short-

term maturities are typically more sensitive to monetary policy.

Figure 6.1: Identified Monetary Policy Shocks Over Time

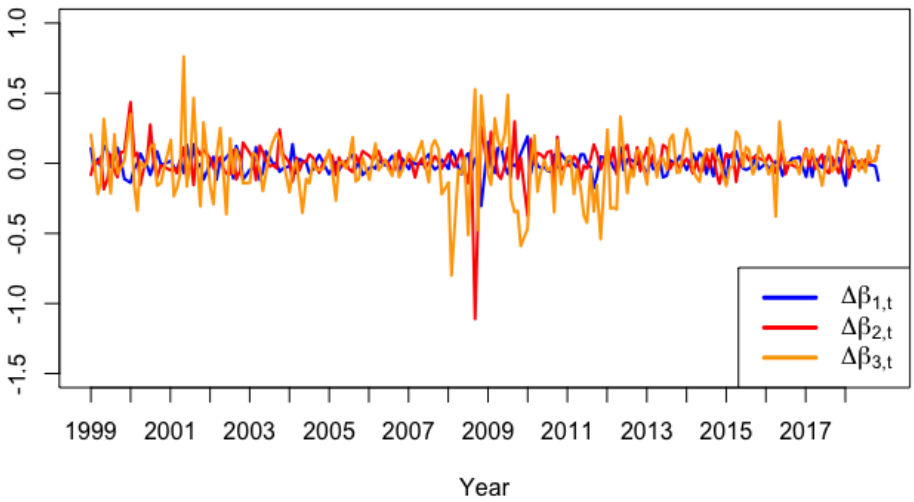



Note: This figure plots the monetary policy shocks identified according to I&R's (2019) approach, expressed in terms of changes in yields at different maturities.

As explained, the shocks can also be understood based on their three components. In panels (a) and (b) of Figure 6.2, I plot the factors individually and jointly as combinations over time. Similar to Figure 6.1, it can be seen that monetary policy shocks have the greatest impacts on all components during the GFC. A more striking feature that emerges from the two figures is that the longer-term components, that is, $\Delta\beta_{3,t}$ and $\Delta(\beta_{3,t}-\beta_{1,t})$ in general have larger realisations over the sample period (except during the GFC). This suggests that simply using a single factor $\Delta\beta_{2,t}$, which proxies the changes in short-term interest rates as conventionally measured, might be inadequate to describe the effects of monetary policy and that additional dimensions are required. Another implication is that monetary policy has not yet lost its effectiveness even in an environment characterised by low and constant interest rates. This is evidenced by the reasonably large magnitude of shocks that involve $\Delta\beta_{3,t}$ towards the end of the sample period, reflecting the Central Bank's ability in affecting the longer-term interest rates.

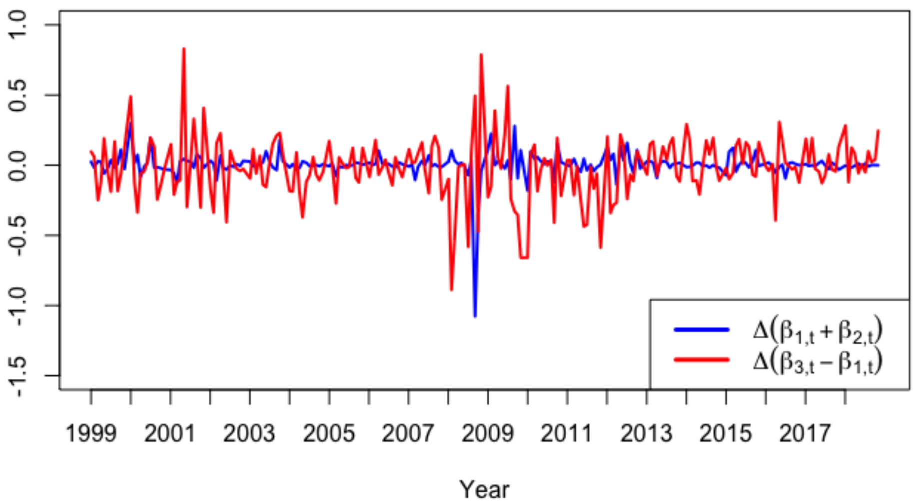
Figure 6.2: Identified Monetary Policy Shocks Over Time

(a) Shock Components



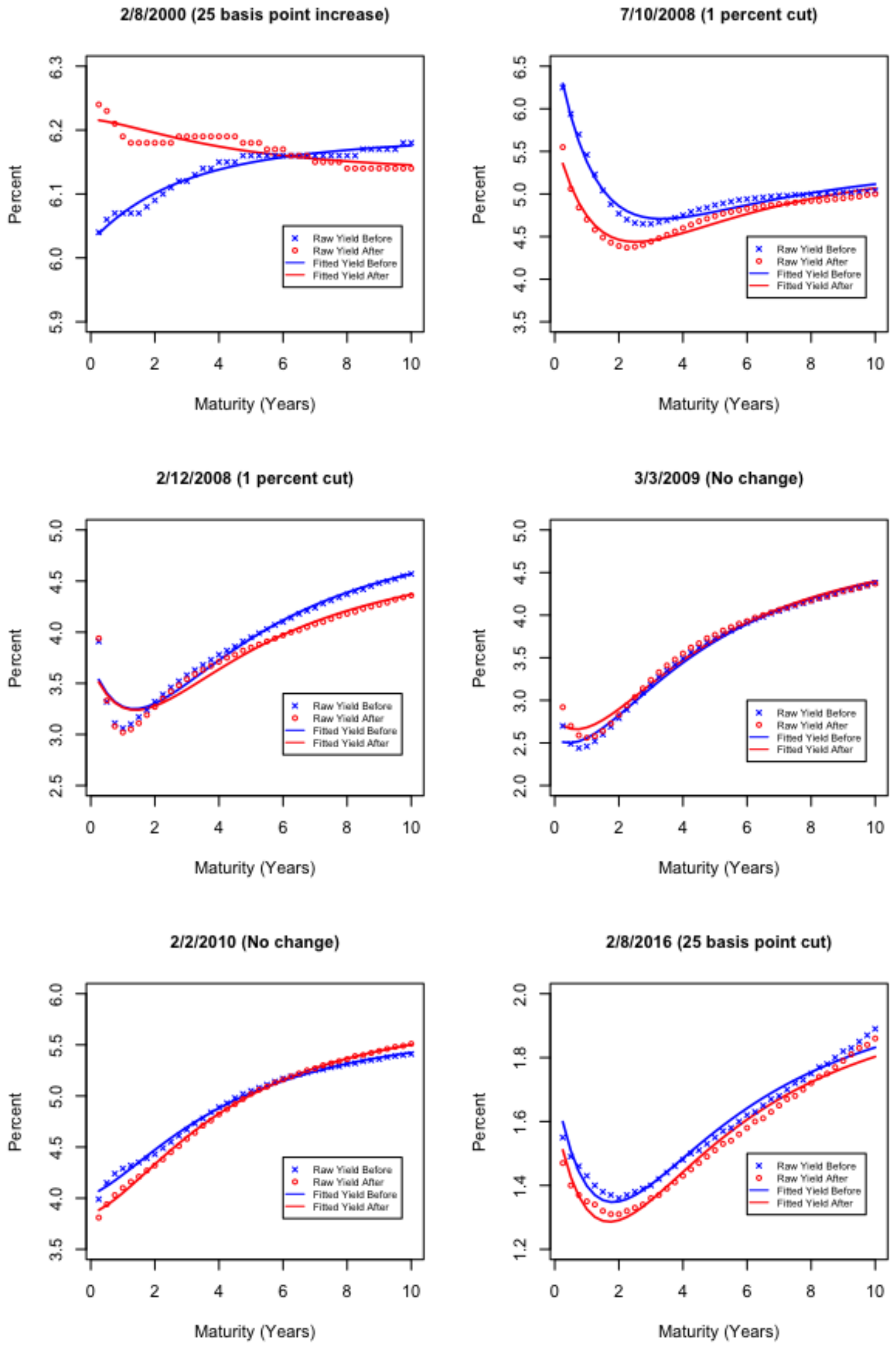

Note: This panel plots the monetary policy shocks identified according to I&R's (2019) approach, expressed in terms of changes in three shock components, i.e. yield curve factors derived from the Nelson and Siegel (1987) model.

(b) Combinations of Shock Components




Note: This panel plots the two different combinations of the shock components, where $\Delta(\beta_{1,t}+\beta_{2,t})$ represents changes in instantaneous yields and $\Delta(\beta_{3,t}-\beta_{1,t})$ captures changes in long-term expectations of future monetary policy.

Figure 6.3: Episodes of Monetary Policy Shocks




Notes: The figure presents six examples of changes in yields after a monetary policy announcement. A monetary policy shock is measured as the difference between the solid red line and the solid blue line. The maturities considered here start from 3 months and increase quarterly up to 10 years, thus giving a total of 40 maturities.

To further motivate and illustrate this shock identification procedure, Figure 6.3 provides six representative examples, with each panel displaying a shift in the yield curve after a monetary policy announcement. The circles and crosses indicate the raw zero-coupon yields at different maturities and the lines indicate the yields estimated parametrically. A monetary policy shock —in other words, the surprise component of a monetary policy action or inaction —is thus identified as the difference between the red line and blue line. An immediate observation from Figure 6.3 is that the Nelson and Siegel (1987) model fits the Australian yield curve very well, which lends weight to the use of parametric yields.

Consider the shock on 7/10/2008 when the RBA announced a 1 percent cut in overnight cash rate (top-right panel of Figure 6.3), the short-term maturities decreased while the longer end barely changed. Now compare this with the shock on 2/12/2008, a cut with the exact same magnitude led to a fall in the longer maturities and left the short-term ones unchanged. This demonstrates that the exact same monetary policy action (same direction and same magnitude of cash rate movement) can lead to a considerably different monetary policy shock.

Shocks can also result in a completely different shape, for example a tilt of the entire function (or a yield curve inversion) on 2/8/2000 versus a parallel downward shift in the curve on 2/8/2016. Furthermore, even when the short-term rates are unmoved, a shock can now be identified by this procedure as long as there is a change in the shape of the yield curve on any announcement day. This is evident in the two shocks on 3/3/2009 and 2/2/2010, which could possibly be ignored if one were to follow the traditional approach of identifying a monetary policy shock.

The shifts in yields on the selected six announcement days can also be summarised by the three yield curve factors and their linear combinations, whose values are reported in Table 6.2. For example, it can be observed that the shock to the slope factor ($\Delta\beta_{2,t}$) on 7/10/2008 is quite significant compared to that on the other announcement days presented. This decrease in $\Delta\beta_{2,t}$ has resulted in a large and negative change in $\Delta(\beta_{1,t}+\beta_{2,t})$, which captures the changes in instantaneous yield. These changes thus correspond to a downward shift in yields on the short end of the yield curve, as evident in the top-right panel of Figure 6.3. On the other hand, the decreases in the medium- and long-term yields on 2/12/2008, shown in the middle-left panel of Figure 6.3, can be explained by the large change in $\Delta(\beta_{3,t}-\beta_{1,t})$.

Table 6.2: Components of Monetary Policy Shocks in Selected Episodes

| **Date** | **Changes in policy rate** | **Summary Statistics** | | | | |
|---|---|---|---|---|---|---|
| | | $\Delta\beta_{1,t}$ | $\Delta\beta_{2,t}$ | $\Delta\beta_{3,t}$ | $\Delta(\beta_{1,t}+\beta_{2,t})$ | $\Delta(\beta_{3,t}-\beta_{1,t})$ |
| 2/8/00 | +0.25 | -0.0836 | 0.2757 | 0.1132 | 0.1921 | 0.1967 |
| 7/10/08 | -1.00 | 0.0331 | -1.1099 | 0.5274 | -1.0768 | 0.4943 |
| 2/12/08 | -1.00 | -0.3047 | 0.2633 | 0.4832 | -0.0414 | 0.7879 |
| 3/3/09 | 0.00 | -0.0006 | 0.2242 | -0.1529 | 0.2236 | -0.1523 |
| 2/2/10 | 0.00 | 0.1922 | -0.3732 | -0.4657 | -0.1811 | -0.6579 |
| 2/8/16 | -0.25 | -0.0182 | -0.0767 | -0.0008 | -0.0949 | 0.0174 |

Notes: This table describes the monetary policy shocks in six different episodes. For each announcement, I report the actual change in the policy cash rate, the estimated values of the shocks to the three yield curve factors and their linear combinations.

These examples illustrate two main points. First, this identification procedure can not only measure the unexpected changes in monetary policy, but also potentially capture other important information that may have been missed by traditional identification methods. In addition to the changes in short-term interest rates, this procedure allows one to understand about how a monetary policy shock impacts people's longer-term expectations of future policy path.

Second, each monetary policy shock ($\varepsilon_t^f(.)$), determined by the level ($\Delta\beta_{1,t}$), slope ($\Delta\beta_{2,t}$) and curvature ($\Delta\beta_{3,t}$) of the term structure, can be different. This means that a shock associated with a monetary policy inaction can also be identified, given that there is a change in shape of the yield curve after an announcement, which is summarised by a combination of changes in the three factors derived parametrically. Thus, only when the yields at *all* maturities remain unchanged after a monetary policy announcement will this approach identify zero monetary policy shock.

## 6.2 Preliminary Analyses

In this section, I conduct some preliminary analyses. For brevity, I only report two important statistics that describe the dynamic behaviour of the house price responses in the following sections. The first being the sum of coefficients on the current and lagged $\Delta\beta_{j,t}$ where $j = 1, 2, 3$ and the t-statistics associated with the test that such a sum is equal to zero. The second is the corresponding long-run effect on the change in house prices, which is calculated by dividing the sum of coefficients on current and lagged values of the beta shock by one minus the sum of coefficients on the lagged changes in house prices, or mathematically,

$$\frac{\sum_{p=0} \Gamma_{j,p}}{(1 - \sum_{p=1} \gamma_p)}$$

The t-statistics corresponding to the test of null hypothesis that such a long-run effect equals to zero, is also reported.

### 6.2.1 Alternative Estimators

Results for the baseline specification with city and year fixed effects (FEs) as shown in equation (5.4) are presented in Column (1) of Table 6.3.[1] As expected, this preferred FE model produces a negative sum of $\Delta\beta_{2,t}$ coefficients and hence a negative long-run effect, suggesting that an increase in short-term interest rates reduces the growth of house prices across Australia.

Nonetheless, as is well known, the lagged dependent variables are typically correlated with the unobserved fixed effects in a dynamic panel model such as the one in equation (5.4). This induces an endogeneity problem, thereby producing inconsistent and biased fixed effects estimators (Nickell, 1981). The resulting bias is known as dynamic panel bias or Nickell (1981) bias. If the bias is present, it has the potential to produce some misleading results. In order to correct for the potential bias, I apply a routine proposed by Arellano and Bond (1991) (hereafter AB) to the preferred FE model in equation (5.4). AB routine involves using lagged levels of dependent variables as instruments after a first-differencing transformation to remove the fixed effects, and hence the name 'difference generalised method of moments (GMM) estimator'.

After some experimentation,[2] the preferred AB specification is a one-step difference GMM estimator with one lag of dependent variable used as instruments and robust standard errors, for which major statistics are summarised in Column (2) of Table 6.3. While some estimates are highly significant, several concerns arise from the implementation of this bias-correction procedure. The difference GMM estimator is designed for micro panel, that is, panel with large cross sectional dimension N and small time dimension T. However, the house price data used in this paper is characterised by $T = 239$ and $N = 8$. With a relatively large T and a long lag length attached to each independent variable, it is not surprising to see that the estimator generates 307 instruments. This problem of instrument proliferation will lead to

[1] A discussion of alternative FE specifications is provided in Appendix C.
[2] See Appendix D for more details.

Table 6.3: Alternative Estimators[a]

| | (1) FE: City, Year | | | (2) AB Difference GMM[b] | | | (3) Pooled OLS | | |
|---|---|---|---|---|---|---|---|---|---|
| | $\Delta\beta_{1,t}$ | $\Delta\beta_{2,t}$ | $\Delta\beta_{3,t}$ | $\Delta\beta_{1,t}$ | $\Delta\beta_{2,t}$ | $\Delta\beta_{3,t}$ | $\Delta\beta_{1,t}$ | $\Delta\beta_{2,t}$ | $\Delta\beta_{3,t}$ |
| Sum[c] | 1.4622 | -0.7707 | -1.3396 | -19.4908 | -3.206 | 7.952 | 1.5431 | -0.788 | -1.3808 |
| | (0.20) | ( -0.22) | ( -1.22) | ( -1.77)* | ( -1.05) | ( 3.50)*** | (0.23) | ( -0.28) | ( -1.31 ) |
| Long-run Effect[d] | 7.2127 | -3.802 | -6.6083 | -9.1043 | -1.4976 | 3.7144 | 8.0697 | -4.1207 | -7.2209 |
| | (0.04) | (0.05) | (1.37) | (2.70) | (0.92) | (8.60)*** | (0.05) | (0.08) | (1.42) |
| Observations | 1720 | | | 1720 | | | 1712 | | |
| Instruments | - | | | 307 | | | - | | |
| Areallano-Bond test for AR(2)[e] | - | | | 0.5554 | | | - | | |
| R-sq | 0.7355[f] | | | - | | | 0.7375 | | |

Notes:
a) All three estimations include year dummies and employ robust standard errors. 24-period lags are used for each beta and change in house prices.
b) This is a one-step Arellano-Bond estimator with one-period lag of dependent variable used as instrument.
c) Test of the null hypothesis that the sum of coefficients is zero, with t statistics displayed in parentheses.
d) Chi-squared test of the null hypothesis that the long-run effect is zero, with chi-square statistics displayed in parentheses.
e) p-value associated with the test of null hypothesis that the second-order autocorrelation of the first-differenced error term is equal to zero.
f) R-sq (within)
* significant at 10%, ** significant at 5%, *** significant at 1%.

overfitting bias and hence imprecise estimates (Roodman, 2009). Furthermore, a common yet arbitrary rule of thumb suggests that the number of instruments should not exceed the number of panel units, which is obviously violated even in this preferred specification.

