

Modelling House Prices in Australian Capital Cities

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Declaration

I hereby declare that this submission is my own work. It contains no material that has previously been submitted to any other university or institution as part of the requirements for a degree or other award.

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TABLE OF CONTENTS

1. INTRODUCTION.....	8
2. ISSUES WITH HOUSE PRICE MODELLING.....	12
2.1 DETERMINANTS OF HOUSE PRICES.....	12
2.2 DEALING WITH NON-STATIONARITY	16
2.3 ACCOUNTING FOR BUBBLES	19
2.4 HOUSE PRICE MODELS WITH VARIATION ACROSS SPACE	20
3. THE THEORETICAL MODEL	22
4. THE DATA.....	27
4.1 DATA CONSTRUCTION	27
4.2 SUMMARY STATISTICS.....	29
4.3 MOVEMENTS IN THE VARIABLES	31
5. EMPIRICAL RESULTS.....	36
5.1 ASSESSING WHETHER THERE IS EVIDENCE OF COINTEGRATION	36
5.2 DYNAMIC ORDINARY LEAST SQUARES (DOLS) ESTIMATION.....	39
5.3 THE BOOM IN HOUSE PRICE TO INCOME RATIOS	43
5.4 STRUCTURAL BREAKS	45
5.5 ADDITIONAL VARIABLES	50
6. CONCLUSION	54
7. APPENDICES.....	56

A1. DATA	56
A2. TABLES	61
A3. FIGURES	63
8. REFERENCE LIST.....	65

LIST OF TABLES AND FIGURES

TABLE 1 – SUMMARY STATISTICS.....	29
FIGURE 1 – SYDNEY HOUSE PRICE TO INCOME RATIO	33
FIGURE 2 – QUARTERLY REAL INTEREST RATE IN SYDNEY	34
FIGURE 3 - HOUSING INTEREST PAYMENTS TO DISPOSABLE INCOME RATIO.....	34
TABLE 2 – OLS REGRESSIONS FOR AUGMENTED ENGLE GRANGER TESTS	37
TABLE 3 – DOLS REGRESSION WITH ONE LEAD AND LAG	40
TABLE 4 – DOLS REGRESSION WITH FOUR LEADS AND LAGS.....	40
FIGURE 4 – REGRESSION RESIDUALS	44
FIGURE 5 – PERTH 1986:2 TO 2007:4	46
FIGURE 6 –PERTH 1986:2 TO 2001:3.....	46
FIGURE 7 – SUP-F STATISTIC	48
TABLE 5 – STRUCTURAL BREAK STATISTICS	48
FIGURE 8 – INF-AEG STATISTIC	50
TABLE 6 – ADDITIONAL AUGMENTED ENGLE GRANGER STATISTICS	53
TABLE A1 – OLS REGRESSIONS WITH ADDITIONAL VARIABLES FOR CAPITAL CITIES	61
TABLE A2 – OLS REGRESSIONS WITH THE NON-LINEAR REAL INTEREST RATE	62
FIGURE A1 - HOUSE PRICE TO INCOME RATIOS FOR MELBOURNE AND BRISBANE	63
FIGURE A2 - HOUSE PRICE TO INCOME RATIOS FOR ADELAIDE AND PERTH	63
FIGURE A3 - HOUSE PRICE TO INCOME RATIOS FOR HOBART AND DARWIN	64
FIGURE A4 - HOUSE PRICE TO INCOME RATIOS FOR CANBERRA AND AUSTRALIA.....	64

ABSTRACT

This paper finds some evidence of a cointegrating relationship between real median house price to real income ratios, real interest rates, and the housing interest payments to disposable income ratio in the Australian data at the capital city level. Dynamic ordinary least squares estimates of the long run equilibrium relationship suggest that the housing interest payments to disposable income ratio plays an important role in the determination of house prices. In particular, rises in this variable account for the increase in real median house price to real income ratios in the early 2000's. The estimates of the coefficient on the real interest rate are inconsistent with the underlying theory, suggesting that there may be an omitted variable problem with these regressions.

1. Introduction

It is relatively standard to model house prices as a function of macroeconomic variables thought to influence the demand and supply of housing. Abelson et al. (2005) find evidence of a cointegrating relationship between Australian house prices and variables that they argue are determinants of house prices over the period 1970 to 2003. They subsequently estimate regressions that provide insights into the long run effect of variables at the national level on house prices in Australia. Otto (2007), however, finds that the manner in which the growth rates of certain variables influence the growth rate of house prices displays considerable variation across Australian capital cities. Consequently, if there is evidence of a cointegrating relationship between house prices and fundamentals at the capital city level in the Australian data, then the estimation of econometric models can provide insights into the long run effect of key variables on house prices in the various capital cities. The objective of this paper is to assess whether there is evidence of a cointegrating relationship between house prices and some determinants of house prices in the Australian data at the capital city level, and if so, to obtain insights into the long run effect of these determinants on house prices.

Insights into the causes of movements' in house prices are valuable, as the changes in the housing market can have a powerful impact on the economy. Housing wealth and housing debt are large components of total household wealth and total household debt. Consequently, movements' in house prices affect the confidence of homeowners and their decision-making process. This can have a direct effect on the economy, through the

influence on key variables such as consumption. Many also argue that the housing market has a wider effect on the economy. Recently, in the U.S., defaults on housing loans have affected the viability of financial institutions worldwide, and are consequently a serious threat to economic growth. This illustrates both the effect that the housing market can have on the economy, and the importance of gaining accurate insights into the manner in which house prices are determined.

There has been considerable movement in house prices and house price to income ratios across Australia in recent times. In the period 2000 to 2007, real median house price to real income ratios rose across the country, by more than 100% in some State capital cities. Meanwhile, movements in house price to income ratios in the other direction are also a cause for concern, such as a 7% reduction in the Sydney real median house price to real income ratio between the beginning of 2004 and the end of 2007, and a 13% reduction in Canberra's real median house price to real income ratio between September 2003 and September 2005. These rapidly changing house price to income ratios pose a number of concerns about their effect on the economy. This paper, however, aims to provide insights into the cause of the movements, rather than information on their effect.

The theoretical model in this paper looks at the demand for housing as a function of the amount that potential homeowners can borrow from lenders, which is a function of real interest rates, real incomes, and a measure of credit conditions. The basis of this model is that most house purchases are mortgage financed, and potential homeowners will purchase a home for the price that lenders will lend them. Lenders in the model choose to

lend homeowners the present value of some proportion of their income over the lifetime of the mortgage. This leads to an equation that relates house price to income ratios to real interest rates and the proportion of disposable income spent on mortgage repayments.

Empirical estimations of this model suggest that there is a cointegrating relationship between these variables in Sydney, Perth, and at the national level. In the other State capital cities, however, there is insufficient evidence to reject a null hypothesis of no cointegration. The inclusion of additional variables, or methods to account for structural breaks, does not change this result.

Cointegrating regressions provide insights into the long run effects of variables on house prices. The estimations of the coefficient on the variable measuring credit conditions, the housing interest payments to disposable income ratio, suggest that this variable plays an important role in the determination of house prices. Dynamic ordinary least squares estimates of the coefficient on this variable suggest that the effect of a one-unit increase in the ratio range from a 3.64% increase in real median house price to real income ratios to a 15.11% increase, depending on the city. These are all statistically significant at the 5% level, apart from the estimates for Darwin. The estimates suggest that the rise in the housing interest payments to disposable income ratio is a contributor to the recent rise in house price to income ratios, in all cities except Darwin. In the context of the theoretical model, this suggests that lenders are lending potential homeowners more as a proportion of their income, which is leading to a rise in the demand for housing, raising house prices independently of incomes.

The estimates of the long run effect of real interest rates on house prices vary across cities. The estimates suggest that a rise in the real interest rate of one percent per quarter leads to an effect on the house price to income ratio between an increase of 11.36% and a decrease of 39.99%, depending on the city. The majority of these estimated coefficients are statistically insignificant at the 10% level. A number of the estimates are positive, which is not consistent with the underlying theory. This may be suggesting that one or more important variables have been omitted from the model.

The structure of the paper is as follows. Section 2 discusses issues that commonly arise in the creation of house price models, and the way that previous authors have dealt with these issues. Section 3 describes the theoretical model that is the basis for the empirical equation. Section 4 looks at the data, describing how it is created, and analysing the features of the data. Section 5 involves an examination of the empirical results, and Section 6 concludes.

2. Issues with House Price Modelling

Econometricians that model house prices as inverted demand or reduced form equations typically need to deal with a number of key issues when specifying their models. These issues require the econometricians to decide on the variables to include in the model, the manner in which to deal with the non-stationary data and account for price bubbles, and whether to allow for variation across space. In this section, I look at the decisions taken by econometricians in previous studies, and the information provided by their empirical estimates.

2.1 Determinants of house prices

Econometricians that model house prices as functions of macroeconomic variables must decide which macroeconomic variables to include as determinants of house prices. There are a variety of theoretical justifications for including different variables. For example, some authors begin with an equation that holds in equilibrium, and these range from the real price of houses equals the real price of household services from ownership (Holly et al., 2006) to the structure value of a house equals the replacement cost (Bourassa & Hendershott, 2001). Despite these different theoretical models, the variables included in the models are similar. Thus, a look at the empirical results of previous studies provides insights into the variables that play an important role in the determination of house prices.

One macroeconomic variable usually included in these types of models is a measure of incomes. Incomes are, theoretically, an essential component of the demand for housing.

One argument is that as income levels rise, more potential homeowners can gain access to finance for housing, and those who can gain access to finance are able to increase the amount that they can access, holding other things constant. This leads to a rise in the demand for housing. The empirical evidence supports this theory, as authors often find that the estimates of coefficients on measures of income in regressions have coefficients that are statistically and economically significant, with the expected positive sign. Econometricians rarely find a cointegrating relationship between house prices and income levels, however, which suggests that other variables might also influence house prices.

Another important variable commonly thought to influence house prices is the mortgage interest rate. One theoretical argument is that as the interest rate increases, the cost of repaying a mortgage increases, and if one argues that potential homeowners borrow an amount of money so that future repayments are a certain proportion of disposable income, then potential homeowners borrowing capacity falls, reducing demand for housing. Due to this strong theoretical link, a measure of the interest rate, either nominal or real, is usually included in house price models.

The empirical evidence is inconclusive on the strength of the effect of the mortgage interest rate on house prices. Some authors, such as Bourassa and Hendershott (2001), and Otto (2007), estimate coefficients on their interest rate terms that are economically and statistically significant, with the *a priori* expected sign. Others, such as Abelson et al. (2005), and Bodman & Crosby (2003), find that the estimated coefficient is either

statistically insignificant, economically insignificant, or both. In some cases, the coefficient even has the incorrect sign. Cameron et al. (2006), and McQuinn & O'Reilly (2008), argue that this may be due to the non-linear effect of interest rates on house prices. Consequently, Cameron et al. (2006) include an interaction of interest rates and an index for credit conditions, and McQuinn and O'Reilly (2008) include a non-linear term involving interest rates. Their empirical results support their argument. Thus, while economic theory suggests that interest rates play an important role in the determination of house prices, the empirical evidence is mixed, which may be a result of the manner in which the variable is included.

Measures of equity returns and the number of households are also often included in house price models, although not as often as the variables mentioned above. Equity returns are considered a substitute for investment in housing and an indicator of financial wealth, and are consequently thought to influence the demand for housing. A change in the number of households is also thought to influence the demand for housing. Consequently, the models of Abelson et al. (2005), Gallin (2006), and Otto (2007), incorporate a variable measuring equity prices, and the models of Bourassa & Hendershott (2001), Bodman & Crosby (2003), and Otto (2007) incorporate a variable measuring the number of households. Empirically, the sign and statistical significance of the estimated coefficients on these variables differ.

Expected capital gains are thought to play a role in the determination of the level of investment in the housing sector, and thus influence the demand for housing, but there are

no data on this variable, so proxies are usually employed. Abelson et al. (2005), and Otto (2007) employ variables measuring economic activity as proxies, since they are thought to affect the future growth rate of house prices, while Gallin (2006) takes the average percentage increase in house prices over the past three years as his proxy. Once again, the empirical estimates of the coefficients on these variables present mixed results.

Some variables are only appropriate in specific types of models. Measures of construction costs, for example, are appropriate in reduced form equations, but are inappropriate in inverse demand equations, as they influence the profits of construction companies, and thus the supply of housing. They are consequently included in the reduced form models of Bourassa & Hendershott (2001), Bodman & Crosby (2003), and Gallin (2006). Measures of the quantity of housing, on the other hand, may be included in inverse demand equations, such as in Abelson et al. (2005), and Otto (2007). Empirically, the sign and statistical significance of the estimated coefficients vary over models.

Some variables are rarely incorporated into house price models. Measures of credit conditions are one such variable. If lenders were to increase the amount they lend to an individual with certain characteristics, then this could have a similar effect to an increase in income, since more potential homeowners can gain access to the finance needed to purchase a house, and those who could already gain access to finance can access more, increasing the demand for housing. In their study of British house prices, Cameron et al. (2006) employ the credit conditions index estimated by Fernandez-Corugedo & Muellbauer (2006). They found that not only are credit conditions a determinant of house

prices, but omitting a variable that accounts for changes in credit conditions could lead to bias in estimates of the effect of interest rates on house prices. Despite this finding, measures of credit conditions are rarely included in house price models, as it is difficult to find appropriate data on this variable.

2.2 Dealing with non-stationarity

Many of the variables mentioned in Section 2.1 are non-stationary, or more specifically, integrated of order one (I(1)), and as a result, regressions involving these variables can be problematic. If the dependent variable or an independent variable is I(1), then there are a number of problems. As a simple univariate example that also applies to the multivariate case, let

$$y_t = \beta x_t + \varepsilon_t \quad (1)$$

If $y_t \approx I(1)$ or $x_t \approx I(1)$, then it is possible that:

$$\varepsilon_t = y_t - \beta x_t \approx I(1) \quad (2)$$

In this case, the t-statistic from an OLS regression of x_t on y_t will not have a limiting standard normal distribution, and the R^2 will converge in probability to a random variable. This can lead to spurious results, where there is a high R^2 , significant t and F statistics, and no true relationship between the variables. Thus, those who model house prices must take steps to ensure that their models do not suffer from this problem.

A number of authors adopt the method of estimating the models in first differences. The first difference of an I(1) variable is a stationary, I(0) variable, so estimating models using the variables in first differences eliminates the possibility of finding spurious

results due to the use of non-stationary variables. Authors employing this method of dealing with non-stationary variables include Bourassa & Hendershott (2001), and Otto (2007). Possible costs of adopting this method are that it does not provide information about the relationship between the variables in levels, and, if there is a cointegrating relationship between the variables, then including an error correction term will improve an equation in first differences. Consequently, some econometricians employ alternative methods of dealing with this type of problem.

A number of authors adopt cointegration techniques in estimating models of house prices. Using the specification in equation (1), and assuming that $y_t \approx I(1)$ and $x_t \approx I(1)$, it is possible that there exists a β such that:

$$\varepsilon_t = y_t - \beta x_t \approx I(0) \tag{3}$$

If this is the case, then there is a cointegrating relationship between y_t and x_t , with β being the cointegrating parameter. Many tests are available to assess whether there is a cointegrating relationship between the variables. One such test is the Augmented Engle Granger (AEG) test. This involves running an Augmented Dickey Fuller (ADF) test on the estimated residuals of a regression of y_t on x_t . The null hypothesis is that there is a unit root in the levels of the residual (i.e. no cointegration), and the alternative hypothesis is that the residual series is stationary. To test the hypothesis, the econometrician must use Engle Granger critical values. Gallin (2006) employs this test, as well as a number of other cointegration tests, and does not find evidence of a cointegrating relationship between house prices and a variety of independent variables in the U.S. data at the city level. Malpezzi (1999), Holly et al. (2006), McQuinn & O'Reilly (2008), and

Abelson et al. (2005), all find that there is evidence of cointegration between house prices and their respective independent variables, in the U.S. (Malpezzi and Holly et al.), Irish, and Australian data respectively. Thus, the evidence on whether there is a cointegrating relationship between house prices and their fundamentals is unclear in the literature, with the result appearing to depend on the data set and model specification.