In fact, the dynamic panel bias diminishes as T increases. According to Nickell (1981), for a reasonably large T, the approximation of the bias is given by,

$$\lim_{N \to \infty} (\hat{\rho} - \rho) \simeq \frac{-(1+\rho)}{T-1}$$

where $\rho$ is the sum of coefficients on lagged house prices in the preferred FE model in equation (5.4). Substituting $\rho = 0.7973$ and $T = 239$ gives a bias of approximately -0.007552. Therefore, with this reasonably large value of T, the dynamic panel bias is quite insignificant.

Another useful way to check the validity of the results produced by the bias-corrected procedure is that, good coefficient estimates should lie within the interval bounded by estimates based on pooled ordinary least squares (OLS) and FE regressions. This is because they produce bias in opposite directions, with pooled OLS estimator biased upwards and FE estimator downwards (Bond, 2002). Therefore, I also run a regression by OLS according to equation (5.4) but less the city FE. Results are reported in Column (3) of Table 6.3. An immediate observation is that the difference GMM estimates obviously fall outside of the range bounded by those of pooled OLS and FE regressions. In fact, the results generated by pooled OLS and FE regressions are qualitatively and quantitatively similar. This suggests that the GMM estimates are very imprecise and that a bias correction procedure is not necessary in this context.

Turning to the comparison between the pooled OLS and FE estimators. While the FE estimator produces a slightly lower R-squared value, it enables one to test for any cross sectional and temporal effects, which are ignored by pooled OLS estimator as it assumes constant model parameters across panel units. Furthermore, the result from the Hausman test shows a failure to reject the null hypothesis of no systematic difference between the pooled OLS and FE estimates, and suggests the FE model fits the data better.[3] Therefore, the FE model for which results are reported in Column (1), is ultimately chosen to be the baseline specification.

[3]Default standard errors are used since the Hausman test does not allow non-default standard errors.

### 6.2.2 Impulse Responses for Baseline Specification

Impulse response functions (IRFs) are commonly constructed by iterating forward the parameters estimated using VAR procedures. Jordà (2005) offers a conceptually equivalent alternative, namely the local projection (LP) methodology, which may be preferred over a standard VAR procedure for a number of reasons: First, LP is robust to misspecification of the data generating process, especially a misspecification of lag length (Brugnolini, 2018; Jordà, 2005). In cases where the model is incorrectly specified, VAR will result in a bias which is exacerbated as horizon increases, thereby producing inaccurate and misleading IRFs. By contrast, using LP can help avoid this type of errors and thus is empirically appealing. Second, it is a better alternative to VAR for panel data, where a large number of dimensions are involved.[4]

Hence, dynamic IRFs are obtained by adapting the LP methodology to the baseline specification in equation (5.4):

$$\begin{aligned}\Delta \mathrm{HP}_{it+h} = & \Gamma_{0,h} + \Gamma_{1,h}(L)\Delta\tilde{\beta}_{1,t} + \Gamma_{2,h}(L)\Delta\tilde{\beta}_{2,t} + \Gamma_{3,h}(L)\Delta\tilde{\beta}_{3,t} + \gamma(L)\Delta \mathrm{HP}_{it-1} \\ & + \delta_i + \lambda_t + u_{it+h} \quad \text{for} \quad h = 0, 1, 2, ..., H,\end{aligned} \tag{6.1}$$

where $h = 0, 1, 2, ..., H$ is the horizon of response. I set $H = 24$ which is equivalent to a response horizon of 2 years. Once again, a lag of 24 months is used for each regressor. According to the LP method, the IRF is basically a sequence of coefficients $\Gamma_{j,h}$ estimated in a series of regressions of $\Delta \mathrm{HP}_{it}$ shifted forward for each horizon h, such that $\Gamma_{j,h}$ represent the responses at time $t + h$ to a shock in $\beta_{j,t}$ at time t, $j = 1, 2, 3$.

Panel (a) of Figure 6.4 displays the accumulated responses of house price growth to a one unit positive shock in each of $\Delta\beta_{1,t}$, $\Delta\beta_{2,t}$, $\Delta\beta_{3,t}$ with the 90% confidence intervals. A positive shock in the three different dimensions of a monetary shock all yields a negative response of house price growth. Given the importance of conventional monetary policy in Australia, it is unsurprising to see that a shock in $\Delta\beta_{2,t}$ produces the most pronounced effect: a one unit increase in the short-term interest rates results in a maximum effect of -6.19% on house price growth in 10 months after the shock. Furthermore, unlike $\Delta\beta_{1,t}$, the response to a shock in $\Delta\beta_{2,t}$ for the first 10 horizons is very precisely estimated, which again reinforces the important role of conventional monetary policy tools in influencing house prices in Australia.

[4] A third advantage is that LP is much easier to implement in the presence of nonlinearities, which present difficulties to VAR models due to the complicated dynamics. However, estimating non-linear effects is not the focus of this study.

Recall that a monetary policy shock ($\varepsilon_t^f(.)$) identified based on I&R's (2019) approach is characterised by $\{\Delta\tilde{\beta}_{1,t}, \Delta\tilde{\beta}_{2,t}, \Delta\tilde{\beta}_{3,t}\}$. That is, the overall response of city house price growth ($\Delta HP_{it+h}$) to a shock ($\varepsilon_t^f(.)$) should be the combined effects of changes in all three yield curve factors after a monetary policy announcement:

$$\begin{aligned} \frac{\partial \Delta HP_{it+h}}{\partial \varepsilon_t^f(.)} &= \sum_j \frac{\partial \Delta HP_{it+h}}{\partial \Delta\tilde{\beta}'_{j,t}} \cdot (\Delta\beta_{j,t} \cdot d_t) \\ &= \Gamma_{1,h}(\Delta\beta_{1,t} \cdot d_t) + \Gamma_{2,h}(\Delta\beta_{2,t} \cdot d_t) + \Gamma_{3,h}(\Delta\beta_{3,t} \cdot d_t) \end{aligned} \tag{6.2}$$

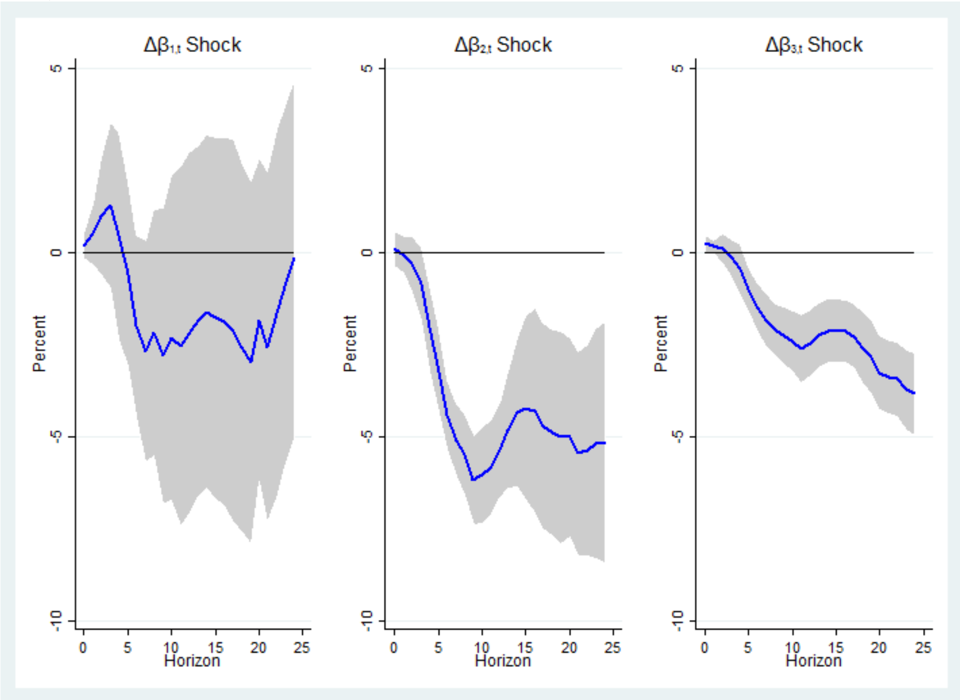
where $\frac{\partial \Delta HP_{it+h}}{\partial \Delta\tilde{\beta}'_t}$ is simply the estimated coefficients obtained from equation (6.1), and $\Delta\beta_{j,t} \cdot d_t$ is the change in the yield curve multiplied by a dummy variable equal to one if there is a monetary policy announcement at time t. It is clear from equation (6.2) that each response can be decomposed into three separate components, representing the effects of a different dimension of a monetary policy shock respectively.

To get a sense of the general effects of an identified monetary policy shock estimated from the baseline specification, I compute the response of house price growth to the average $\Delta\beta_{j,t}$ shocks that are associated with a normalised 25 basis point rise in cash rate, as shown earlier in Table 6.1. More specifically, I feed $\Delta\beta_{1,t} = 0.024$, $\Delta\beta_{2,t} = -0.090$ and $\Delta\beta_{3,t} = 0.346$ into equation (6.2), in turn generating the response shown in Panel (b) of Figure 6.4.

Several features are worth mentioning. Firstly, a positive monetary policy shock leads to a deterioration in house price growth, which is consistent with empirical evidence (Giuliodori, 2004; Jarocinski and Smets, 2008). Secondly, this negative response is somewhat delayed —the growth rate only starts to decline in about 5 months after the shock, reflecting the substantial inertia exhibited in house prices. Lastly, it can be observed that the overall negative impact is quite small, with an accumulated response of about 1% realised in the end of the 24-month horizon. This is because $\Delta\beta_{j,t}$ shocks associated with a 25 basis point rise in cash rate are averaged over the whole sample and are therefore quite small in size. As illustrated in Section 6.3.3, a much larger house price response is generated when I feed in the actual $\Delta\beta_{j,t}$ shocks arising from historical monetary policy announcements.

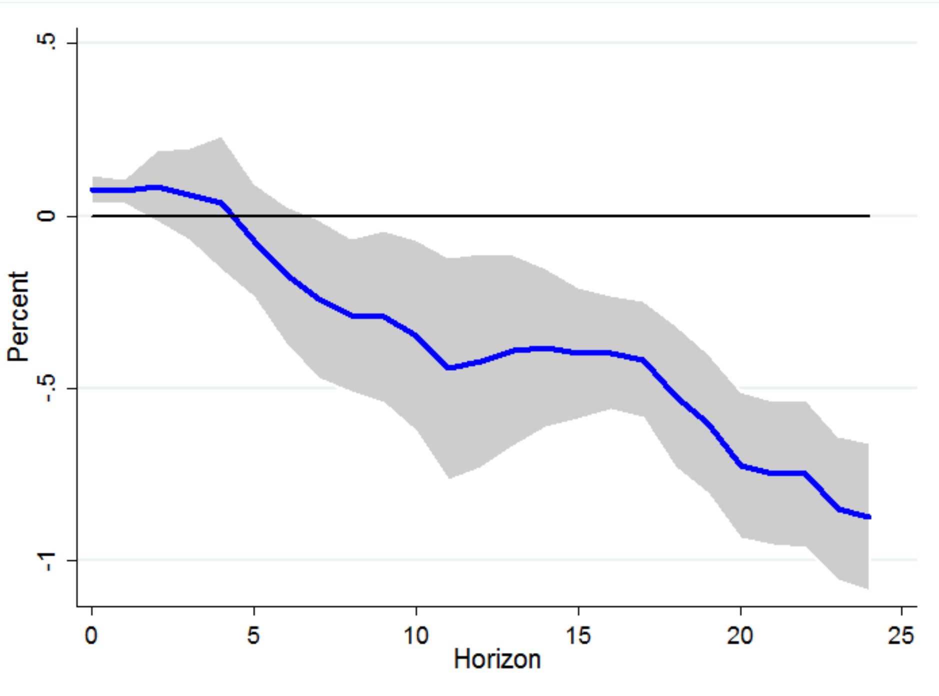
Figure 6.4: Impulse Responses from Baseline Specification

(a) Accumulated Response of House Price Growth to a Shock in $\Delta\beta_{1,t}$, $\Delta\beta_{2,t}$ and $\Delta\beta_{3,t}$




Notes: Each sub-panel presents the accumulated impulse response of house price growth to a one percent increase in $\Delta\beta_{1,t}$, $\Delta\beta_{2,t}$ and $\Delta\beta_{3,t}$ respectively. The grey shaded areas indicate the 90% confidence intervals.

(b) Cumulative HP Response to Average $\Delta\beta_{j,t}$ Shocks associated with a normalised 25 bp rise in cash rate




Note: This figure plots the cumulative response of house price growth obtained from feeding the average $\Delta\beta_{j,t}$ shocks associated with a normalised 25 bp rise in cash rate (Table 6.1) into the baseline specification.

### 6.2.3 Preliminary Evidence on City Heterogeneity

Thus far, it is assumed that a monetary policy shock exerts the same effect on all city house price growth rates. However, as discussed in Chapter 4, the substantial variation in house prices across Australia indicates that it is likely that the effects of monetary policy on house price growth rates also vary across cities. Before introducing city heterogeneity into my model, I provide some preliminary evidence on the differential effects of conventional monetary policy shocks, proxied by changes in $\Delta\beta_{2,t}$ in two ways. The focus on $\Delta\beta_{2,t}$ is motivated by the fact that Australia has not yet moved to the zero lower bound, so it is reasonable to expect changes in short-term interest rates to most likely produce heterogeneous effects on city house prices.

One straightforward way to examine if a monetary policy shock affects house price growth differently across cities is to estimate a simple OLS regression separately for each city. Formally, I estimate:

$$\Delta HP_t = \Gamma_0 + \Gamma_1(L)\Delta\tilde{\beta}_{1,t} + \Gamma_2(L)\Delta\tilde{\beta}_{2,t} + \Gamma_3(L)\Delta\tilde{\beta}_{3,t} + \gamma(L)\Delta HP_{t-1} + \lambda_t + u_t \quad (6.3)$$

where $\lambda_t$ represents a set of year dummies. Notations for other variables are the same as those in equation 5.4. A lag length of $p = 24$ is selected for each of the right hand side variables in all city regressions, except for Sydney which uses a lag of $p = 42$ on the house price changes.[5]