If there is evidence of cointegration, then econometricians tend to estimate two equations. The first is a long run equilibrium equation, which looks similar to estimating equation (1). Inference on an Ordinary Least Squares (OLS) regression of (1) may still be invalid, however, as there is often correlation between the error term, ε_t in the example, and Δx_t , Δx_{t+s} , and Δx_{t-s} , where s is close to t . Consequently, leads and lags of Δx_t are often included in the regression. The estimate of β from this equation, sometimes referred to as the dynamic ordinary least squares (DOLS) estimator due to Stock & Watson (1993), is often employed, such as in Abelson et al. (2005) and McQuinn & O'Reilly (2008). This provides information on the long run effect of the levels of variables on the levels of house prices.

Econometricians frequently estimate a second equation, referred to as the short run error correction equation, which, in the univariate example, might look as follows:

$$\Delta y_t = \delta_0 + \delta_1(y_{t-1} - \hat{\beta}_{DOLS}x_{t-1}) + \sum_{j=0}^m \delta_{j+2}\Delta x_{t-j} + \sum_{k=1}^p \delta_{k+m+2}\Delta y_{t-k} + \nu_t \quad (4)$$

Malpezzi (1999), Abelson et al. (2005), and McQuinn & O'Reilly (2008) estimate this equation. The coefficient on the error correction term, δ_1 , is referred to as the speed of

adjustment parameter, as it is the speed at which the variables adjust back to the long run equilibrium, once they are out of equilibrium. Those that find evidence of cointegration, and estimate long run equilibrium and error correction equations, find that adjustment to equilibrium is relatively slow. McQuinn & O'Reilly (2008), for example, estimate it at approximately 0.05, i.e. 5% error correction per quarter, while Malpezzi (1999) estimates it at 2.7% per year. Thus, estimating this equation can provide information on the short run movements of cointegrated variables. It is not estimated in this paper, however, due to time constraints, and because the objective of the paper is to assess whether there is evidence of cointegration, and if possible to find information on the long run relationship between the variables.

2.3 Accounting for bubbles

There is no consensus on the manner in which one should account for house price bubbles, and as a result, authors have attempted to do so in a variety of ways. Bourassa & Hendershott (2001), and Bodman & Crosby (2003) adopt a similar technique. They employ two terms, a one period lag of the rate at which house prices appreciate, and a one period lag of the difference between the equilibrium and actual real price levels. To obtain a measure of the equilibrium price level, however, they employ the fitted values of a regression without the two terms as a measure of the equilibrium appreciation rate, and allow a house price level that they assume is in equilibrium to appreciate at this rate. They subsequently include the bubble terms, altering the estimates, so they calculate new fitted values, and re-estimate the equations. They continue this process until the estimates stabilise.

Abelson et al. (2005) adopt an alternative technique to account for bubbles. They argue that bubbles in house prices will lead to a faster adjustment back to equilibrium. The authors define a boom observation as an observation where real house price growth over the past year has been greater than two percent. In their error correction equation, they allow for a variation in the speed of adjustment parameter between boom and non-boom years. Empirically, however, they find that the difference in the speed of adjustment is relatively small.

Some authors do not account for bubbles at all, as there are flaws with the methods outlined above. These authors instead choose to analyse the differences between the actual and predicted values, suggesting that when the difference is both positive and large, there may be a bubble, as fundamentals are not determining the actual price. This method, as with the others, has flaws. Consequently, there is no consensus on how to account for bubbles in house price models. I do not attempt to account for them.

2.4 House price models with variation across space

While many authors model house prices at the national level, some allow for variation in the way house prices are determined in different areas. The advantage of this is that if there is variation in the manner in which house prices are determined in different areas, then allowing for this variation might provide a more accurate and informative estimation of the way house prices are determined. The disadvantage to estimating models that vary across space is that it is more difficult to obtain data on variables for specific areas within

a nation, which can result in fewer observations, and the need to use data for a larger area. Despite the disadvantages, Bodman & Crosby (2003), Cameron et al. (2006), Holly et al. (2006), Gallin (2006), and Otto (2007) estimate models that allow for variation at some level smaller than the national level. The majority of these papers find that there is variation in the way house prices are determined beneath the national level. Thus, in assessing whether there is evidence of a cointegrating relationship between house prices and fundamentals, I allow for differences in the way house prices are determined across different capital cities by estimating regressions for each capital city.

A number of authors estimate regional house price models that require the author to account for the relationship between regions. Cameron et al. (2006), and Holly et al. (2006) allow for variation across regions in the U.K. and U.S. respectively by estimating models for a number of different regions. Both account for the spatial interactions and interdependencies, particularly between contiguous regions. Bodman & Crosby (2003) and Otto (2007), in estimating models for the State capital cities in Australia, do not incorporate methods to account for these interactions, apart from a lag of Sydney house prices in the models for other cities in Otto (2007). Accounting for spatial interactions and interdependencies may be unnecessary when modelling Australian capital cities, since there is a vast distance between the cities. Thus, I do not account for these interactions in this paper.

3. The Theoretical Model

The theoretical model that I employ as the basis of my empirical equation is an adaptation of the model in McQuinn and O'Reilly (2008). The foundation of the model is the idea that most potential homeowners finance the purchase of their house with a mortgage. These potential homeowners are therefore financially constrained by the amount that they can borrow from lenders, given their income. Since homeowners purchase housing with a mortgage, the assumption is that other variables, such as financial wealth, do not play a role in the demand for housing. Consequently, the demand for housing is a function of the amount that the potential homeowner can borrow.

Lenders in the model aim to have borrowers repay some proportion of their income every period, for the term of the mortgage. Thus, one is able to calculate the amount that potential homeowners can borrow according to the following equation,

$$B_{it} = \kappa_t Y_{it} \left(\frac{1 - (1 + R_t)^{-\tau}}{R_t} \right) \quad (5)$$

where B_{it} is the amount that can be borrowed in capital city i at time t , κ_t is the proportion of disposable income that is spent on mortgage repayments at time t , Y_{it} is disposable income per household in capital city i at time t , R_t is the real mortgage interest rate at time t , and τ is the length of the mortgage. I implicitly assume in the equation that the manner in which lenders behave does not differ across capital cities, but does differ over time, and that the term of the mortgage is constant over both time and capital cities.

Due to the assumption that the demand for housing is primarily a function of the amount that the individual can borrow, the demand equation looks as follows:

$$P_{it}^D = \left[\kappa_t Y_{it} \left(\frac{1 - (1 + R_t)^{-\tau}}{R_t} \right) \right] (Q_{it}^D)^{-\mu_i} \quad (6)$$

where P_{it}^D is the demand price for housing, Q_{it}^D is the quantity of housing demanded, and μ_i is the negative inverse of the own price elasticity of the demand for housing, which I assume varies across different cities. The supply of housing is inelastic in the short run, so the inverted housing supply equation is:

$$P_{it}^S = \delta_i (Q_{it}^S)^{\phi_i} \quad (7)$$

where P_{it}^S is the supply price of housing, δ_i is a constant, Q_{it}^S is the quantity of housing supplied, and ϕ_i is the inverse of the own price elasticity of supply of housing. In equilibrium, $P_{it}^S = P_{it}^D$. Setting these to be equal, and solving for the equilibrium quantity, Q_{it}^E , gives

$$Q_{it}^E = \left[\frac{\kappa_t Y_{it} \left(\frac{1 - (1 + R_t)^{-\tau}}{R_t} \right)}{\delta_i} \right]^{\frac{1}{(\phi_i + \mu_i)}} \quad (8)$$

Substituting this quantity into the inverse supply equation, and solving for the equilibrium price, P_{it}^E , gives

$$P_{it}^E = \delta_i^{\frac{\mu_i}{(\phi_i + \mu_i)}} \kappa_t^{\frac{\phi_i}{(\phi_i + \mu_i)}} Y_{it}^{\frac{\phi_i}{(\phi_i + \mu_i)}} \left(\frac{1 - (1 + R_t)^{-\tau}}{R_t} \right)^{\frac{\phi_i}{(\phi_i + \mu_i)}} \quad (9)$$

Taking the logs of these values,

$$\begin{aligned}\log(P_{it}^E) = & \frac{\mu_i}{(\phi_i + \mu_i)} \log(\delta_i) + \frac{\phi_i}{(\phi_i + \mu_i)} \log(\kappa_t) \\ & + \frac{\phi_i}{(\phi_i + \mu_i)} \log(Y_{it}) + \frac{\phi_i}{(\phi_i + \mu_i)} \log\left(\frac{1 - (1 + R_t)^{-\tau}}{R_t}\right)\end{aligned}\quad (10)$$

In order to estimate (10), I impose a number of restrictions. Firstly, I restrict the value of the coefficient on $\log(Y_{it})$ to one, and re-arrange to get the log of the house price to income ratio as the dependent variable,

$$\log\left(\frac{P}{Y}\right)_{it} = \frac{\mu_i}{(\phi_i + \mu_i)} \log(\delta_i) + \frac{\phi_i}{(\phi_i + \mu_i)} \log(\kappa_t) + \frac{\phi_i}{(\phi_i + \mu_i)} \log\left[\frac{1 - (1 + R_t)^{-\tau}}{R_t}\right] \quad (11)$$

Imposing this restriction clarifies a key implication in the model. House price to income ratios are often looked at as indicators of whether there is a bubble in housing prices. This is based on the idea that a rise in the house price to income ratio is not sustainable in the long term, so house price to income ratios will fall back to some long-term level. This implies that there should be a stable long run relationship between house prices and income levels. It is often found, however, that house prices and incomes are not cointegrated. This rearrangement of the model clarifies the idea that, according to this model, real interest rates might fall, or lenders might decide to lend more to individuals as a proportion of their income, increasing potential homeowners' ability to obtain funds to purchase housing, which can lead to a permanent rise in house prices that is independent of income.

To illustrate the second restriction imposed on the model, I first rearrange the equation,

$$\log\left(\frac{P}{Y}\right)_{it} = \frac{\mu_i}{(\phi_i + \mu_i)} \log(\delta_i) + \frac{\phi_i}{(\phi_i + \mu_i)} \log(\kappa_t) + \frac{\phi_i}{(\phi_i + \mu_i)} \log[1 - (1 + R_t)^{-\tau}] - \frac{\phi_i}{(\phi_i + \mu_i)} \log(R_t) \quad (12)$$

The second restriction involves setting the coefficient on $\log[1 - (1 + R_t)^{-\tau}]$ to zero. To include this variable, one must estimate the constant τ . Without data on the variable, this amounts to guesswork. McQuinn and O'Reilly (2008) calculate their variable with τ estimated as 15, 20, 25 and 30 years, in four different regressions, and find that altering it does not have a strong effect on their results. Consequently, the primary model restricts this coefficient to zero, and an alternative model incorporates the variable with τ equal to 80 quarters.

This leads to the empirical equation,

$$\log\left(\frac{P}{Y}\right)_{it} = \beta_{0i} + \beta_{1i} \log(\kappa_t) + \beta_{2i} \log(R_{it}) + \varepsilon_{it} \quad (13)$$

where ε_{it} is the error term. The coefficients vary over each State capital city, so essentially this involves nine different regressions, one for each State capital city, and one at the national level. There is a subscript added to the real interest rate, as the measure of the real interest rate varies between cities.

Some might argue that there is an endogeneity problem when estimating a model with a measure of the proportion of disposable income that goes towards mortgage repayments included as a determinant of house prices. This argument suggests that as house prices rise, individuals must borrow more as a proportion of income to pay for a house.

According to the theoretical model, however, lenders decide how much to lend potential homeowners based on the proportion of their income that they think the homeowner will be able to repay. This decision is made independently of house prices and the house price to income ratio, since there is a one to one relationship between house prices and incomes. Thus, according to the theoretical model, a rise in house prices does not affect lenders decisions, so there is no endogeneity problem.

4. The Data

4.1 Data construction

To create a series for real median house prices for the period 1986:2 to 2007:4, I employ a number of series, all of which are available from the Australian Bureau of Statistics (ABS) website. Median house price series for each capital city are available for the period 2002:1 to 2006:4. Creating observations for the median house price for 1986:2 to 2001:4 and 2007:1 to 2007:4 is possible using price indices for each capital city. Taking a weighted average of the capital city median house prices, with the weights being the number of working age people in the capital city divided by the number of working age people in the sum of the capital cities, provides an estimate of the median house price series for the country. The series for the number of working age people in each capital city is available for the entire period (i.e., 1986:2 to 2007:4). I multiply median house prices by 1000, to convert them to dollar terms. I employ the consumer price index excluding housing, which is available for each capital city, and nationally, over the entire period, to convert these dollar values to observations for real median house prices. I adjust all measures of the Consumer Price Index (CPI) to exclude the effect of the Goods and Services Tax (GST). In all conversions from nominal to real, the base period is 1989-90, for which the CPI is 100.

To obtain a measure of real median house price to real income ratios for each city, and one for the national level, I employ seasonally adjusted average weekly earnings series. The ABS website contains quarterly observations on this variable for each State, and for

Australia, over the entire period. I multiply these observations by 52 to convert them to a measure of average annual earnings per capita. I convert these values into real terms using CPI's, which are available for each city, and nationally, for the entire period. The real median house price to real income ratio for each State capital city is the real median house price in each city divided by real average annual earnings per capita in the relevant State.

For the nominal interest rate, monthly observations on the standard variable interest rate provided by banks on housing loans are available from the Reserve Bank of Australia (RBA) website for the entire period. I take the average over the quarter to obtain a quarterly measure. This does not vary by city or State. I employ the CPI's for each city again to convert nominal interest rates to real interest rates. This provides a measure of the real interest rate that varies by city.

An ideal measure of κ_t is the proportion of disposable income that the lenders believe will go towards interest repayments on housing loans over the length of the mortgage, a ratio between zero and one. Closely related to this is the 'Interest payments to disposable income – Housing' ratio, which is available on the RBA website for the entire period. It is the ratio of total interest payments made on housing loans to annual disposable income. This variable is only available at the national level. Consequently, I employ this national level variable as a measure of κ_t for all the capital cities.

4.2 Summary statistics

Table 1 – Summary Statistics

Area	Mean	Minimum	Maximum	ADF statistic on levels	ADF statistic on difference
House price to income ratio					
Syd	8.13	4.82	12.25	-1.874	-3.337
Mel	5.45	4.01	9.18	-0.025	-3.265
Bri	5.59	3.47	9.57	-1.008	-3.119
Ade	5.11	3.87	8.31	-0.06	-3.182
Per	5.5	3.66	10.03	-0.724	-3.842
Hob	4.44	3.58	7.56	-0.165	-3.046
Dar	5.31	3.46	9.37	-0.036	-3.228
Can	5.81	4.37	8.55	-0.918	-3.48
Aus	6.38	4.33	9.83	-0.185	-3.523
Housing interest payments to disposable income ratio					
Aus	5.88	4	10.4	-0.068	-3.319
Quarterly real interest rate (%)					
Syd	1.51	0.67	2.64	-1.291	-5.314
Mel	1.54	0.73	2.56	-1.046	-6.323
Bri	1.53	0.68	2.63	-0.924	-5.538
Ade	1.53	0.38	2.54	-1.055	-6.897
Per	1.52	0.68	2.97	-1.152	-6.413
Hob	1.55	0.74	2.58	-1.202	-5.631
Dar	1.59	0.7	2.74	-1.233	-5.256
Can	1.55	0.66	2.55	-1.052	-6.833
Aus	1.53	0.72	2.60	-1.073	-5.712

ADF statistic refers to the Augmented Dickey Fuller test statistic. The 5% critical values for the Dickey-Fuller test range from -2.904 to -2.907.