Table 6.4: OLS Estimates

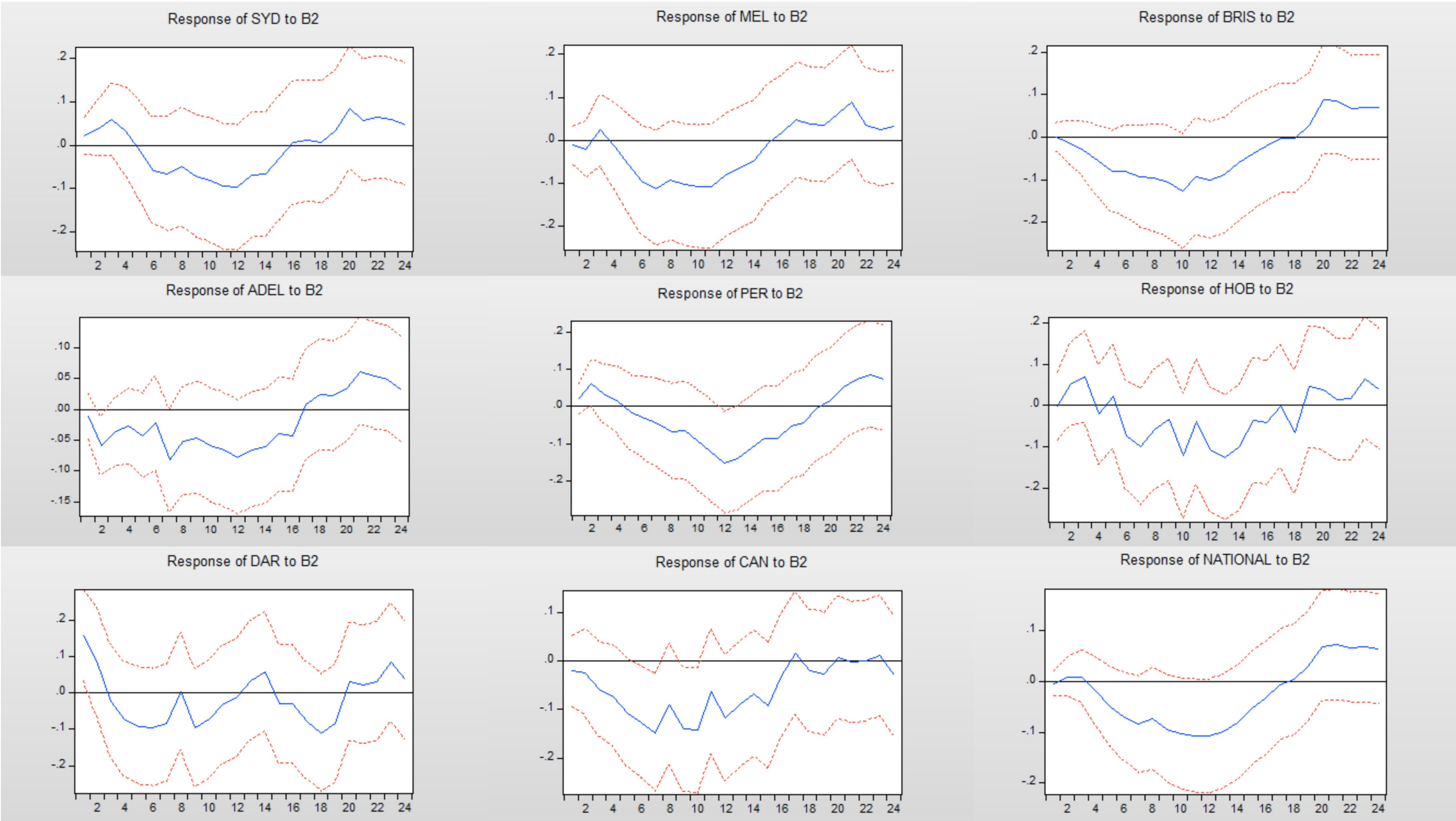
| | Sydney | Melbourne | Brisbane | Adelaide | Perth | Hobart | Darwin | Canberra | National |
|---|---|---|---|---|---|---|---|---|---|
| Sum of $\Delta\beta_{2,t}$ | -5.183 | -5.821 | -0.971 | -8.655 | -15.602 | 8.94 | 34.973 | -18.757 | -0.931 |
| | ( -0.59) | ( -1.05 ) | ( -0.20) | ( -1.78)* | ( -2.08)** | -0.71 | (2.50)** | ( -1.63) | ( -0.25) |
| Observations | 197 | 215 | 215 | 215 | 215 | 215 | 215 | 215 | 215 |

Notes:
A lag length of 24 is attached to each regressor for all city measures except for Sydney, which uses a lag of 42 attached to the change in house prices.
t statistics displayed in parentheses.
* significant at 10%, ** significant at 5%, *** significant at 1%.

Table 6.4 reports the sum of coefficients on $\Delta\beta_{2,t}$ for the city and national regressions, with a full set of estimates provided in Appendix E. As expected, a tighter conventional monetary policy captured by $\Delta\beta_{2,t}$ dampens house price growth for most cities. Importantly, there is substantial variation in the size of the $\Delta\beta_{2,t}$ estimates across cities, supporting my initial hypothesis of the presence of city

[5] The longer lag length used in the Sydney regression is to keep the sum of coefficients on lagged house price changes below 1, in order to get a stable dynamic. This also reflects the stickiness of house prices in the city.

Figure 6.5: Impulse Responses to a Shock in $\Delta\beta_{2,t}$




Note: This figure presents the impulse responses of house price growth for each city and a national measure, to a one standard deviation positive shock to conventional monetary policy (proxied by $\Delta\beta_{2,t}$) based on a structural VAR. The red lines represent $\pm 1$ standard-error bands.

heterogeneity in the effects of a conventional monetary policy shock.

To present a second set of evidence, I estimate a simple two-variable structural VAR, with $\Delta\beta_{2,t}$ being ordered first and followed by $\Delta HP_{it}$, again for each capital city and a national measure.[6] The impulse responses of house price growth to a one percent increase in $\Delta\beta_{2,t}$ using a lag of 24 months are displayed in Figure 6.5.[7] Two features stand out. First, a contractionary monetary policy shock reduces house price growth rates in all cities, broadly consistent with the OLS estimates presented earlier. This result also lends some methodological support to my baseline results estimated using the LP method. Second, the differential adjustment dynamics observed in the city responses further reinforce that a conventional monetary policy shock affects house price growth differently across cities.

To summarise, both the OLS estimates and impulse responses from the structural VAR provide some initial evidence on two important results: 1) as expected, a positive shock in conventional monetary policy, or $\Delta\beta_{2,t}$, negatively impacts the growth rates of house prices in Australia; 2) this negative impact is heterogeneous across different cities.

### 6.2.4 Extended Model

Motivated by the preliminary evidence presented in the previous section, I now formally test if there is city heterogeneity in the effects of conventional monetary policy shocks on house price changes. To do this, I add the interaction terms between the lagged values of $\Delta\beta_{2,t}$ and city dummies to the preferred FE model in equation (5.4) which gives:[8]

$$\begin{aligned}\Delta HP_{it} = \Gamma_0 + \Gamma_1(L)\Delta\hat{\beta}_{1,t} + \Gamma_2(L)\Delta\hat{\beta}_{2,t} + \Gamma_3(L)\Delta\hat{\beta}_{3,t} + \gamma(L)\Delta HP_{it-1} \\ + \Gamma_4(L)(\Delta\hat{\beta}_{2,t} \times CITY_i) + \delta_i + \lambda_t + u_{it}\end{aligned} \tag{6.4}$$

where $CITY_i$ represents a set of city dummy variables. Once again, a lag of 24 months is adopted for all regressors. Since Sydney is the base city, the marginal effect of $\Delta\beta_{2,t}$ on its house price growth is simply $\Gamma_2$. The marginal effect of $\Delta\beta_{2,t}$ on house price growth for all other cities is then given by $(\Gamma_2 + \Gamma_4)$.

[6] This ordering is consistent with the order of data release in a given month: Monetary policy decisions are typically announced at the beginning of month, whereas house price data are released at the end of month.

[7] Note that for consistency purpose, this lag length is common to all cities including Sydney.

[8] This means that the effects of both $\Delta\beta_{1,t}$ and $\Delta\beta_{3,t}$ are assumed to be homogeneous across cities.

I then test if these interaction terms are jointly significant using $F$-tests, for which results are reported in Table 6.5. Indeed, the null hypothesis that the interaction terms with $\Delta\beta_{2,t}$ are jointly equal to zero, is strongly rejected at the 5 percent significance level. This therefore lends support to my earlier results and is generally in line with previous empirical findings that the effects of changes in mortgage rates vary across capital cities in Australia (Otto, 2007; Valadkhani et al., 2019).

To check if this heterogeneity is unique to $\Delta\beta_{2,t}$, I also interact the city dummies with $\Delta\beta_{1,t}$ and $\Delta\beta_{3,t}$ in separate regressions. As seen in Table 6.5, the interaction terms with $\Delta\beta_{3,t}$ are significant at the 10 percent level. Nonetheless, they are not added to equation (6.4) for two main reasons. First, given that Australia has not yet moved to the zero lower bound, it is believed that the effects of conventional monetary policy ($\Delta\beta_{2,t}$), rather than the effects of long-run expectations ($\Delta\beta_{3,t}$), are more important. The second reason is simply to conserve the degrees of freedom. More details on testing the presence of city heterogeneity in the effects of a monetary policy shock are provided in Appendix F.

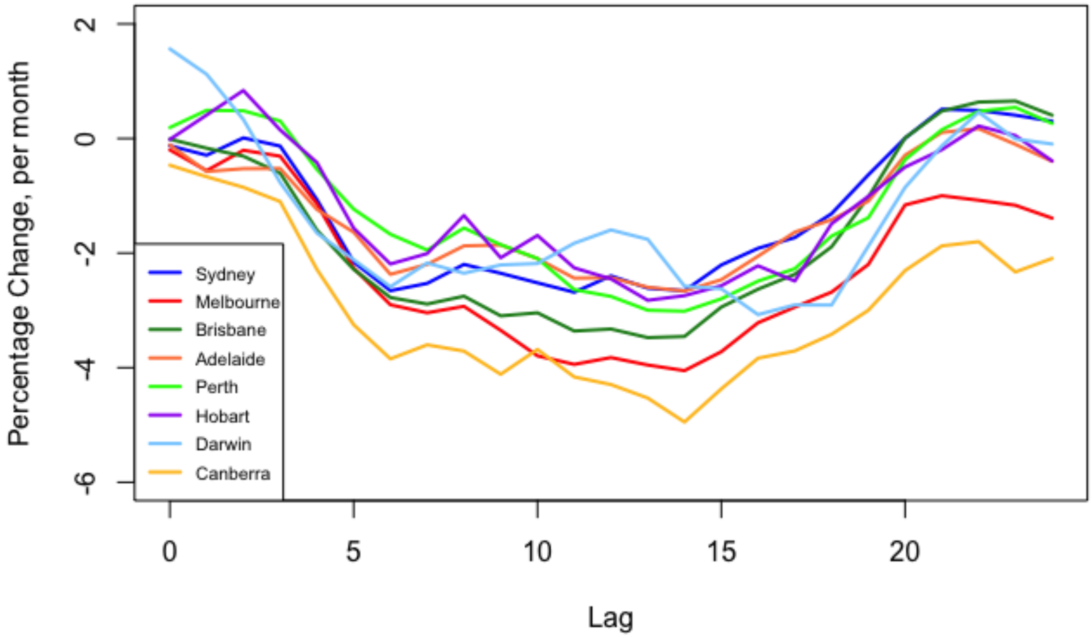
Table 6.5: Joint Hypothesis Test on Interaction Terms[a]

| | (1)<br>$\Delta\beta_{1,t}$ **Interaction** | (2)<br>$\Delta\beta_{2,t}$ **Interaction** | (3)<br>$\Delta\beta_{3,t}$ **Interaction** |
|---|---|---|---|
| $F$-stat[b] | 0.95 | 1.24 | 1.15 |
| $p$-value | (0.6707) | (0.0228)*** | (0.0967)** |

Notes:
a) The preferred FE model shown in equation (5.4) is estimated with the addition of interaction terms with each beta shock in separate regressions.
b) Test of the null hypothesis that the all of the interaction terms are equal to zero, with p-values displayed in parentheses.
* significant at 10%, ** significant at 5%, *** significant at 1%.

Figure 6.6 plots the cumulative marginal effects of $\Delta\beta_{2,t}$ on house price growth for each city against the 24 lags. As is evident, a positive $\Delta\beta_{2,t}$ shock results in varying degrees of cumulative falls in city house price growth rates, providing a direct evidence that the effects of conventional monetary policy shocks do differ across cities. More specifically, stronger negative responses are observed in Canberra, Melbourne and to a lesser extent, Brisbane. These cities have peak effects of -4.95%, -4.05% and -3.48% respectively, which are all realised in 13-14 months. It is somewhat surprising to see that Sydney, the most populous city in Australia, has a relatively weak response to a conventional monetary policy shock. Overall, a positive conventional monetary shock leads to a quite sizeable reduction in house price growth rates across cities.

Figure 6.6: Cumulative Marginal Effects of $\Delta\beta_{2,t}$




Note: This figure presents the cumulative marginal effects of $\Delta\beta_{2,t}$ on city house price growth rates, estimated from the extended model in equation (6.4).

## 6.3 Results for Heterogeneous House Price Responses

### 6.3.1 Alternative Extended Model

The results from the previous section provide direct evidence that the effects of a conventional monetary policy shock do differ across cities in Australia. However, the extended model does not address the issue of what drives these differential house price responses. In this section, I explore one possible mechanism.

Many argue that the soaring property prices in some major cities in Australia such as Sydney and Perth can be attributed to the failure of housing supply to quickly respond to the surge in demand (Gurran and Phibbs, 2013; Yates, 2011). While the sluggish responses of housing supply can be associated with the 'inherent' incapability to develop land and construct new dwellings, Productivity Commission (2004) acknowledges another significant supply constraint, which is the restrictions on building constructions placed by the local and state governments. In order to capture these supply-side influences on house price growth, I turn to the data on the number of building approvals. In the absence of reliable data on planning restrictions in Australia, the number of building approvals can be a good proxy as it reflects the extent to which the government restricts the addition of new properties to the housing market.[9] The tighter the government restrictions, the fewer buildings will be approved and the more constrained housing supply is.

Nonetheless, considering merely the number of dwellings approved in each state gives only a partial picture as it makes sense for new housing supply to concentrate in states with a more rapid population growth and hence higher house prices such as New South Wales and Victoria (Appendix G). To enhance the comparability of housing supply between states, I scale the number of building approvals by the size of population in each state and results are displayed in Panel (a) of Figure 6.7.[10] Additionally, I report the corresponding summary statistics in Table 6.6. One immediate observation is how the supply of new houses is disproportionately distributed across states. The fact that the average of the highest four state building approvals per 10000 head of population is about 1.5 times larger than that of the lowest four, further highlights the considerable heterogeneity.

Table 6.6: Average Building Approvals per 10000 Head of Population

| | Western Australia | Australian Capital Territory | Victoria | Queensland | South Australia | New South Wales | Northern Territory | Tasmania |
|---|---|---|---|---|---|---|---|---|
| **Mean** | 8.7019 | 8.2734 | 7.9068 | 7.5054 | 5.7725 | 5.5098 | 5.2808 | 4.2067 |
| | **Mean of Highest Four** | | | | **Mean of Lowest Four** | | | |
| | 8.0969 | | | | 5.1924 | | | |

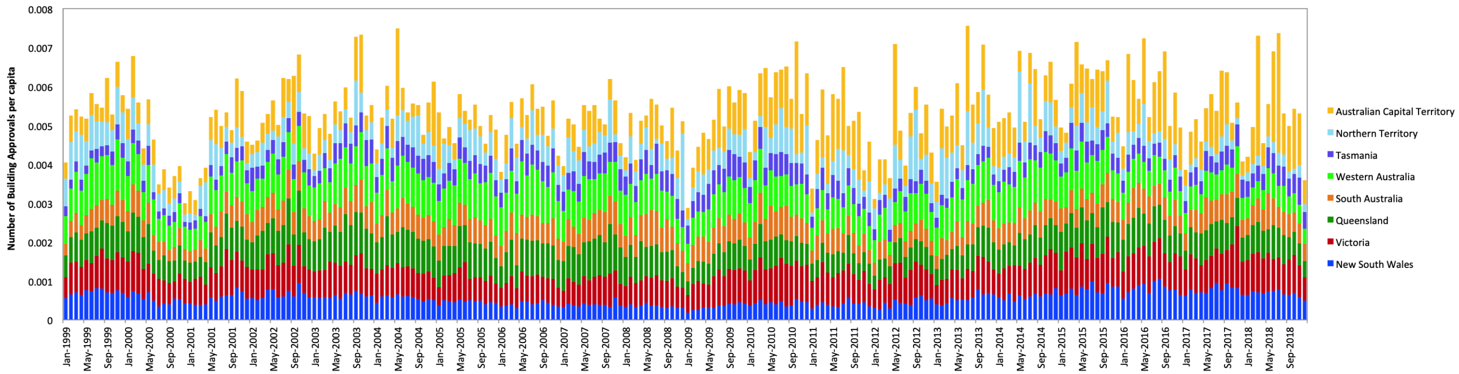
In Panel (b), I reproduce the cumulative marginal effects of $\Delta\beta_{2,t}$ on the city house price growth rates as shown previously in Figure 6.6. A comparison between Panel (a) and (b) reveals some interesting patterns. For example, Canberra, where average building approvals in the corresponding state are the second highest on a per capita basis, has the strongest negative response to a positive shock in conventional monetary policy. Melbourne and Brisbane, with slightly lower average building approvals, have seen smaller cumulative declines in house price changes. In general, a city with a higher level of building approvals appears to have a larger house price response to a conventional monetary policy shock. One exception is observed in Perth, which has the highest state building approvals on average but a much less

[9] I recognise that to some extent building approvals also depend on the number of applications to build new houses and thus reflect the demand for housing. Nonetheless, the number of dwellings approved is ultimately determined by government policies, neighbours' preferences for new developments, amongst other supply side influences.

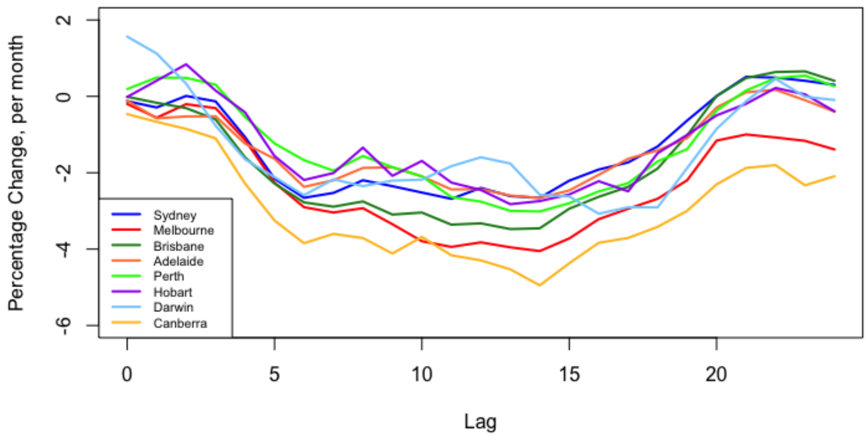
[10] I employ state-level data because city-level data is only available from 2001. Furthermore, since the population series are only available on a quarterly basis, I assume that the size of state population remains constant within a given quarter, which should be reasonable given that population is very slow-moving.