The first three columns of Table 1 report means, minimums, and maximums of the variables in levels. There are 87 quarterly observations on all of the variables, from 1986:2 to 2007:4. For the house price to income ratios, the means, minimums, and maximums indicate that there is a level difference between the various cities. The house price to income ratio with the highest level is in Sydney, while the lowest is in Hobart. The housing interest payments to disposable income ratio is not restricted to the zero one interval, as it is total housing interest repayments to annual disposable income. It ranges from 4 to 10.4. The means, minimums, and maximums for the quarterly real interest rates indicate that there is not much variation between the cities, which is due to the similarity between the CPI's for the various cities.

The last two columns of Table 1 report Augmented Dickey Fuller statistics, which are for testing whether the variables have a unit root. I calculate Augmented Dickey Fuller tests statistics on the logarithms of the house price to income ratios, as well as the levels of the housing interest payments to disposable income ratio, quarterly real interest rates, and the differences of these variables. The null hypothesis is that there is a unit root in the variable, while the alternative is that the variable is stationary.

I make a number of decisions before obtaining the test statistics. Firstly, I choose the number of lags of the dependent variable to include as independent variables by including a certain number of lags, running the Augmented Dickey Fuller regression, obtaining the residuals, and testing for serial correlation up to order four. I choose the minimum

amount of lags required to eliminate serial correlation. The number of lags in the regressions ranges from three to six. I also choose not to include a trend term in the regressions. Thus, one obtains the test statistics above, for which the 5% critical values range from -2.904 to -2.907, depending on the number of observations, which depends on the number of lags.

The one exception to the decisions above is for the test on the level of the housing interest payments to disposable income ratio, where there is a trend term and ten lags, and the 5% critical value is -3.474. The regression to obtain the statistic on the difference of the variable, however, includes no trend and four lags.

The test statistics for the levels of the variables, which range from -0.025 to -1.874, are greater than their critical values. Consequently, the null hypothesis of a unit root in the levels of the series is not rejected. Adding a trend term to the Augmented Dickey Fuller equations that produce these statistics does not change the conclusions. The test statistics for the differences of the variables, which range from -3.046 to -6.897, are all less than the critical values. Thus, the conclusions from the tests are that all the variables have a unit root in the levels, but are stationary in the differences, suggesting they are I(1) series.

4.3 Movements in the variables

Plots of the variables over time provide an insight into the movements of the variables, and assist with the understanding of the relationships underlying the regressions in Section 5. As there are 19 variables employed, the charts included for the real median house price to real income ratio and real interest rate are for Sydney only, with

differences between Sydney and the other cities noted. For the real median house price to real income ratios, the charts for the other capital cities, and for Australia, are included in the Appendix as Figures A1 to A4.

Figure 1 plots the real median house price to real income ratio for Sydney from 1986:2 to 2007:4. The house price to income ratio in Sydney rises 52% from 1986:2 to 1989:4. The Melbourne, Adelaide, and Australian ratios also rise in this early period, although their rise is less pronounced. The ratios in the rest of the cities do not change much over this period. Throughout the 1990's, the ratios are relatively stable throughout the country, although the slight movements vary between cities. In the Sydney series, as in the Melbourne, Adelaide, and Australian series, the ratio declines over the first half of the 1990's, and rises over the second half, while the reverse occurs in Brisbane, Hobart and Canberra. In Darwin, there is a rapid rise in the first half of the 1990's, before a slow decline towards the end of the decade, while the ratio for Perth is steady over the whole decade. In the early 2000's, there is a rise in the house price to income ratio in every capital city in Australia, with the ratios in cities such as Brisbane rising by as much as 74.9% from 2001:1 to 2004:1, while in cities such as Darwin, the rise is only 13% over the same period. This illustrates that the timing of the rise varies by city. In the mid 2000's, the house price to income ratio ceases to rise in every city except Darwin, with some, such as Sydney, declining towards the end of the period, while others flatten before rising again.

Figure 1 – Sydney House Price to Income Ratio

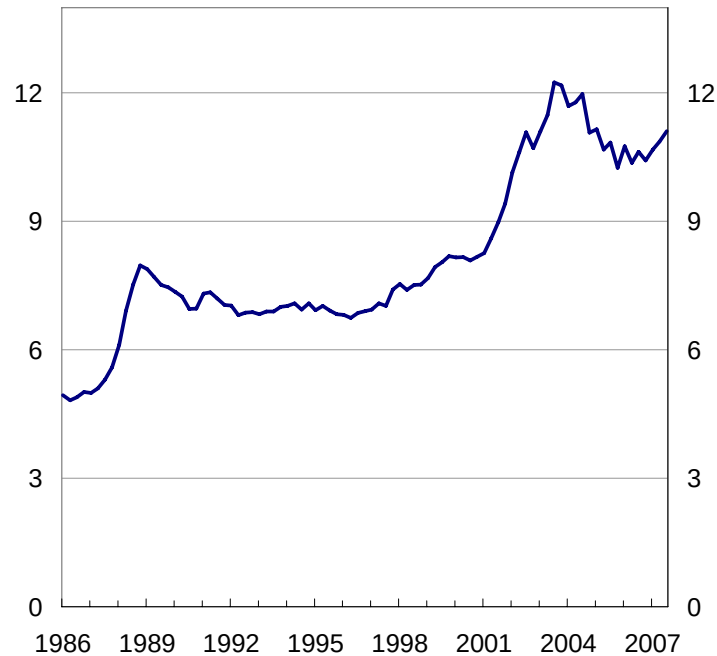


Figure 2 is a plot of the quarterly real interest rate in Sydney from 1986:2 to 2007:4. This chart indicates that there is a sharp rise in quarterly real interest rates from the late 1980's to early 1990's, from 1.5% per quarter in 1989:2 to 2.6% per quarter in 1991:4 in Sydney. For the majority of the period from 1992 through to 2001, real interest rates decline, falling to a low of 0.67% in Sydney. Following this, they begin to rise again. Movements in the quarterly real interest rate in the other State capital cities are similar to this.