Figure 6.7: Building Approvals and Marginal Effects of $\Delta\beta_{2,t}$

(a) The Number of Building Approvals per Capita




(b) Cumulative Marginal Effects of $\Delta\beta_{2,t}$ in the Extended Model




Notes: Panel (a) reproduces the cumulative marginal effects of $\Delta\beta_{2,t}$ on city house price growth rates, estimated from the extended model in equation (6.4). Panel (b) presents the number of building approvals per capita for each state over the sample period 1999-2018, for which data are derived from the ABS.

responsive house price growth to a shock in $\Delta\beta_{2,t}$. Overall, it can be seen that the patterns displayed in Panel (a) and (b) broadly match each other and this similarity suggests two important points. First, the differences in the level of state building approvals with respect to population size might have been largely captured by the city dummies in equation (6.4). Second, the city heterogeneity in the building approvals per capita might be important in explaining the heterogeneous responses of house price growth to a conventional monetary policy shock.

In light of the observations, I refine the model in equation (6.4) to the following form:

$$\begin{aligned}\Delta HP_{it+h} =&\Gamma_{0,h} + \Gamma_{1,h}(L)\Delta\tilde{\beta}_{1,t} + \Gamma_{2,h}(L)\Delta\tilde{\beta}_{2,t} + \Gamma_{3,h}(L)\Delta\tilde{\beta}_{3,t} + \gamma(L)\Delta HP_{it-1}\\ &+ \Gamma_{4,h}(L)\Delta\tilde{\beta}_{2,t} \times BA_{it-1} + \delta_i + \lambda_t + u_{it+h} \quad \text{for} \quad h = 0, 1, 2, ..., H,\end{aligned} \tag{6.5}$$

which is my alternative extended model. Specifically, I make two amendments to the previous extended model in equation (6.4). First, instead of interacting $\Delta\tilde{\beta}_{2,t}$ with a set of city dummies, I interact them with one-month lagged number of dwellings approved per capita for each state ($BA_{it-1}$). The use of the number of dwellings approved per capita follows Otto (2007) who adopts the same variable to proxy for the effects of government restrictions on new housing supply. Thus, the coefficients $\Gamma_{4,h}$ capture the effects of $\Delta\tilde{\beta}_{2,t}$ on $\Delta HP_{it+h}$ that are due to this supply-side influence and are critical in explaining the heterogeneous house price responses. Second, I adapt the LP methodology to the model in order to construct the dynamic IRFs.

The use of one-month lagged number of dwellings approved per capita is motivated by several reasons. The first concerns the LP method. To date, there is no consensus in the literature on how to deal with contemporaneous causation, so excluding the contemporaneous term in the equation can help alleviate the endogeneity in the impact impulse responses. The second reason comes from the supply-side effects of monetary policy. As pointed out by Mishkin (2007), one of the housing-related channels of monetary transmission mechanism is associated with the direct effects of monetary policy on housing supply. He explains that higher short-term interest rates would increase the cost of building new houses, thereby reducing construction activities. To this end, the one-month lag will reduce some endogenous concerns. Additionally, unlike building commencements or completions, building approvals is the first variable in housing supply chain, meaning that they tend to rise before actual housing supply rises.[11] This characteristic of building approvals being a

[11] The earlier timing of building approvals in the housing supply chain has two other implications.

leading indicator can also help allay the endogeneity of short-term interest rates and housing supply.

### 6.3.2 Impulse Responses for Alternative Extended Model

Given the alternative extended model, I modify the response of city house price growth ($\Delta \mathrm{HP}_{\mathrm{it+h}}$) to a shock ($\varepsilon^{\mathrm{f}}_{\mathrm{t}}(.)$) as follows:

$$\begin{aligned}\frac{\partial \Delta \mathrm{HP}_{\mathrm{it+h}}}{\partial \varepsilon^{\mathrm{f}}_{\mathrm{t}}(.)} &= \sum_{\mathrm{j}} \frac{\partial \Delta \mathrm{HP}_{\mathrm{it+h}}}{\partial \Delta \tilde{\beta}'_{\mathrm{j,t}}} \cdot (\Delta \beta_{\mathrm{j,t}} \cdot \mathrm{d_t}) \\ &= \Gamma_{1,\mathrm{h}}(\Delta\beta_{1,\mathrm{t}} \cdot \mathrm{d_t}) + (\Gamma_{2,\mathrm{h}} + \Gamma_{4,\mathrm{h}} \times \mathrm{BA}_{\mathrm{it-1}})(\Delta\beta_{2,\mathrm{t}} \cdot \mathrm{d_t}) + \Gamma_{3,\mathrm{h}}(\Delta\beta_{3,\mathrm{t}} \cdot \mathrm{d_t})\end{aligned} \tag{6.6}$$

where notations follow those of equation (6.2).

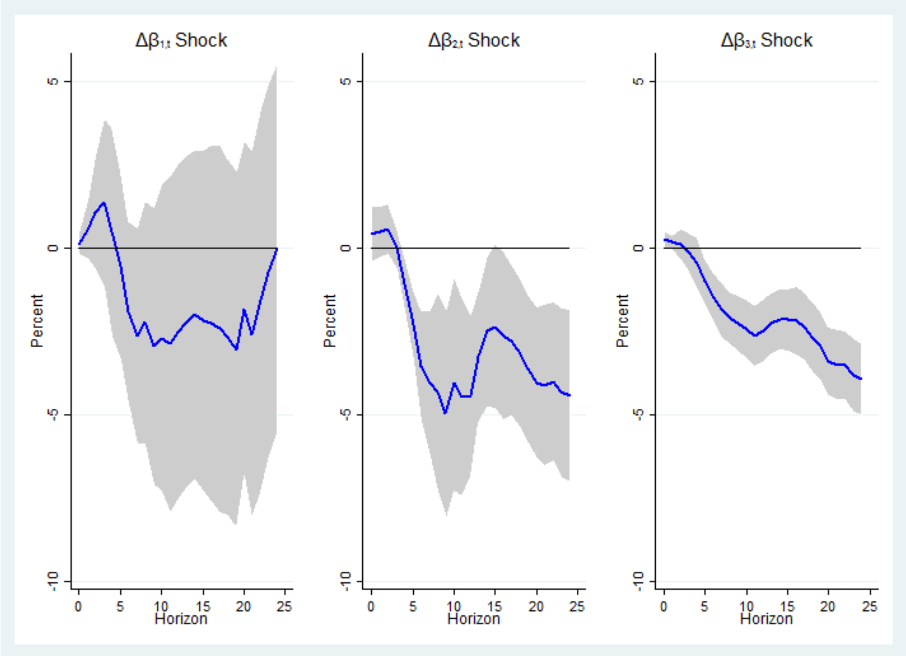
The accumulated responses of house price growth to a one unit increase in each of $\{\Delta\tilde{\beta}_{1,\mathrm{t}}, \Delta\tilde{\beta}_{2,\mathrm{t}}, \Delta\tilde{\beta}_{3,\mathrm{t}}\}$ are displayed in Panel (a) of Figure 6.8 with the 90% confidence intervals. As mentioned earlier, according to the LP methodology, these are simply a sequence of $\{\Gamma_{1,\mathrm{h}}, \Gamma_{2,\mathrm{h}}, \Gamma_{3,\mathrm{h}}\}$ at each horizon $\mathrm{h} = 0, 1, ..., 24$ respectively. As is clear from Panel (a), after a small initial rise, all house price growth deteriorates in response to a positive shock in each of $\{\Delta\tilde{\beta}_{1,\mathrm{t}}, \Delta\tilde{\beta}_{2,\mathrm{t}}, \Delta\tilde{\beta}_{3,\mathrm{t}}\}$. It should be stressed that the middle sub-panel merely plots the effects of a shock in $\Delta\tilde{\beta}_{2,\mathrm{t}}$ without taking into account the parametric interaction with the number of state building approvals per capita, that is, $\Gamma_{2,\mathrm{h}}$. Thus, to get a more meaningful interpretation, I compute and plot in Panel (b) of Figure 6.8 the cumulative marginal effects of a one percent increase in $\Delta\tilde{\beta}_{2,\mathrm{t}}$, calculated at the average one-month lagged level of building approvals per capita for each state, or $(\Gamma_{2,\mathrm{h}} + \Gamma_{4,\mathrm{h}} \times \overline{\mathrm{BA}}_{\mathrm{it-1}})$.

Two important results emerge from Panel (b). The first is associated with the general path of response to a shock. As is clear, all cities experience declines in house price growth within the first three months and reach peak effects in 9-10 months. Nonetheless, compared to the marginal effects generated from the interaction between $\Delta\tilde{\beta}_{2,\mathrm{t}}$ and city dummies (Figure 6.6), those generated from the interaction with the building approvals per capita instead are stronger. For comparison, peak effects in Canberra, Melbourne and Brisbane based on the former specification are -4.95%, -4.05% and -3.48%, whereas those based on the latter specification are -6.68%, -6.56% and -6.43%.

First, it suggests that a longer lag is not necessary. Second, this means that the number of building approvals is a better measure of housing supply than the number of building commencements or the number of building completions.

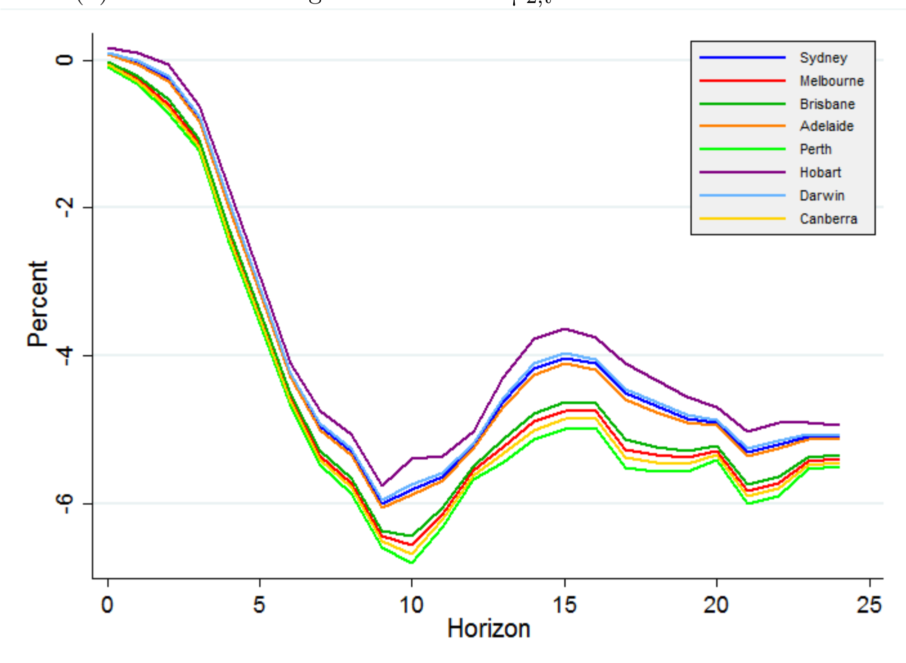
Figure 6.8: Results from Extended Model with Parametric Interaction

(a) Accumulated Response of House Price Growth to a Shock in $\Delta\beta_{1,t}$, $\Delta\beta_{2,t}$ and $\Delta\beta_{3,t}$




Notes: Each sub-panel presents the accumulated impulse response of house price growth to a one percent increase in $\Delta\beta_{1,t}$, $\Delta\beta_{2,t}$ and $\Delta\beta_{3,t}$ respectively. It should be noted that the response to $\Delta\beta_{2,t}$ in the model sub-panel only shows the main effect without taking into account the parametric interaction with state building approvals, i.e. $\Gamma_{2,h}(L)$. The grey shaded areas indicate the 90% confidence intervals.

(b) Cumulative Marginal Effects of $\Delta\beta_{2,t}$ on House Price Growth




Note: This figure plots the cumulative marginal effects of $\Delta\beta_{2,t}$, calculated at the average one-month lagged number of building approvals per capita for each state, i.e. $(\Gamma_{2,h} + \Gamma_{4,h} \times \overline{\mathrm{BA}}_{it-1})$.

These larger point estimates suggest that incorporating the number of building approvals per capita produces a larger impact of a shock in conventional monetary policy on house price growth across Australia. Moreover, the fact that the growth rates show only limited signs of recovering even after 24 months demonstrates the persistent effects of conventional monetary policy shocks.

Second, it can be observed that the responses of house price growth to a shock in $\Delta\tilde{\beta}_{2,t}$ do vary across cities. Despite falling at a similar rate initially, the city house price growth rates start to diverge in 6 months after the shock. The growth in house prices in Perth is the most sensitive to a conventional monetary shock, with a peak negative response of -6.82% realised in 10 months. Disregarding Perth, it is remarkable to see that the relative sensitivity of city house price growth rates is very similar to that shown in Figure 6.6. Nonetheless, relative to Figure 6.6, the extent to which the effects of a conventional monetary policy shock differ across cities becomes smaller, as evidenced by the smaller gaps between responses.

In sum, the results indicate that a higher number of dwellings approved per capita appears to amplify the response of a city house price growth to a shock in conventional monetary policy ($\Delta\tilde{\beta}_{2,t}$). Furthermore, despite a smaller degree of heterogeneity in city responses, a positive conventional monetary policy shock leads to an overall sizeable deterioration in house price growth rates when differences in building approvals are accounted for. In particular, impulse responses show that house price growth in Perth would decline by a massive 6.82% in 10 months after a one percent shock in $\Delta\tilde{\beta}_{2,t}$.

### 6.3.3 Effects of Historical Monetary Policy Announcements

Unlike the traditional identification approach, one key advantage of I&R's (2019) approach is that each announcement will produce a different impulse response according to equation (6.6), depending on how the components $\{\Delta\tilde{\beta}_{1,t}, \Delta\tilde{\beta}_{2,t}, \Delta\tilde{\beta}_{3,t}\}$ vary after a shock. Thus, to show some real policy applications, I consider the effects of six historical monetary policy announcements in Figure 6.9: two rate cuts (Panels (a) and (b)), two rate rises (Panels (c) and (d)) and another two from a no change in cash rate (Panels (e) and (f)). These six examples correspond to dates on which some of the largest shocks (captured by the changes in the three yield curve factors $\{\tilde{\beta}_{1,t}, \tilde{\beta}_{2,t}, \tilde{\beta}_{3,t}\}$) are observed following an announcement. For each example, I show 1) the shift in the yield curve within a one-day window around the policy announcement, 2) the cumulative responses of city house price growth and,[12] 3) the decomposition of Sydney house price response into three components according to equation (6.6).[13] Estimates of each monetary policy shock $\{\Delta\tilde{\beta}_{1,t}, \Delta\tilde{\beta}_{2,t}, \Delta\tilde{\beta}_{3,t}\}$ are also reported on top of each panel.

I first analyse the results at an aggregate level. An inspection of Figure 6.9 illustrates that each monetary announcement, irrespective of the monetary policy action or inaction, yields vastly different house price responses. Panels (a) and (c) display results that are generally in line with the theoretical prediction: a cut (rise) in the short-term interest rate leads to a boost (decline) in the house price growth. A decomposition suggests that the response is mostly explained by changes in $\Delta\tilde{\beta}_{2,t}$ in both instances, meaning that the changes in cash rate are largely unanticipated by the market.

Simply looking at the changes in cash rate, the house price responses in panels (b) and (d) appear to be at odds with what economic theory would predict. Considering first the announcement example in Panel (d), it can be seen that the yield curve has moved in the opposite direction to the change in the policy rate. In particular, the declines in medium-term yields may suggest that the market expects falling interest rates in the future, thereby boosting the nationwide house price growth over the forecast horizon despite the announcement of a rise in cash rate. The fact that changes in $\Delta\tilde{\beta}_{3,t}$, that is, the effects of monetary policy on the medium-term expectations, have played a vital role at all horizons in explaining the positive response lends further weight to this possibility. Overall, the movement in yield curve and the resulting house price response make economic sense. By contrast, the results shown in Panel (b) are surprising, in that a fall in cash rate has caused a downward shift in the long-end of the yield curve, which in turn results in falling house price growth in all cities. As I explain in the next section, this counterintuitive response may be associated with the prevailing market conditions.

Some striking patterns emerge when I compare panels (b) and (f). Despite the different nature of monetary policy announcements, both shocks generate qualitatively similar results. In both instances, the long-end of the yield curve shifts downward whereas house price growth rates in all cities decline immediately after the policy announcement. Thus, it appears that the direction of response of

[12] Since I am estimating the predicted responses to an actual monetary policy announcement, the cumulative marginal effects of $\Delta\tilde{\beta}_{2,t}$ (the second component) are calculated at the observed one-month lagged number of building approvals for each state.

[13] I only show the decomposition of Sydney response because the contributions of the three components are broadly similar across other cities.

house price growth depends on how the shape of the yield curve changes after an announcement, more than the actual movement in the cash rate itself. However, the two sets of results do differ quantitatively. Notice that the announcement of a 1 percent cut in cash rate on 2/12/2008 (Panel (b)) yields a negative response, more than double the size of response associated with a zero change in cash rate on 4/12/2008 (Panel (f)). As such, changes in the cash rate seem to play a role in influencing the magnitude of the house price response. Based on these particular announcements, it can be seen that a monetary policy action produces a larger effect on house price growth than a monetary policy inaction. Despite the similar patterns of the yield curve and house price movements, the contributions of each yield curve factor vary across the two shocks.

Turning to panels (e) and (f), it can be seen that a monetary policy inaction can also generate effects on city house price growth. While the cash rate is unchanged, both announcements have moved the yields, mostly because the decisions are unexpected. Crucially, it is the changes in yields that allow a monetary policy shock to be identified, in turn producing economic impacts.

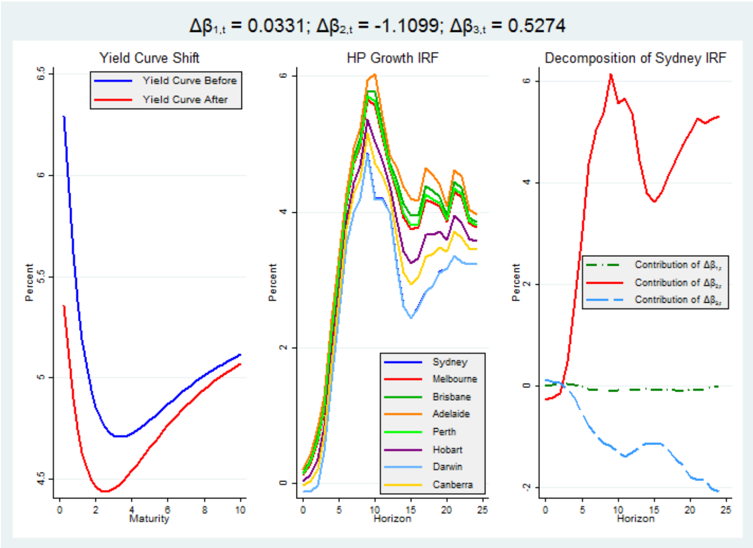
In terms of the heterogenous house price responses, unlike the cumulative marginal effects, the impulse responses for these six historical monetary policy announcements in Figure 6.9 are computed at the *observed* lagged number of dwellings approved per capita. As a result, it can be seen that the patterns of city heterogeneity differ across announcements. For example, the differences between city house price responses to shocks on 4/3/2008 and 6/6/2001 are relatively limited when compared to other shocks. This can be attributed to the extremely similar levels of building approvals per capita across states observed in the month prior to the shocks.

Estimating the house price responses to specific announcements can also reflect the trends in building approvals with respect to population in each state. A closer look at Figure 6.9 shows that house price growth in Canberra has been more reactive to monetary policy shocks particularly those occur in the second half of the sample period. For instance, a shock associated with a 25 basis point rise in cash rate on 2/11/2010 has resulted in a response in Canberra almost twice as large as the responses in other cities, reflecting the much looser supply constraints around that period. If one simply looks at the cumulative marginal effects of $\Delta\tilde{\beta}_{2,t}$ computed at the average values, different variations in the city responses would be missed.

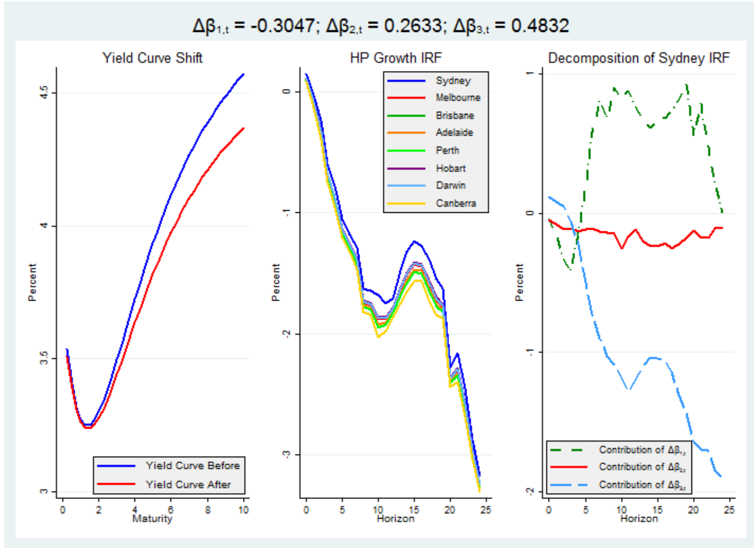
Taken together, these findings demonstrate that each monetary policy announcement can produce a distinct effect on house price growth, irrespective of whether there is a change in the policy rate. In addition, it appears that how the yield curve shifts after a shock plays a bigger role in determining the responses of house price growth. The main takeaway is that the shift in the yield curve ultimately depends on how each monetary policy announcement is perceived by the market participants, which are summarised by the changes in the three yield curve components.

Figure 6.9: Yield Curve Shifts, Cumulative Responses and their Decomposition

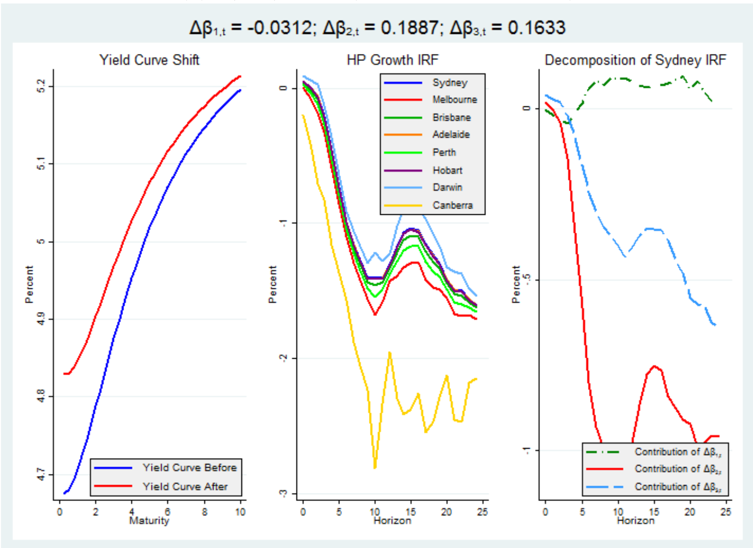
(a) 7/10/2008 (1 percent cut)



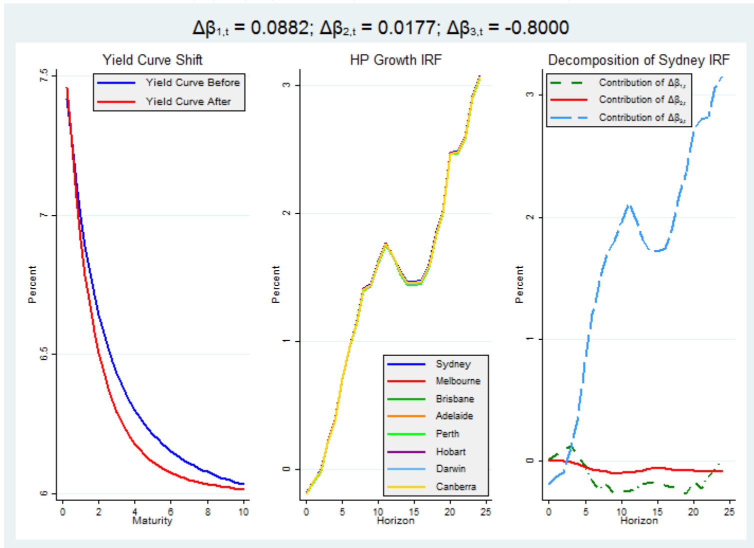

(b) 2/12/2008 (1 percent cut)



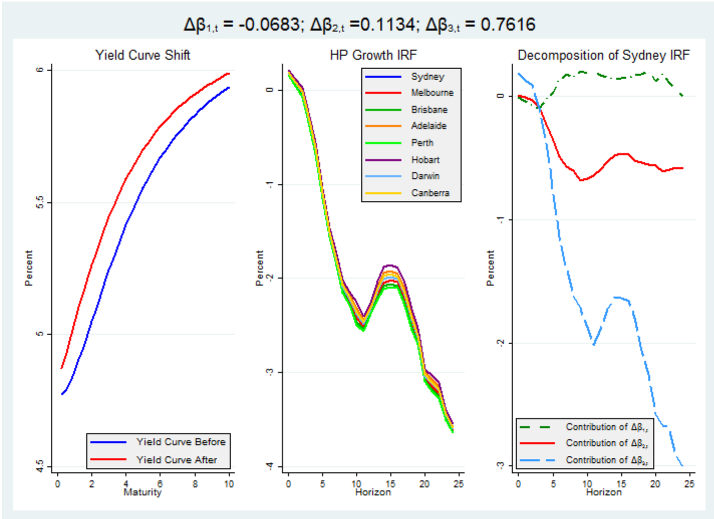

(c) 2/11/2010 (25 basis point rise)



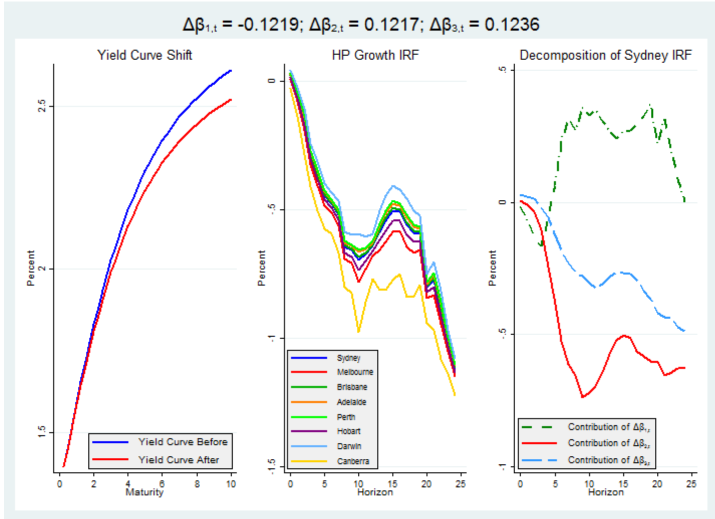

(d) 4/3/2008 (25 basis point rise)



(e) 6/6/2001 (No change)




(f) 4/12/2018 (No change)




Notes: This figure presents the effects of six representative monetary policy announcements. In each panel: I show 1) the yield curve before and after a policy announcement, 2) the cumulative responses of city house price growth calculated at the observed one-month lagged number of building approvals for each state, and 3) a decomposition of Sydney house price response into three components.

# Chapter 7
# Discussion

I now discuss the results presented in Chapter 6. I first give some implications derived from the findings regarding the effects of monetary policy inactions. This is then followed by a discussion of the links between the shifts in the yield curve and the changes in house price growth in response to a monetary policy shock. Finally, I provide an explanation for the observed heterogenous house price responses.

## 7.1 Effects of Monetary Policy Inactions

The last Chapter establishes that each monetary policy shock is distinct as it changes the shape of the yield curve differently, depending on how an announcement is perceived by the market. As a result, the effect on house price growth in each instance can be vastly different.

Using two monetary announcements of a no change in cash rate as examples, I also show that the effects associated with a monetary policy inaction can now be estimated, because a shock is identified based on a change in the entire yield curve rather than a change in the policy interest rate. This has significant implications for the conduct of monetary policy in Australia. At present, the cash rate has not hit the zero lower bound, albeit close. To the extent that the RBA were to push the cash rate to the zero lower bound in the future, they need to influence the market expectations of future path of monetary policy. My results suggest that even when a monetary policy announcement does not change the short-end of the yield curve, it still has the ability to affect the long-term rates. This implies that even if the ability to move the cash rate was constrained, the RBA might still have the potential to affect the economy via its influences on the yield curve.

On the methodological front, this reveals the advantages of I&R's (2019) identification approach, which contrasts with the traditional identification approach. The traditional literature would consider the same change in the short-term interest rate as the same monetary policy shock, when they are obviously very different according to Figure 6.9. Furthermore, in the case where there is no change in the cash rate, the traditional approach would likely identify a zero shock. But my results show that even a monetary policy inaction has the ability to cause a huge impact if the decision is largely unanticipated by the market. Therefore, I&R's (2019) identification approach offers a way to identify a monetary policy shock during both conventional and unconventional times in a unified framework.

## 7.2 Links between Changes in Yields and House Price Responses

When considering a housing loan, it may seem natural to think of the long-term interest rates as the most relevant rates due to the long life of a mortgage. While this may be true in countries with mostly fixed rate mortgages such as the US, the impact of longer-term interest rates on the housing market is not that direct in Australia. As mentioned in the Introduction, the predominant type of mortgage in Australia has a variable rate. Thus, the mortgage rate is more closely linked to the policy rate (Debelle, 2004), as evidenced by the nearly one-for-one relationship depicted in Figure 1.1. An implication is that movements in the short-term interest rates, particularly when they are unexpected, should have a pronounced impact on mortgages and hence house prices. This view is also supported by Mishkin (2007), who explains that countries that have a high share of variable-rate mortgages are subject to a higher sensitivity to short-term interest rate changes and more volatile housing activity.

Returning back to the findings in Figure 6.9 where I present the effects of six representative monetary policy announcements, a comparison between panels (a) and (b) reveals some interesting insights. In Panel (a), a 1 percent cut in the cash rate on 7/10/2008 appears to be largely unexpected by the market, and it has caused a significant reduction in the short-term interest rates relative to the longer rates. On the other hand, a cut of the same magnitude on 2/12/2008 causes a downward shift in the long-end of the yield with short-term rates almost unaffected. Despite the different direction of response, the monetary policy shock in Panel (a) has resulted in a much larger response of house price growth than that shown in Panel (b). Similarly, the shock associated with the announcement of a zero change in cash rate on 4/12/2018 shown in Panel (f) has caused long rates to decrease, leading to only a limited negative impact on house price changes (although this might be partially due to the monetary policy inaction). These examples thus support the view conjectured by Debelle (2004) and Mishkin (2007) that movements at the shorter end of the yield curve have larger effects through their effects on mortgage interest rates.

It should be emphasised that this is not to say the long-term interest rates do not matter. According to the expectations theory, the long-term interest rates essentially represent the expected average variable rate over the life of a mortgage loan.[1] Importantly, movements in the long end of the yield curve indicate changes in people's expectations about the future economy underlying the changes in future short-term interest rates. One possible explanation for the counterintuitive house price responses in panels (b) and (f) is that the prevailing economic conditions play a bigger role in driving the house price responses. The expectations of a lower future short-term rate reflected in the downward shifts in the long end of the yield curve in both episodes may signify that the market is uncertain about the future economic prospects. The heightened uncertainty may significantly dampen the momentum of house prices, thereby contributing to the negative house price responses. This explanation is further supported by the observation that there was a major house price downturn across the country surrounding both years of 2008 and 2018.

Moreover, it appears that when combined with short-term interest rates, changes in the longer-term yields can have another interesting interpretation. This is evidenced by the monetary policy shock on 6/6/2001 in Panel (e), where the announcement has caused all yields to increase by a similar extent. Mishkin (2007) suggests that this parallel upward shift in the entire yield curve might indicate a change in the interest rate regime. Indeed, there have been subsequently 12 instances of a rise in the cash rate from June 2001 to August 2008. As reflected in the large contribution by the changes in $\Delta\tilde{\beta}_{3,t}$, this future change from a low to high interest rate environment might not have anticipated by the market participants. The market could interpret this announcement as the RBA's negative revisions in the future housing market conditions. This consequently dampens households' incentives to take out new mortgages and hence leads to declines in house price growth.

An important caveat to the above discussion is that the observations may or may not hold for all monetary policy announcements. To test the link between changes in yields at different maturities and responses of house price growth more systematically, one can again resort to I&R's (2019) identification approach. Their approach has a key strength in that it produces three estimates of the effects of a monetary policy shock for every announcement $\{\Delta\tilde{\beta}_{1,t}, \Delta\tilde{\beta}_{2,t}, \Delta\tilde{\beta}_{3,t}\}$, with each relating to a different dimension of the shock. In particular, $\Delta\tilde{\beta}_{3,t}$ captures the effects from a monetary policy announcement that change the market expectations regarding the future path of interest rates. By exploiting

[1]Technically speaking, the most relevant yields should be those with a much longer time to maturity such as 25- or 30-year yields. My results on yield curve shifts only include yields with maturities up to 10 years due to data unavailability.

this meaningful interpretation, one can compare the size of the change in this longer-run estimate relative to that in a short-run estimate following all observed announcements. This thus gives a better evaluation of the role played by the longer-term interest rates in affecting house price growth in Australia.