Figure 2 – Quarterly Real Interest Rate in Sydney

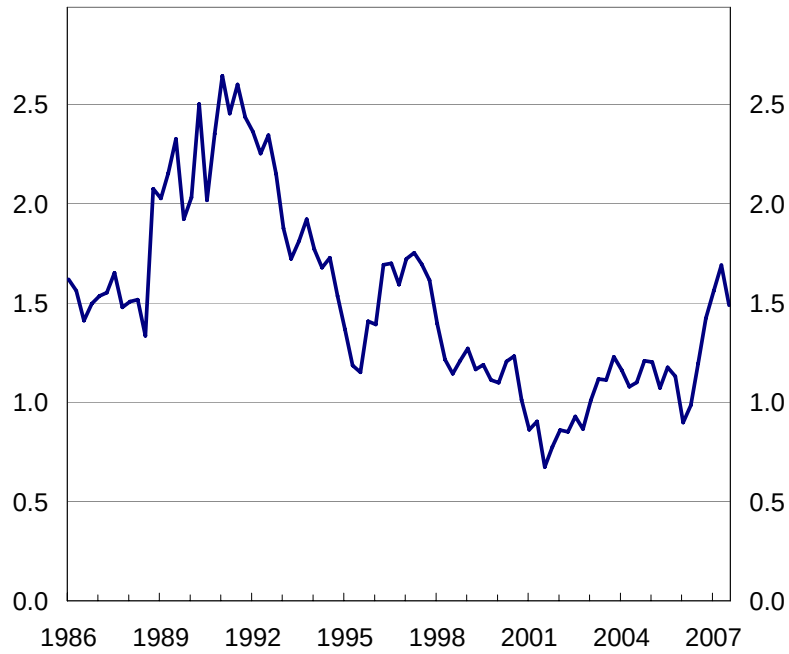


Figure 3 - Housing Interest Payments to Disposable Income Ratio

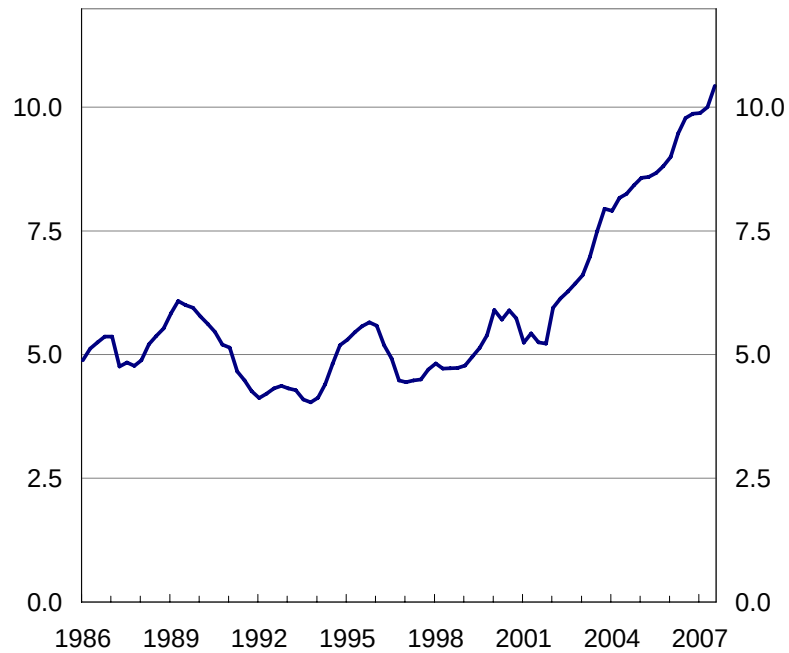


Figure 3 is a plot of the housing interest payments to disposable income ratio for Australia. This plot indicates that the ratio fluctuates around five before 2002:1. From

2002:1 onwards, the ratio rises. Over the period 2001:1 to 2004:1, there is a rise of 38.7% in this ratio, which is similar to the average rise in the house price to income ratios of 48.9% over this same period. When the house price to income ratios decline towards the end of the period, the amount of housing interest repayments continues to rise, which may be related to rising real interest rates, rather than an expansion of credit.

5. Empirical Results

5.1 Assessing whether there is evidence of cointegration

This section presents and analyses the results from various estimations based on equation (13), focusing specifically on whether there is evidence of a cointegrating relationship between the logarithm of real median house price to real income ratios, real interest rates and housing interest payments to disposable income ratios in the Australian data at the capital city level. In order to assess whether there is evidence of cointegration, I run ordinary least squares (OLS) regressions for each capital city and Australia. When there is evidence of cointegration, these estimates will be consistent, but the standard errors will be invalid. Despite this, these regressions allow the calculation of Augmented Engle Granger (AEG) statistics, which I employ to test the null hypothesis of no cointegrating relationship between the variables against the alternative of a cointegrating relationship. Table 2 reports the coefficient estimates and the AEG statistics.

To calculate AEG statistics, one must choose the number of lags of the dependent variable to include as independent variables, in order to ensure that the tests do not suffer from serial correlation problems. To choose the number of lags, I first run the Augmented Engle Granger regression of the difference of the residuals from the OLS regression on a one period lag of the levels of the error and a one period lag of the difference of the error. I take the residuals from this regression and test for serial correlation of order four. If this regression suggests that there is serial correlation in the Augmented Engle Granger equation, I add a two period lag of the difference of the OLS residual to the Augmented

Engle Granger equation and re-estimate it, until it suggests there is no serial correlation. This method leads to the use of two to six lags in the Augmented Engle Granger equations that produce the test statistics in Table 2.

Table 2 – OLS Regressions for Augmented Engle Granger Tests

Variable	Syd	Mel	Bri	Ade	Per	Hob	Dar	Can	Aus
R	-.1070	-.1546	-.0319	-.0104	.0019	-.0030	-.2829	-.0243	-.0901
κ	.0962	.1132	.1164	.1170	.1328	.1252	.0689	.0842	.1070
β_0	1.664	1.236	1.052	.9346	.8902	.7309	1.681	1.284	1.335
AEG	-3.913	-3.290	-2.865	-2.920	-3.969	-2.455	-1.944	-2.632	-3.737

The dependent variable is the log of the real median house price to real average income ratio. R refers to the real interest rate, κ refers to the housing interest payments to disposable income ratio, and β_0 refers to the constant. AEG refers to the Augmented Engle Granger statistic. The 10% critical value for an Augmented Engle Granger test with three variables in the system and 100 observations is -3.59 (Engle and Yoo, 1987). There are 87 observations.

If the test statistic is lower than the 10% critical value, I reject the null hypothesis of no cointegrating relationship between the variables and conclude in favour of cointegration. Using the 10% critical value for a regression with 100 observations from Engle and Yoo (1987), - 3.59, as an approximate critical value, there is sufficient evidence to reject the null hypothesis of no cointegration at the 10% level for the Sydney, Perth, and Australian regressions. For the Melbourne regression, there is insufficient evidence to reject the null hypothesis of no cointegrating relationship at the 10% level. The test statistic, however, is close to the critical value of -3.59, and the coefficient estimates are similar to the estimates for Sydney, which supports an argument that there may be a cointegrating

relationship between the variables. There is insufficient evidence to reject the null hypothesis of no cointegrating relationship in the Brisbane, Adelaide, Hobart, Darwin, and Canberra regressions. Thus, according to these results, there is some evidence of a cointegrating relationship between the logarithm of the real median house price to real income ratio, the real interest rate, and the housing interest payments to disposable income ratio at the capital city level, but not for all State capital cities.

The OLS estimates of the coefficients are consistent when there is a cointegrating relationship between the variables, so they can provide us with some insights. Apart from the estimate of the coefficient on the real interest rate in the Perth regression, the signs on these estimates coincide with *a priori* expectations. The estimates of the coefficients on the housing interest payments to disposable income ratio are positive, suggesting an increase in this ratio leads to a rise in the real median house price to real income ratio. These estimates, in the regressions for Sydney, Melbourne, Perth, and Australia, all have similar magnitudes, suggesting that a one-unit increase in κ_t leads to an increase in the real median house price to real income ratio of 9.62 to 13.28 percent. In the regressions for the other cities, where there is less confidence in the estimates due to the lack of evidence of cointegration, the effect of a one-unit increase in κ_t is between 6.89 percent and 12.52 percent. In accordance with *a priori* expectations, the estimates of the coefficients on the real interest are all negative, except for the estimate of the coefficient in the Perth regression, which is positive, but economically insignificant. The estimates suggest that a 100 basis point rise in the real interest rate leads to a decrease in the real median house price to real income ratio between 9 and 16 percent in Sydney, Melbourne,

and Australia, whereas it leads to a 0.19 percent increase in this ratio in Perth. In the other State capital cities, for which there is less confidence in the estimates, the estimated effect of a 100 basis point rise in the quarterly real interest rate ranges from a 0.3% decrease to a 28.3% decrease. Performing statistical inference on these estimates is not possible, as the standard errors are invalid. Consequently, in the next section I calculate the dynamic ordinary least squares (DOLS) estimates, which, when the variables are cointegrated, are consistent and efficient, with asymptotically valid standard errors.

5.2 Dynamic ordinary least squares (DOLS) estimation

As described in Section 2.2, DOLS estimation involves adding leads, lags, and contemporaneous differences of the independent variables to the regression. The specification for the regressions for each city becomes

$$\log\left(\frac{P}{Y}\right)_t = \beta_0 + \beta_1 \kappa_t + \beta_2 R_t + \beta_3 \Delta \kappa_t + \beta_4 \Delta R_t + \sum_{j=1}^J [\beta_{j+4} \Delta \kappa_{t+j} + \beta_{j+6} \Delta \kappa_{t-j} + \beta_{j+8} \Delta R_{t+j} + \beta_{j+10} \Delta R_{t-j}] + \varepsilon_t \quad (14)$$

This model requires a decision on the number of lags to employ. I vary the number of leads and lags, J , between one and four, but the results are similar, so I only report the results for the two extremes. Additionally, since the regressions reveal evidence of serial correlation, I employ Newey and West (1987) standard errors, setting the maximum lag to consider in the autocorrelation structure to eight. Tables 3 and 4 report the results, with the standard errors in brackets.

Table 3 – DOLS Regression with One Lead and Lag

Variable	Syd	Mel	Bri	Ade	Per	Hob	Dar	Can	Aus
R	-.0888 (.0727)	-.1633* (.0850)	.0074 (.0807)	.0038 (.0526)	.0356 (.0327)	.0005 (.0644)	-.3217** (.0739)	.0026 (.0679)	-.0710 (.0610)
κ	.0895** (.0188)	.1102** (.0126)	.1233** (.0164)	.1201** (.0129)	.1400** (.0163)	.1317** (.0154)	.0569** (.0233)	.0837** (.0152)	.1059** (.0125)
β_0	1.674** (.2103)	1.268** (.2009)	.9501** (.1714)	.8898** (.1397)	.7979** (.1510)	.6909** (.1625)	1.813** (.2585)	1.238** (.1671)	1.311** (.1616)
AEG	-3.336	-3.269	-2.903	-3.091	-4.376	-2.341	-2.137	-2.807	-3.443

Notes for both Table 3 and 4 are beneath Table 4.

Table 4 – DOLS Regression with Four Leads and Lags

Variable	Syd	Mel	Bri	Ade	Per	Hob	Dar	Can	Aus
R	-.0976 (.0680)	-.1929* (.0983)	.1136 (.1226)	.0626 (.0842)	.0533 (.0352)	.0550 (.1022)	-.3999** (.1474)	.0721 (.0965)	-.0538 (.0583)
κ	.0957** (.0230)	.1140** (.0234)	.1411** (.0230)	.1369** (.0189)	.1455** (.0175)	.1511** (.0201)	.0364 (.0287)	.0948** (.0219)	.1151** (.0191)
β_0	1.667** (.2082)	1.304** (.2707)	.6654** (.3043)	.6946** (.2153)	.7471** (.1304)	.4909* (.2472)	2.0672** (.3969)	1.055** (.2295)	1.237** (.1826)
AEG	-2.074	-2.120	-2.589	-2.783	-4.869	-2.533	-1.719	-2.595	-2.629

The dependent variable is the log of the real median house price to real average yearly income ratio. R refers to the real interest rate. κ refers to the housing interest payments to disposable income ratio. β_0 and refers to the constant. * indicates statistical significance at the 10% level. ** indicates statistical significance at the 5% level. The 10% critical value for the Augmented Engle Granger test is approximately 3.59 (Engle and Yoo, 1987). Parentheses indicate Newey and West (1987) standard errors.

An analysis of the estimates can provide insights into both the determinants of house prices in Australian capital cities and the validity of the model. The signs of the estimates

of the coefficients on the housing interest payments to disposable income ratio (κ_t) are all positive, which is consistent with the expectations of these coefficients. For the Sydney, Melbourne, Perth and Australian regressions, the magnitudes of these estimates range from 0.0895 to 0.1455. The estimates of the coefficients on κ_t in the regressions for the other cities, for which there is less confidence in the estimates, are similar, ranging from 0.0837 to 0.1511, excluding Darwin, for which the estimates are lower. All of these estimates, excluding the estimate for Darwin in Table 4, are statistically significant at the 5% level. Interpreting these estimates, a one-unit increase in the housing interest payments to disposable income ratio in the long run leads to a rise in the real median house price to real income ratio of 8.37 to 15.11 percent. Thus, the DOLS regressions suggest that housing interest payments to disposable income ratios have a considerable impact on real median house price to real income ratios in Australian capital cities.

The magnitudes of the DOLS estimates of the coefficients on R_{it} , the real interest rates, display substantial differences across cities. The negative signs on the coefficient estimates in the regressions for Sydney, Darwin, and Australia are consistent with expectations. Interpreting the coefficient on R in the Sydney regression in Table 3, an increase in the real interest rate in the long run of 100 basis points leads to an 8.88% decrease in the real median house price to real income ratio. The estimates for Darwin suggest that real interest rates have a much stronger impact on the real median house price to real income ratio. The estimates of the coefficients for the other capital cities, however, have a positive sign, suggesting that a rise in the real interest rate leads to a rise in the real median house price to real income ratio. This result is inconsistent with

economic theory, as one would expect negative coefficients on all of the estimates. This may be suggesting that an important variable has been omitted from the model. The statistical significance of the estimates of the coefficients on R_{it} varies by city. The majority of the estimates of the coefficients on R_{it} , 14 out of 18, are statistically insignificant at the 10% level.

A comparison of the DOLS estimates in Table 3 and 4 with the OLS estimates in Table 2 reveals that including leads and lags of the differenced independent variables reduces the support for the model specification in certain ways, while adding to it in other ways. The Augmented Engle Granger statistics for the DOLS regressions differ to the statistics from the OLS regressions, with some, such as Perth, increasing, while others, such as Sydney, decrease. Based on the statistics from the DOLS regressions, the null hypothesis of no cointegrating relationship can only be rejected at the 10% level for Perth, although a number of the test statistics are close to this level when J , the number of leads and lags of the differenced independent variables, is one. Despite this, I maintain the conclusion that there is evidence of cointegration in the regressions for Sydney, Perth, Australia, and possibly Melbourne, based on the OLS results. In terms of the coefficient estimates, the OLS estimates of the coefficient on R_{it} , in the regressions for all capital cities excluding Perth, are negative, although some are economically insignificant. When incorporating the leads and lags of the differenced independent variables, the estimates of a number of these coefficients are close to zero, but the signs are positive, which differs from *a priori* expectations. Estimating the model with DOLS does, however, reveal that a number of these estimates are statistically insignificant. Additionally, the DOLS and OLS estimates

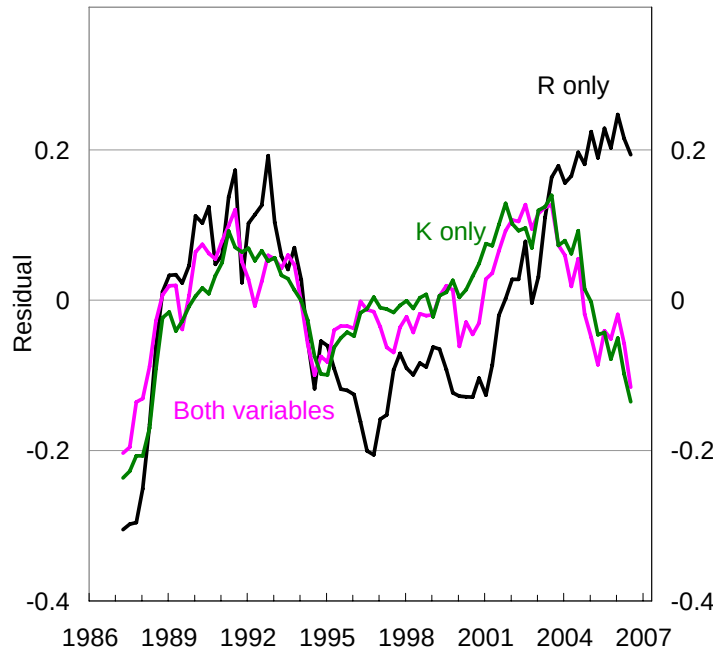
of the coefficients on κ_t all have the expected positive sign, and the magnitudes are similar, which supports the inclusion of this variable in the model. Thus, there are both positive features and negative features of the DOLS estimates, relative to the OLS estimates.

5.3 The boom in house price to income ratios

Given that there is some evidence of cointegration, the estimates from the regressions provide some insights into the determination of house prices in some Australian capital cities over the sample period. One feature of these estimates is that the rise in the total housing interest payments to annual disposable income ratio drives the increase in real median house price to real income ratios in the early 2000's.