To sum up, my findings reveal the importance of exploring the links between changes in yields and house prices. Particularly, the extent to which the growth in house prices responds to a monetary policy shock may depend on whether short- or longer-term interest rates get affected most. Based on the six selected examples, it seems that changes in the short-term interest rates may have a larger effect on house price growth. A more comprehensive analysis is required to assess whether this holds true throughout the entire sample period.

## 7.3 Heterogeneous House Price Responses

The discussion so far is focused on the overall response of house price growth to a monetary policy shock. Now, I turn to analysing the differential effects of a conventional monetary policy shock on city house price growth.

In Chapter 6, I exploit the variation in building approvals per capita to explore whether or not this can provide an explanation for the differential effects of conventional monetary policy on city house prices. Differences in the levels of building approvals per capita reflect to some extent the different housing supply elasticities between states. Tighter government regulations will translate to a smaller number of dwellings approved per capita, which in turn indicates a more inelastic housing supply. For example, the low level of building approvals per capita in New South Wales as displayed in Panel (a) of Figure 6.7, reflects the state's unresponsiveness of housing supply. As pointed out by Moran (2007), this arises from the fact that Sydney, which dominates the New South Wales housing market, has the most restrictive planning regime in Australia. Gitelman and Otto (2012) also confirm later that the housing supply is relatively inelastic for Sydney.

The standard demand and supply framework suggests that cities with a more inelastic housing supply such as Sydney would experience a larger house price response following a monetary policy shock, whereas cities that are less restricted would respond with more supply adjustments. Nonetheless, my finding that —a city with a higher number of building approvals per capita and hence a more elastic housing supply sees a bigger drop in its house price growth after a positive conventional monetary policy shock —contradicts the prediction of economic theory.

Given this puzzling result, I illustrate one possible explanation using Sydney and Canberra as examples. Both empirical evidence and data on building approvals per capita shown earlier suggest that Sydney has a more inelastic housing supply than Canberra. Now consider that the two markets are hit by a monetary policy shock associated with a cut in the cash rate, this causes the demand curve to shift towards the right. The degree of this demand shift will depend on how responsive housing demand in each city is to interest rate changes. If Canberra has a relatively higher interest rate demand elasticity for reasons such as a greater willingness of households to borrow, it will generate a larger shift in demand in response to a given change in interest rates. Consequently, this can result in a bigger increase in house price growth relative to Sydney despite having a more elastic supply curve. Therefore, differences in city's responsiveness of demand to interest rate changes may offer an explanation to why house prices in a less restricted city would experience a more pronounced effect from a monetary policy shock.

All in all, while data on the level of building approvals per capita broadly matches the cumulative marginal effects of $\Delta\tilde{\beta}_{2,t}$, the model produces results that are counterfactual based on the conventional demand and supply perspective. It therefore cannot be concluded that this channel via the heterogeneity in building approvals per capita contributes to the heterogeneous house price responses to a monetary policy shock. Given the complex dynamics of demand and supply, I recognise that evaluating which mechanism drives the heterogeneous responses of city house price growth is by no means easy. These results, though puzzling, show an attempt at explaining the differences in city house price responses by incorporating a housing supply measure,[2] which is important in the analysis of house prices (Williams, 2009). That said, further investigation into the relationship between housing supply and house prices is warranted.

[2] A common explanation for the differential effects of monetary policy in the literature is that a city with a higher level of house prices means a larger amount of debt will be required to finance the housing purchases. As a result, housing demand and hence house prices will be more exposed to changes in short-term interest rates.

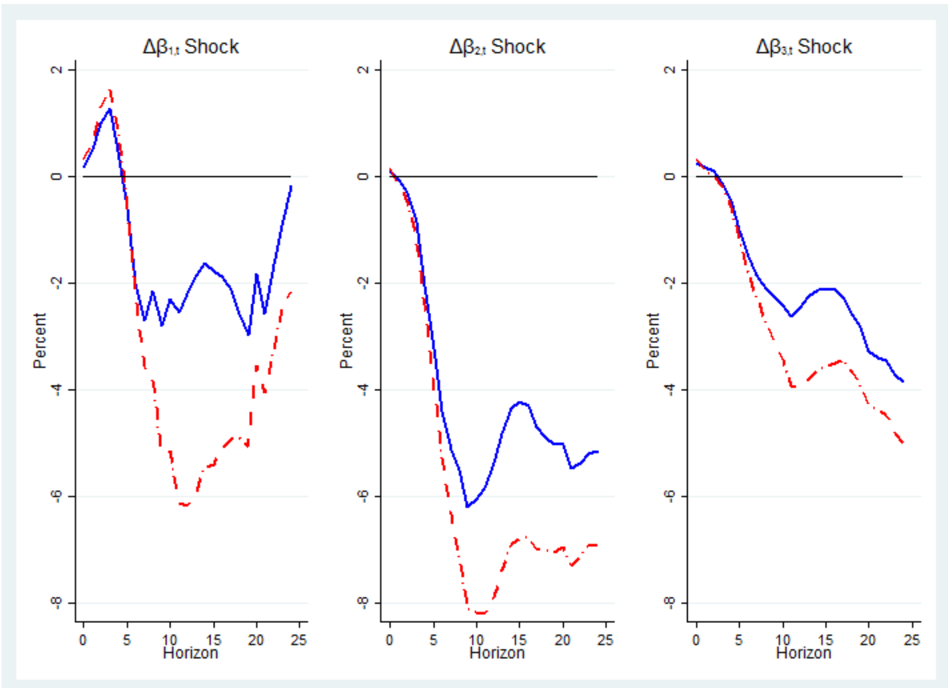
# Chapter 8
# Robustness Checks

This Chapter presents results obtained from three robustness checks: (1) using a longer lag length on house prices, (2) interacting $\Delta\beta_{2,t}$ with a dummy variable indicating the GFC period, and (3) considering an alternative measure of housing supply.

## 8.1 Alternative Lag Length

First, I test if my baseline results are robust to an alternative lag length of house price growth. This is motivated by the substantial persistence of house prices in Australia as discussed throughout the paper. Specifically, I consider 36 lags of the dependent variable in the baseline specification in equation (5.4). It is true that using more lags generally produces a larger negative effect on house price growth rates, which is particularly evident in the response to a shock in $\Delta\beta_{1,t}$. Nonetheless,

Figure 8.1: Cumulative HP Responses using a Different Lag Length



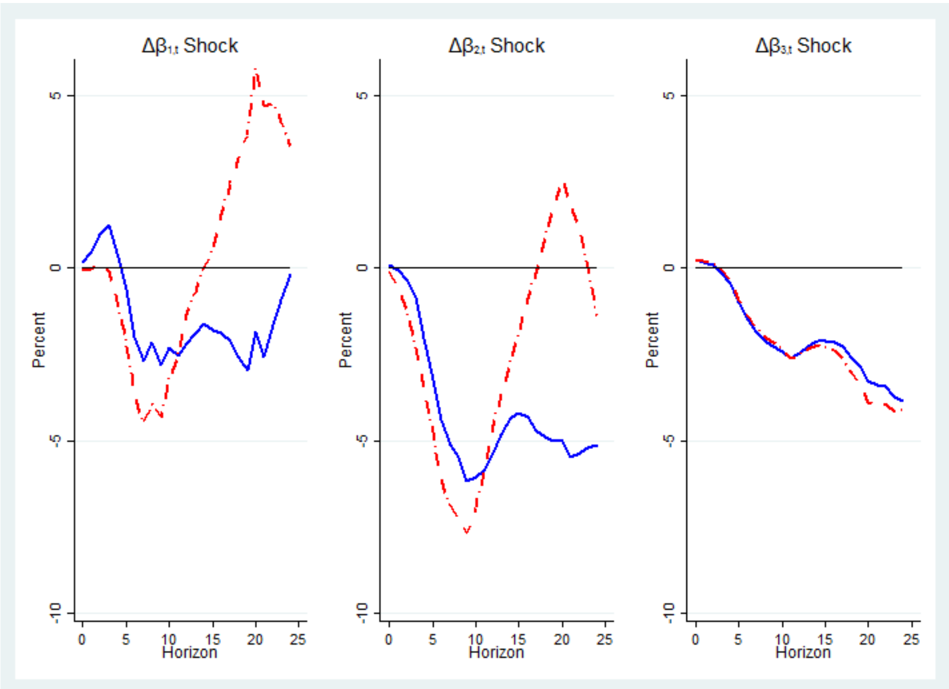

Notes: Blue line represents the original cumulative impulse responses (24 lags on house prices) and red line represents those generated by the benchmark model with the use of 36 lags.

all impulse responses are similar qualitatively, with a one unit increase in each of $\{\Delta\beta_{1,t}, \Delta\beta_{2,t}, \Delta\beta_{3,t}\}$ driving declines in house price growth rates. Moreover, the responses obtained from the two different lag lengths are almost indistinguishable from each other over the shorter horizons, providing further support for the baseline results.

## 8.2 GFC Dummy Variable

To see if the interaction between monetary policy and house price growth is affected by the prevailing economic conditions, I conduct a robustness test where I add a dummy variable representing the GFC period (10/2008-06/2009) to the baseline specification in equation (5.4) with city and year fixed effects.[1] Figure 8.2 shows the original impulse responses of house price growth and those generated by a model with the inclusion of a GFC dummy to a one unit increase in each of $\{\Delta\beta_{1,t}, \Delta\beta_{2,t}, \Delta\beta_{3,t}\}$.

Figure 8.2: Comparison between Benchmark Model with and without GFC Dummy




Notes: Blue line represents the original cumulative impulse responses and red line represents those generated by the baseline model with the addition of the GFC dummy.

As shown, the responses to a shock in $\Delta\beta_{3,t}$ exhibit remarkably similar dynamics over the entire horizon. My findings also indicate that, despite a quicker recovery at the longer horizons, the short-run dynamics of house price growth in response to a

[1]Note that this model does not allow for differences in slope coefficients, hence the effects of all three dimensions of a monetary policy shock are the same for all cities.

positive shock in $\Delta\beta_{1,t}$ and $\Delta\beta_{2,t}$ respectively do not change significantly, in terms of both the direction of response and the timing of peak effect. With regards to the size of peak effect, the model with a GFC dummy produces only marginally larger negative effects on house price growth. Overall, my baseline results are quite robust to the control of the GFC period.

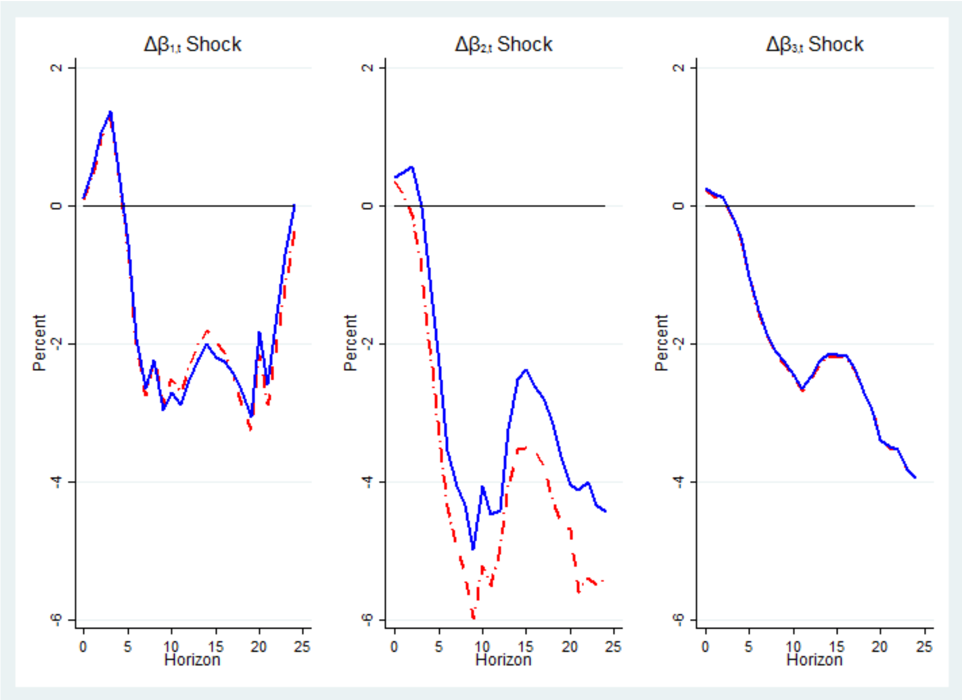
## 8.3 Alternative Measure of Housing Supply

In Section 6.3.1, I consider an alternative extended model in an attempt to understand the mechanism underlying the heterogeneous house price responses. In particular, I interact $\Delta\beta_{2,t}$ with the one-month lagged level of state building approvals per capita in equation (6.5). As a second set of robustness checks, I use instead the number of building approvals as the measure of new housing supply and illustrate how sensitive my results are to the absence of population control. As seen in Panel (a) of Figure 8.3, failing to control for population size produces both qualitatively and quantitatively similar cumulative responses of house price growth. The responses to a positive shock in $\Delta\beta_{3,t}$ even appear to coincide with one another.

While the impulse responses are generally insensitive to alternative measure of housing supply, Panel (b) shows that cumulative marginal effects of $\Delta\beta_{2,t}$ are markedly different to the baseline results shown in Panel (b) of Figure 6.8. The biggest difference can perhaps be observed in the house price growth in Sydney, where sensitivity to a conventional monetary policy has greatly increased when the state population is not accounted for. Overall, the results demonstrate that the heterogenous responses of city house price changes are sensitive to how the housing supply is measured. More importantly, they highlight that controlling for population growth is crucial for comparing between the differential effects of a conventional monetary policy shock across cities, thereby lending support to my measure of housing supply.

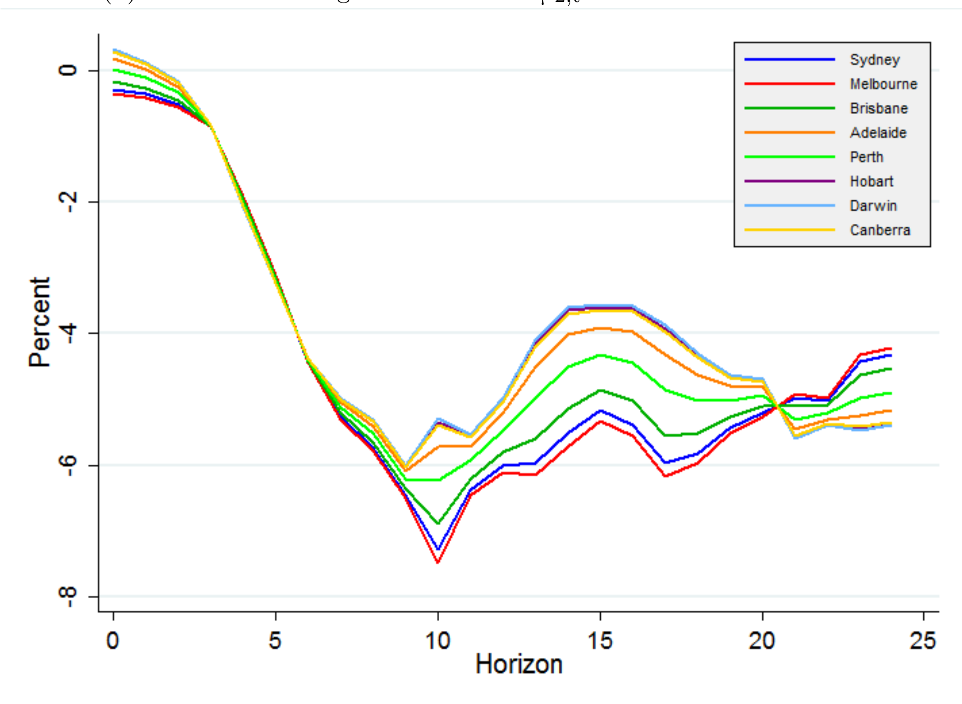
Figure 8.3: Comparison between Housing Supply Measured with and without Population Control

(a) Accumulated Response of House Price Growth to a Shock in $\Delta\beta_{1,t}$, $\Delta\beta_{2,t}$ and $\Delta\beta_{3,t}$




Notes: Blue line represents the original cumulative impulse responses and red line represents those generated by interacting $\Delta\beta_{2,t}$ with one-month lagged level of state building approvals.

(b) Cumulative Marginal Effects of $\Delta\beta_{2,t}$ on House Price Growth




Note: This figure plots the cumulative marginal effects of $\Delta\beta_{2,t}$, calculated at the average one-month lagged number of building approvals for each state, i.e. $(\Gamma_{2,h} + \Gamma_{4,h} \times \overline{BA}_{it-1})$.

# Chapter 9
# Concluding Remarks

## 9.1 Limitations and Future Research Avenues

There are several limitations inherent in my results presented in this paper, which point toward avenues for future research. First, the impulse responses relating to the heterogeneous effects on city house price growth rates are presented in the absence of confidence bands. Part of the reason for this is that I recognise that by construction, there will be serial correlation in the error terms produced by the local projection methodology and correcting for the standard errors will undoubtedly complicate the research. One can typically correct this using the heteroskedasticity and autocorrelation consistent estimator (Newey and West, 1986). Another useful technique is to bootstrap the standard errors. In the context of my research, if the constructed confidence intervals for each city house price growth rates do not overlap one another, this will indicate that the heterogeneity in the city responses to a monetary policy shock is statistically significant and provide further support for my results.

The second limitation relates to my attempt to use the differences in the level of building approved per capita to explain the city heterogeneity in the effects of monetary policy on house price growth. Specifically, my initial motivation for using this measure is to proxy for the effects of supply-side influences such as planning regulations, thereby indicating a degree of housing supply elasticity. Nonetheless, building approvals likely suffer from endogeneity because while government restrictions do play a role in the approval process, the number of residential buildings approved also depends on how many people want to build a house. Ultimately, building approvals are an outcome of both housing demand and supply. In this regard, my analysis on the heterogeneous effects of monetary policy can be improved by research that conducts precise estimations of housing supply elasticities.[1] They can then be used to interact with the exogenous monetary policy shock, thereby testing whether the effects of monetary policy depend on housing supply elasticities in Australia, as with Aastveit and Anundsen (2018).

[1] Note however this is somewhat limited by the availability of data on planning restrictions.

I also acknowledge that the effects of monetary policy on house prices are more than purely a demand or a supply story. To date, empirical studies that examine the relationship between housing supply and house prices are limited, particularly in the Australian context. Further research on this area would be useful in building a more holistic understanding of the interactions between both demand and supply forces in the housing market, which ultimately determine how house prices in different cities would react to monetary policy.

Third, the models presented in this paper do not include any control variables. Despite the remarkable similarity between the patterns of city house price responses displayed in panels (a) and (b) of Figure 6.7, it could be that the city dummies happen to capture a lot of noise, instead of picking up the differences in the building approvals per capita. Hence, including a wide range of controls that have an impact on house price growth such as income and inflation will help provide a clearer picture of the relationship between monetary policy and house prices.

Lastly, my results generated based on the historic monetary policy announcements provide some evidence of asymmetric effects of monetary policy. Given the growing importance of capturing non-linearities in the relationship between interest rates and house prices (Lim and Tsiaplias, 2016), it will be interesting to extend my paper by formally assessing whether or not the effects on city house price growth are subject to the direction and/or the size of a monetary policy shock.

## 9.2 Conclusion

To my knowledge, this paper is the first that applies Inoue and Rossi's (2019) identification approach to the Australian yield curve data and studies the implications of multi-dimensional monetary policy for the city-level house prices in Australia. The work is partly inspired by the increasing instances of a zero change in cash rate, and partly by the deficiency in the Australian housing market literature regarding the city heterogeneity in the monetary policy effects. My results are two-fold and are summarised as follows.

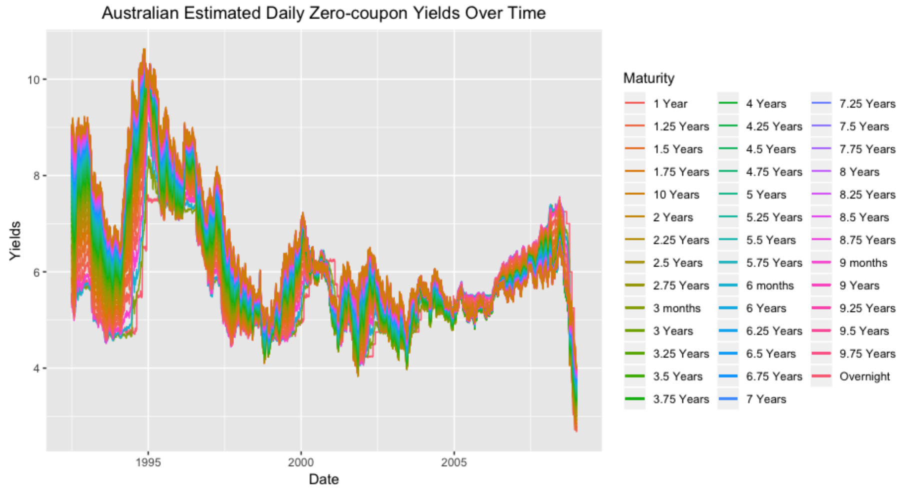
First, I successfully demonstrate that a monetary policy shock can be identified irrespective of whether there is a change in the policy interest rate. Importantly, each monetary policy announcement can produce a distinct effect on house price growth, depending on how it is perceived by the market participants. Given the current interest rate environment in Australia, this finding offers support to future applications of this shock identification approach to unconventional monetary policy.

Second, my results show that a positive monetary policy shock leads to a sizeable deterioration in house price growth across Australia. Additionally, I find that the effects of a conventional monetary policy shock are heterogeneous across the eight capital cities in the country. While the channel via differences in new housing supply explored in this paper cannot be taken as the driver of the observed heterogeneous responses, this does not rule out its potential in explaining the heterogeneous monetary policy effects. The substantial heterogeneity in housing supply elasticities indicates that policy makers should pay attention to how this can affect the house price dynamics. It also suggests that more efforts should be made to obtain a plausible explanation as to why city house price growth responds to a monetary policy shock in a way presented in this paper.

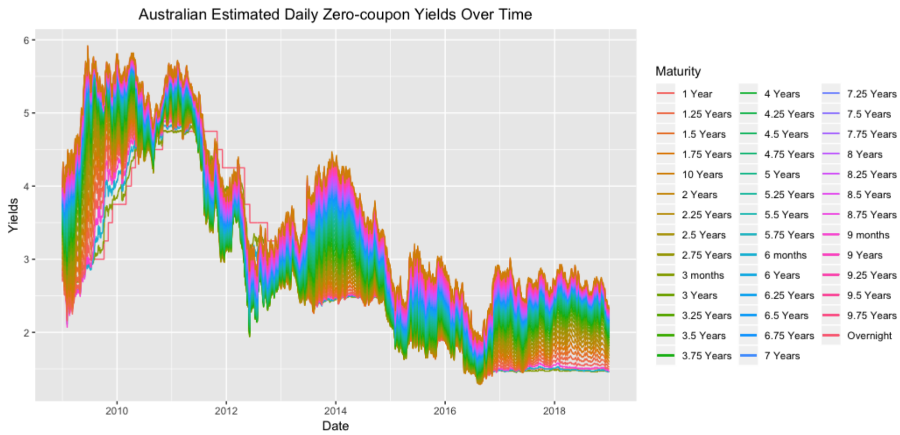
# Appendix A
# Inferred Zero-Coupon Yields

Figure A.1: Australian Estimated Daily Zero-Coupon Yields Over Time

(a) Period 1992-2008




(b) Period 2009-2018

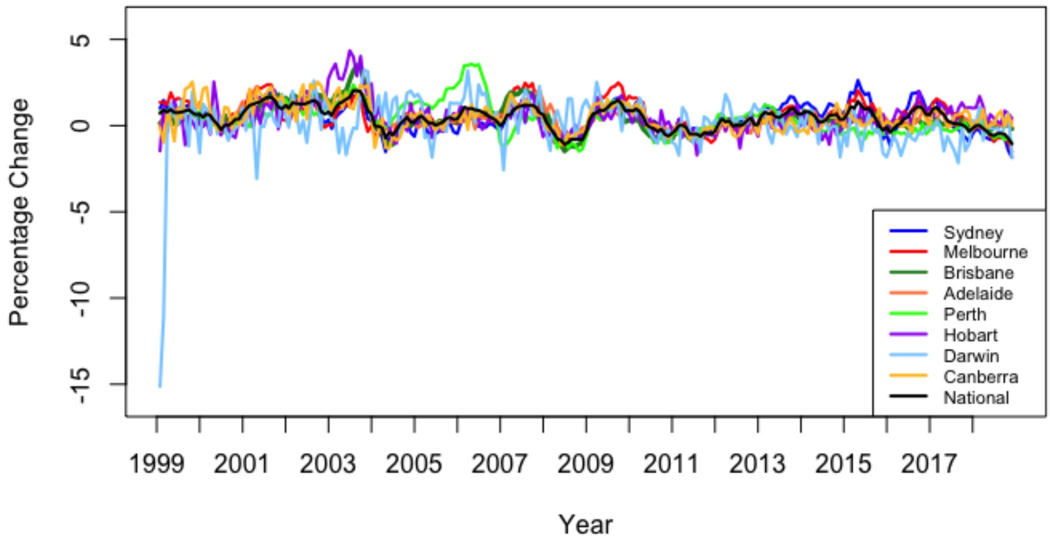



Notes: The figures plot the inferred zero-coupon yields over the periods of 1992-2008 (Panel (a)) and 2009-2018 (Panel (b)). The yields are estimated at maturities in quarterly increment out to 10 years, thus giving a total of 41 maturities. Data are available on the RBA's website.

# Appendix B

# Descriptive Statistics

Figure B.1: Monthly Changes in Hedonic Home Value Indexes




Source: CoreLogic; Monthly changes are calculated as the log difference multiplied by 100.

Table B.1: Descriptive Statistics for the Level of House Prices

| | **Sydney** | **Melbourne** | **Brisbane** | **Adelaide** | **Perth** | **Hobart** | **Darwin** | **Canberra** | **National** |
|---|---|---|---|---|---|---|---|---|---|
| **Mean** | 104.418 | 91.321 | 82.189 | 86.089 | 82.377 | 85.527 | 79.606 | 84.910 | 91.682 |
| **Std** | 33.843 | 32.850 | 22.520 | 23.303 | 24.693 | 24.756 | 21.890 | 23.451 | 26.468 |
| **Max** | 178.113 | 157.163 | 108.599 | 117.064 | 109.201 | 133.807 | 108.933 | 120.815 | 137.400 |
| **Min** | 51.733 | 38.299 | 38.140 | 39.702 | 36.757 | 39.066 | 44.762 | 36.221 | 43.360 |
| | | | | **Contemporaneous Correlations** | | | | | |
| **Sydney** | 1.000 | | | | | | | | |
| **Melbourne** | 0.970 | 1.000 | | | | | | | |
| **Brisbane** | 0.826 | 0.903 | 1.000 | | | | | | |
| **Adelaide** | 0.861 | 0.937 | 0.991 | 1.000 | | | | | |
| **Perth** | 0.685 | 0.793 | 0.940 | 0.927 | 1.000 | | | | |
| **Hobart** | 0.843 | 0.913 | 0.974 | 0.975 | 0.895 | 1.000 | | | |
| **Darwin** | 0.648 | 0.773 | 0.881 | 0.893 | 0.940 | 0.814 | 1.000 | | |
| **Canberra** | 0.856 | 0.933 | 0.985 | 0.994 | 0.915 | 0.980 | 0.875 | 1.000 | |
| **National** | 0.952 | 0.984 | 0.955 | 0.973 | 0.865 | 0.955 | 0.814 | 0.968 | 1.000 |

# Appendix C
# Fixed Effects Specifications

Recall that my baseline specification given by equation (5.4) is a dynamic panel model estimated with city and year fixed effects. This Appendix presents this baseline specification together with three other specifications considered. Details on the notations can be found in Section 5.2. The four models are given by equations (C.1)-(C.4), with results displayed in Table A3.

**Model (1): City Fixed Effects**

$$\Delta HP_{it} = \Gamma_0 + \Gamma_1(L)\Delta\tilde{\beta}_{1,t} + \Gamma_2(L)\Delta\tilde{\beta}_{2,t} + \Gamma_3(L)\Delta\tilde{\beta}_{3,t} + \gamma(L)\Delta HP_{it-1} + \delta_i + u_{it} \quad (C.1)$$

where $\delta_i$ captures the city fixed effects.

**Model (2): City and Month Fixed Effects**

$$\Delta HP_{it} = \Gamma_0 + \Gamma_1(L)\Delta\tilde{\beta}_{1,t} + \Gamma_2(L)\Delta\tilde{\beta}_{2,t} + \Gamma_3(L)\Delta\tilde{\beta}_{3,t} + \gamma(L)\Delta HP_{it-1} + \delta_i + \eta_t + u_{it} \quad (C.2)$$

where $\delta_i$ and $\eta_t$ capture the city and month fixed effects respectively.

**Model (3): City and Year Fixed Effects (Baseline)**

$$\Delta HP_{it} = \Gamma_0 + \Gamma_1(L)\Delta\tilde{\beta}_{1,t} + \Gamma_2(L)\Delta\tilde{\beta}_{2,t} + \Gamma_3(L)\Delta\tilde{\beta}_{3,t} + \gamma(L)\Delta HP_{it-1} + \delta_i + \lambda_t + u_{it} \quad (C.3)$$

where $\delta_i$ and $\lambda_t$ capture the city and year fixed effects.

**Model (3): City, Month and Year Fixed Effects**

$$\Delta HP_{it} = \Gamma_0 + \Gamma_1(L)\Delta\tilde{\beta}_{1,t} + \Gamma_2(L)\Delta\tilde{\beta}_{2,t} + \Gamma_3(L)\Delta\tilde{\beta}_{3,t} + \gamma(L)\Delta HP_{it-1} + \delta_i + \eta_t + \lambda_t + u_{it} \quad (C.4)$$

where $\delta_i$, $\eta_t$ and $\lambda_t$ capture the city, month and year fixed effects.

Table C.1: Fixed Effects Regressions[a]

| | (1) | | | (2) | | | (3) | | | (4) | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | $\Delta\beta_{1,t}$ | $\Delta\beta_{2,t}$ | $\Delta\beta_{3,t}$ | $\Delta\beta_{1,t}$ | $\Delta\beta_{2,t}$ | $\Delta\beta_{3,t}$ | $\Delta\beta_{1,t}$ | $\Delta\beta_{2,t}$ | $\Delta\beta_{3,t}$ | $\Delta\beta_{1,t}$ | $\Delta\beta_{2,t}$ | $\Delta\beta_{3,t}$ |
| Sum[b] | -1.7327 | -1.425 | -0.0617 | -2.5297 | -1.4223 | -0.0526 | 1.4622 | -0.7707 | -1.3396 | -1.0538 | 0.6389 | -0.8873 |
| | ( -1.71) | ( -4.42 )*** | ( -0.20) | ( -2.74 )** | ( -4.74)*** | (-0.17) | (0.20) | ( -0.22) | ( -1.22) | ( -0.15) | (0.17) | (-0.81) |
| Long-run Effect[c] | -13.1804 | -10.8403 | -0.4696 | -19.1232 | -10.7516 | -0.3974 | 7.2127 | -3.802 | -6.6083 | -5.1741 | 3.1371 | -4.3566 |
| | (2.66) | ( 24.91)*** | (0.04) | ( 7.50)*** | ( 24.96)*** | (0.03) | (0.04) | (0.05) | (1.37) | (0.02) | (0.03) | (0.62) |
| City FE | Yes | | | Yes | | | Yes | | | Yes | | |
| Month FE | No | | | Yes | | | No | | | Yes | | |
| Year FE | No | | | No | | | Yes | | | Yes | | |
| Observations | 1720 | | | 1720 | | | 1720 | | | 1720 | | |
| R-sq (within) | 0.7174 | | | 0.7262 | | | 0.7355 | | | 0.7446 | | |
| 'Correct' R-sq[d] | - | | | 0.7138 | | | 0.4741 | | | 0.4708 | | |

Notes:
a) All estimations are obtained using robust standard errors.
b) Test of the null hypothesis that the sum of coefficients on the current and lagged measures of a beta is zero, with t statistics displayed in parentheses.
c) Chi-squared test of the null hypothesis that the long-run effect is zero, with chi-square statistics displayed in parentheses.
d) This is the proportion of variance explained by explanatory variables, obtained by subtracting the R-squared based on the regression on time dummies, from the R-squared based on the regression of the full model (including all explanatory variables and time dummies).
* significant at 10%, ** significant at 5%, *** significant at 1%.

# Appendix D
# Arellano Bond Bias-Correction Procedure

In Chapter 6, I attempt to deal with the Nickell (1981) bias by applying a routine proposed by Arellano and Bond (1991) to my baseline FE model in equation (5.4). The difference GMM estimator can typically be implemented using Stata command xtabond. In finding the most appropriate specification, it is important to limit the instrument count, which tends to increase with T. So I also consider an alternative command xtabond2 developed by Roodman (2009). This is because in addition to its ability to limit the maximum number of lag used in the instrument set as with xtabond, xtabond2 provides a sub-option to additionally collapse the instruments, which is not available in the command xtabond. Several other variants of the difference GMM regression are attempted, including xtabond or xtabond2, robust or default standard error, collapsed or exploded instrument set, excluding $\Delta\beta_{1,t}$ or not (Since $\Delta\beta_{1,t}$ and $\Delta\beta_{2,t}$ might be collinear). After some experimentation, the preferred AB specification, displayed in Column (2) of Table 6.3, is a one-step difference GMM estimator with one lag of dependent variable used as instruments and robust standard errors, which is estimated using the command xtabond. Note that while xtabond2 is able to generate a smaller number of instruments, the command is not chosen due to its poor performance in terms of statistical significance of individual coefficients and system stability. In contrast, the preferred specification produces a larger number of statistically significant coefficients and is able to induce a stable dynamic system. Moreover, almost all of the significant coefficients on lagged $\Delta\beta_{1,t}$ and $\Delta\beta_{2,t}$ have the anticipated negative sign. Almost all coefficients on $\Delta\beta_{3,t}$ have a positive sign with most of them being statistically different from zero. Lastly, it passes the test for second-order autocorrelation of the first-differenced error term.

# Appendix E

# Separate OLS Regressions

Table E.1: Separate OLS Regressions for the Eight Cities and a National Measure

| | **Sydney** | | | **Melbourne** | | | **Brisbane** | | | **Adelaide** | | | **Perth** | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | beta 1 | beta 2 | beta 3 | beta 1 | beta 2 | beta 3 | beta 1 | beta 2 | beta 3 | beta 1 | beta 2 | beta 3 | beta 1 | beta 2 | beta 3 |
| Sum | -38.558 | -5.183 | -4.634 | -37.861 | -5.821 | 1.492 | -17.19 | -0.971 | 0.644 | -17.282 | -8.655 | 3.724 | -32.864 | -15.602 | -4.999 |
| | (-1.93)* | ( -0.59) | ( -1.19 ) | ( -2.56 )** | ( -1.05 ) | -0.58 | ( -1.59) | ( -0.20) | -0.38 | ( -1.35) | ( -1.78)* | ( 2.09)** | (-2.12)** | ( -2.08)** | ( -1.93)* |
| Long-run Effect | -599.663 | -80.602 | -72.072 | -191.541 | -29.45 | 7.549 | -33.522 | -1.893 | 1.255 | -28.612 | -14.329 | 6.166 | -56.168 | -26.665 | -8.544 |
| | -0.01 | -0.01 | -0.01 | -0.37 | -0.31 | -0.37 | -1.48 | -0.04 | -0.16 | -0.81 | -1.45 | -2.39 | (3.64)* | -1.66 | -1.52 |
| Observations | | 197 | | | 215 | | | 215 | | | 215 | | | 215 | |

| | **Hobart** | | | **Darwin** | | | **Canberra** | | | **National** | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | beta 1 | beta 2 | beta 3 | beta 1 | beta 2 | beta 3 | beta 1 | beta 2 | beta 3 | beta 1 | beta 2 | beta 3 |
| Sum | 8.747 | 8.94 | 2.479 | 14.755 | 34.973 | 6.74 | -52.045 | -18.757 | 5.895 | -14.296 | -0.931 | 0.781 |
| | -0.33 | -0.71 | -0.59 | -0.47 | (2.50)** | -1.29 | (-1.90)* | ( -1.63) | -1.65 | ( -1.61) | ( -0.25) | -0.52 |
| Long-run Effect | 8.754 | 8.947 | 2.481 | 9.567 | 22.675 | 4.37 | -51.781 | -18.662 | 5.865 | -41.64 | -2.712 | 2.274 |
| | -0.1 | -0.56 | -0.39 | -0.23 | (3.07)* | -1.83 | -1.49 | -1.35 | ( 2.80)* | -0.98 | -0.06 | -0.37 |
| Observations | | 215 | | | 215 | | | 215 | | | 215 | |

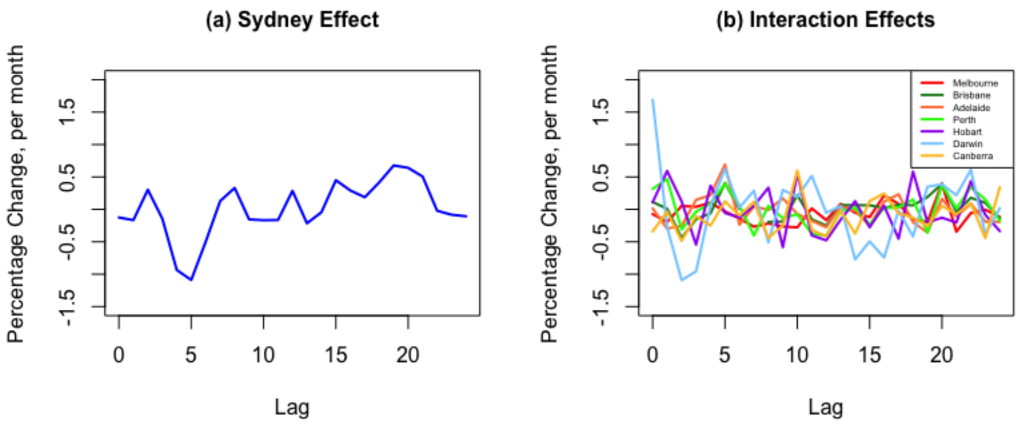
Notes:
a) OLS estimations are obtained using a lag length of 24 attached to each beta and change in house prices for all city measures except for Sydney, which uses a lag of 42 attached to the change in house prices.
b) Test of the null hypothesis that the sum of coefficients on the current and lagged measures of a beta is zero, with t statistics displayed in parentheses.
c) Chi-squared test of the null hypothesis that the long-run effect is zero, with chi-square statistics displayed in parentheses.
* significant at 10%, ** significant at 5%, *** significant at 1%.

# Appendix F

# Testing for Heterogeneous Effects of Conventional Monetary Policy Shocks

Figure F.1 shows visually the dynamic effects of $\Delta\beta_{2,t}$ on the changes in house prices for Sydney and all other cities. The varying patterns of the interaction effects in Panel (b) are a direct evidence that the effects of conventional monetary policy shocks do differ across cities, as a one-unit increase in $\Delta\beta_{2,t}$ leads to a different magnitude of incremental effect on different cities at each lag. However, inspecting Panel (b) might lead one to wonder if the city heterogeneity is driven by Darwin. Thus, I also interact $\Delta\beta_{2,t}$ with all cities but Sydney (base) and Darwin and test if the interaction terms remain significant. Results illustrated in Table F.1 once again confirm that the significant effects of $\Delta\beta_{2,t}$. Interestingly, the $\Delta\beta_{3,t}$ interaction terms become insignificant when Darwin is removed, which reinforces my choice of only interacting $\Delta\beta_{2,t}$. Results for the extended model are provided in Table F.2.

Figure F.1: Dynamic Effects of Interacting with $\Delta\beta_{2,t}$




Notes: Panel (a) plots the marginal effect for Sydney ($\Gamma_2$) against the 24 lags, whereas Panel (b) displays the interaction effects ($\Gamma_4$) for each of the remaining cities.

Table F.1: Joint Hypothesis Test on Interaction Terms Excluding Darwin[a]

| | (1) $\Delta\beta_{1,t}$ **Interaction** | (2) $\Delta\beta_{2,t}$ **Interaction** | (3) $\Delta\beta_{3,t}$ **Interaction** |
|---|---|---|---|
| $F$-stat[b] | 0.97 | 1.21 | 1.09 |
| $p$-value | (0.5794) | (0.0507)** | (0.2244) |

Notes:
a) The preferred FE model shown in equation (5.4) is estimated with the addition of interaction terms with each beta shock in separate regressions.
b) Test of the null hypothesis that the all of the interaction terms are equal to zero, with p-values displayed in parentheses.
* significant at 10%, ** significant at 5%, *** significant at 1%.

Table F.2: Results for the Extended Model

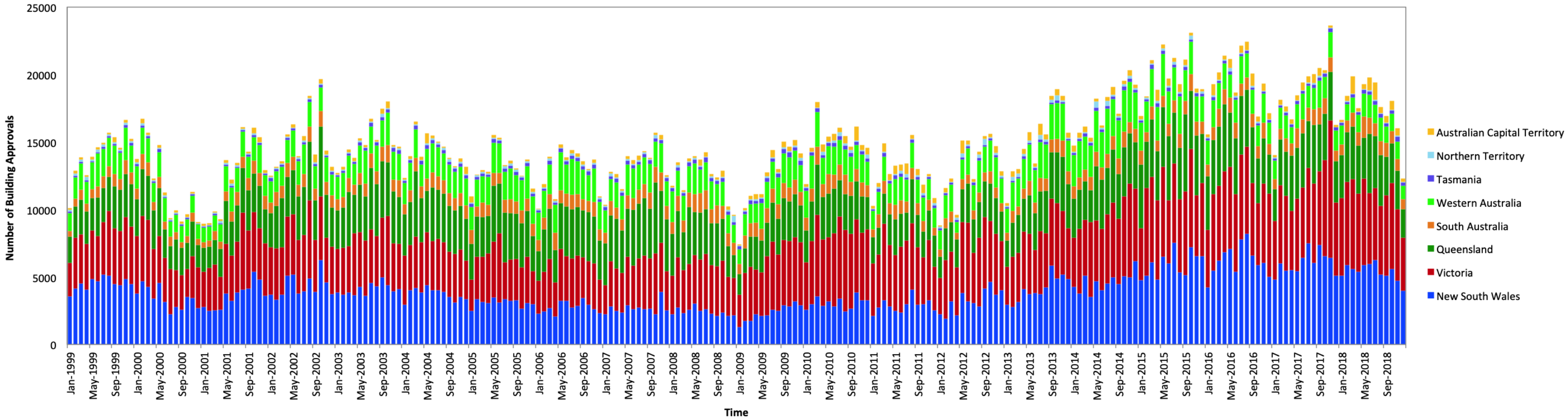
| | $\Delta\beta_{1,t}$ | $\Delta\beta_{2,t}$ | $\Delta\beta_{3,t}$ |
|---|---|---|---|
| Sum | 1.7318 | 0.3008[a] | -1.1422 |
| **Melbourne** | | | |
| Sum of Interaction Effects | | -1.6909 | |
| **Brisbane** | | | |
| Sum of Interaction Effects | | 0.3472 | |
| **Adelaide** | | | |
| Sum of Interaction Effects | | -0.6987 | |
| **Perth** | | | |
| Sum of Interaction Effects | | -0.0375 | |
| **Hobart** | | | |
| Sum of Interaction Effects | | -0.6920 | |
| **Darwin** | | | |
| Sum of Interaction Effects | | -0.3975 | |
| **Canberra** | | | |
| Sum of Interaction Effects | | -2.3942 | |
| Observations | | 1720 | |
| R-sq (within) | | 0.759 | |

Notes:
a) This represents the Sydney main effect.

# Appendix G
# Building Approvals

Figure G.1 displays the number of dwellings approved in each state over the sample period 1999-2018. I show that controlling for population in each state would give drastically different patterns of new housing supply and hence enhance comparability between states.

Figure G.1: The Number of Building Approvals




Note: This figure presents the number of building approvals for each state over the sample period 1999-2018, for which data are derived from the ABS.

# Bibliography


Aastveit, K. and Anundsen, A. K. (2018). Asymmetric effects of monetary policy in regional housing markets.

Abelson, P., Joyeux, R., Milunovich, G., and Chung, D. (2005). Explaining house prices in australia: 1970–2003. *Economic Record*, 81:S96–S103.

Arellano, M. and Bond, S. (1991). Some tests of specification for panel data: Monte carlo evidence and an application to employment equations. *The review of economic studies*, 58(2):277–297.

Bond, S. R. (2002). Dynamic panel data models: a guide to micro data methods and practice. *Portuguese economic journal*, 1(2):141–162.

Brugnolini, L. (2018). About Local Projection Impulse Response Function Reliability. CEIS Research Paper 440, Tor Vergata University, CEIS.

Christiano, L. J., Eichenbaum, M., and Evans, C. L. (1996a). Identification and the effects of monetary policy shocks. *Financial factors in economic stabilization and growth*, pages 36–74.

Christiano, L. J., Eichenbaum, M., and Evans, C. L. (1996b). The effects of monetary policy shocks: evidence from the flow of funds. *The Review of Economics and Statistics*, 78(1):16–34.

Christiano, L. J., Eichenbaum, M., and Evans, C. L. (1999). Monetary policy shocks: What have we learned and to what end? *Handbook of macroeconomics*, 1:65–148.

Debelle, G. (2004). Household debt and the macroeconomy. *BIS Quarterly Review*.

Dvornak, N. and Kohler, M. (2007). Housing wealth, stock market wealth and consumption: a panel analysis for australia. *Economic Record*, 83(261):117–130.

Elbourne, A. (2008). The uk housing market and the monetary policy transmission mechanism: An svar approach. *Journal of Housing Economics*, 17(1):65–87.

Evans, C. L. and Kuttner, K. N. (1998). *Can VARs describe monetary policy?*, volume 9812. Federal Reserve Bank of New York.

Finlay, R. and Chambers, M. (2009). A Term Structure Decomposition of the Australian Yield Curve. *The Economic Record*, 85(271):383–400.

Fry, R. A., Martin, V. L., and Voukelatos, N. (2010). Overvaluation in australian housing and equity markets: wealth effects or monetary policy? *Economic Record*, 86(275):465–485.

Gitelman, E. and Otto, G. (2012). Supply elasticity estimates for the sydney housing market. *Australian economic review*, 45(2):176–190.

Giuliodori, M. (2004). Monetary Policy Shocks and the Role of House Prices Across European Countries. DNB Working Papers 015, Netherlands Central Bank, Research Department.

Gurkaynak, R. S., Sack, B. P., and Swanson, E. T. (2004). Do actions speak louder than words? the response of asset prices to monetary policy actions and statements.

Gurran, N. and Phibbs, P. (2013). Housing supply and urban planning reform: The recent australian experience, 2003–2012. *International journal of housing policy*, 13(4):381–407.

Gürkaynak, R. S. and Wright, J. H. (2012). Macroeconomics and the term structure. *Journal of Economic Literature*, 50(2):331–367.

Haug, A. A., Karagedikli, Ö., and Ranchhod, S. (2005). Monetary policy transmission mechanisms and currency unions: A vector error correction approach to a trans-tasman currency union. *Journal of Policy Modeling*, 27(1):55–74.

Himmelberg, C., Mayer, C., and Sinai, T. (2005). Assessing high house prices: Bubbles, fundamentals and misperceptions. *Journal of Economic Perspectives*, 19(4):67–92.

Iacoviello, M. M. (2000). House prices and the macroeconomy in europe: results from a structural var analysis.

Inoue, A. and Rossi, B. (2019). The effects of conventional and unconventional monetary policy: a new approach.

Jarocinski, M. and Smets, F. (2008). House prices and the stance of monetary policy.

Jordà, Ò. (2005). Estimation and inference of impulse responses by local projections. *American economic review*, 95(1):161–182.

Jordà, Ò., Schularick, M., and Taylor, A. M. (2015). Betting the house. *Journal of International Economics*, 96:S2–S18.

Lim, G. and Tsiaplias, S. (2016). Non-Linearities in the Relationship between House Prices and Interest Rates: Implications for Monetary Policy. Melbourne Institute Working Paper Series No. 2/16, Melbourne Institute of Applied Economic and Social Research, The University of Melbourne.

Mishkin, F. S. (2007). Housing and the monetary transmission mechanism. Working Paper 13518, National Bureau of Economic Research.

Moran, A. (2007). Land regulations, housing prices and productivity. *Agenda: A Journal of Policy Analysis and Reform*, 14(1):35.

Nelson, C. and Siegel, A. F. (1987). Parsimonious modeling of yield curves. *The Journal of Business*, 60(4):473–89.

Newey, W. K. and West, K. D. (1986). A simple, positive semi-definite, heteroskedasticity and autocorrelationconsistent covariance matrix. Working Paper 55, National Bureau of Economic Research.

Nickell, S. (1981). Biases in dynamic models with fixed effects. *Econometrica*, 49(6):1417–1426.

Otto, G. (2007). The growth of house prices in australian capital cities: What do economic fundamentals explain? *Australian Economic Review*, 40(3):225–238.

Piazzesi, M. (2010). Chapter 12 - affine term structure models. In AÏT-SAHALIA, Y. and HANSEN, L. P., editors, *Handbook of Financial Econometrics: Tools and Techniques*, volume 1 of *Handbooks in Finance*, pages 691 – 766. North-Holland, San Diego.

Productivity Commission (2004). First Home Ownership. Technical report, University Library of Munich, Germany.

Romer, C. D. and Romer, D. H. (1989). Does monetary policy matter? a new test in the spirit of friedman and schwartz. Working Paper 2966, National Bureau of Economic Research.

Roodman, D. (2009). A Note on the Theme of Too Many Instruments. *Oxford Bulletin of Economics and Statistics*, 71(1):135–158.

Sims, C. A. (1980). Macroeconomics and reality. *Econometrica*, 48(1):1–48.

Sims, C. A. (1992). Interpreting the Macroeconomic Time Series Facts: The Effects of Monetary Policy. Cowles Foundation Discussion Papers 1011, Cowles Foundation for Research in Economics, Yale University.

Swanson, E. T. (2007). What we do and don't know about the term premium. *FRBSF Economic Letter*.

Valadkhani, A., Nguyen, J., and O'Brien, M. (2019). Asymmetric responses of house prices to changes in the mortgage interest rate: evidence from the australian capital cities. *Applied Economics*, 51(53):5781–5792.

Wadud, I. M., Bashar, O. H., and Ahmed, H. J. A. (2012). Monetary policy and the housing market in australia. *Journal of Policy Modeling*, 34(6):849–863.

Williams, D. M. (2009). House prices and financial liberalisation in Australia. Economics Series Working Papers 432, University of Oxford, Department of Economics.

Yates, J. (2011). Housing in Australia in the 2000s: on the agenda too late? In *The Australian Economy in the 2000s*, RBA Annual Conference Volume (Discontinued). Reserve Bank of Australia.