Charts of the error terms of various regressions depict this result. Figure 4 plots the regression residuals for three regressions using data at the Australian level. The black line, titled Both Variables, is a plot of the residuals from the regression in equation (14) for Australia. The other two lines are plots for similar residuals, except that the regression that produces the residuals for the green line, titled K only, does not include the real interest rate terms from equation (14), while the regression that produces the residuals for the pink line, titled R only, does not include the housing interest payments to disposable interest rate terms.

Figure 4 – Regression Residuals



This chart indicates that the rise in the housing interest payments to disposable income ratio accounts for the increase in the real median house price to real income ratio in the early 2000's. In this period, the residual series titled R only rise more than the residuals from the other two regressions, which decline towards zero in a similar fashion, falling beneath zero in the late 2000's. This indicates that, according to the regression estimates, the housing interest payments to disposable income ratio accounts for the real median house price to real income ratio rise in Australia in the early 2000's, with similar results for Sydney, Perth, and the other Australian capital cities, excluding Darwin. In the context of the theoretical model, this result suggests that increases in lending to potential homeowners with a certain income in the early 2000's increased the capacity of individuals to purchase housing, increasing the demand for housing and thus the real median house price to real income ratio.

5.4 Structural breaks

If there is a structural break in the relationship, the error term from an OLS regression may appear to be non-stationary due to the break, which may lead the Augmented Engle Granger test to fail to reject the null hypothesis of no cointegration (Gabriel et al., 2003). Consequently, in this section I test whether there is evidence of a structural break, and upon finding evidence of this, I assess whether there is evidence of cointegration when allowing for a structural break.

Scatter plots of the dependent variable and real interest rates suggest that there may be a structural break. Figure 5 is a scatter plot of these data series for Perth, over the entire period. The trendline for the scatter plot over the entire period is downward sloping, suggesting that a rise in the real interest rate is associated with a reduction in the real median house price to real income ratio, as one would expect. Figure 6 is the same plot but restricted to the period before 2001:4. The trendline in the scatter plot for this period is almost flat, suggesting that a rise in the real interest rate has almost no effect on the real median house price to real income ratio. Thus, the effect of the real interest rate on the real median house price to real income ratio does not appear to be constant over the entire sample.

Figure 5 – Perth 1986:2 to 2007:4

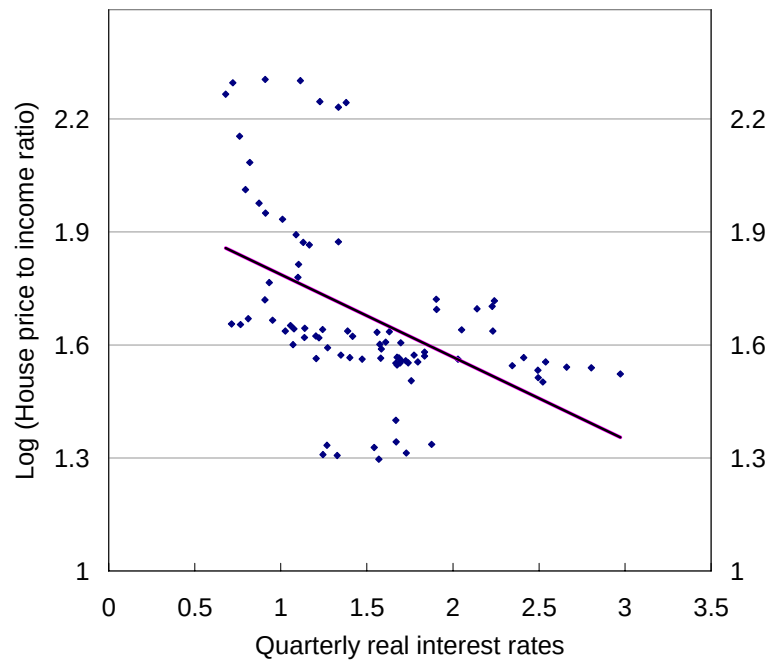
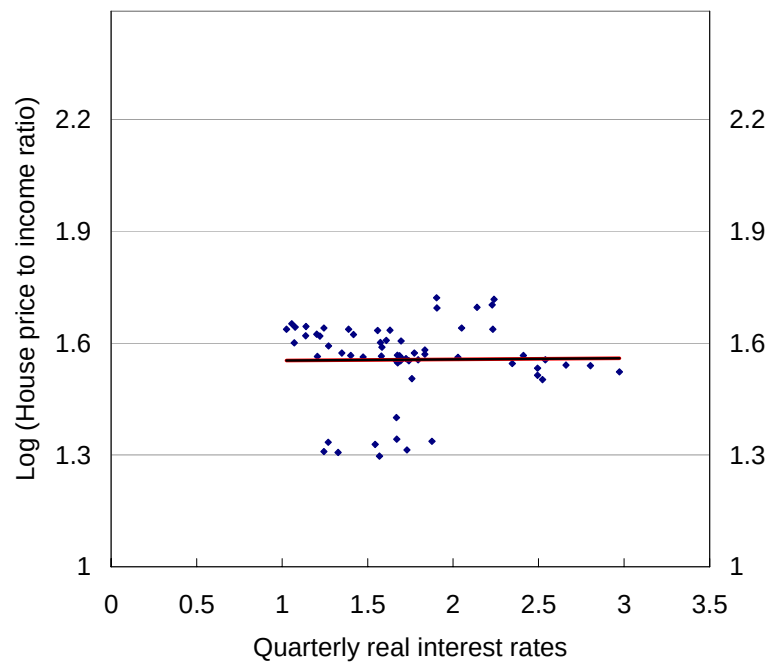


Figure 6 –Perth 1986:2 to 2001:3



Consequently, I calculate sup-F test statistics from Hansen (1992) to test for a structural break in the relationship. These statistics require a sequence of regressions based on (14) for each city. These regressions are of the form

$$\log\left(\frac{P}{Y}\right)_t = \beta_0 + \beta_1 \kappa_t + \beta_2 R_t + \beta_3 D_t + \beta_4 (\kappa_t \times D_t) + \beta_5 (R_t \times D_t) + \beta_6 \Delta \kappa_t + \beta_7 \Delta R_t + \sum_{j=1}^2 [\beta_{j+6} \Delta \kappa_{t+j} + \beta_{j+8} \Delta \kappa_{t-j} + \beta_{j+10} \Delta R_{t+j} + \beta_{j+12} \Delta R_{t-j}] + \varepsilon_t \quad (15)$$

where D_t represents a dummy variable that is zero for $t = 1986:2$ to $t = m$, and one for $t = m+1$ to $t = 2007:4$. I exclude the city subscript for convenience. The sequence of regressions for each city involves OLS regressions of (15) with $m = 1989:3$, $m = 1989:4 \dots m = 2004:4$, each with Newey & West (1987) standard errors. Each regression in the sequence has an associated F-test statistic for a null hypothesis of

$$H_0 : \beta_3 = \beta_4 = \beta_5 = 0$$

against the alternative hypothesis

$$H_1 : \beta_3, \beta_4, \text{ or } \beta_5 \neq 0$$

This creates a sequence of F-statistics for each city. The maximum of these statistics for each city is the sup-F statistic. Let z be the period at which the maximum statistic occurs.

The sup-F test is a test of the null hypothesis,

$$H_0 : \beta_{q,t=1986:2} = \beta_{q,t=1986:3} = \dots = \beta_{q,t=2007:4} \text{ for } q = 0, 1, \text{ and } 2.$$

against the alternative,

$$H_1 : \beta_{q,t=1986:2} = \beta_{q,t=1986:3} = \dots = \beta_{q,z-1} \text{ for } q = 0, 1, \text{ or } 2. \\ \beta_{q,z} = \beta_{q,z+1} = \dots = \beta_{q,t=2007:4} \text{ for } q = 0, 1, \text{ or } 2.$$

This test is simply a method of testing for a structural break when the period of the structural break, z , is unknown, and the variables are non-stationary.

Figure 7 – Sup-F Statistic

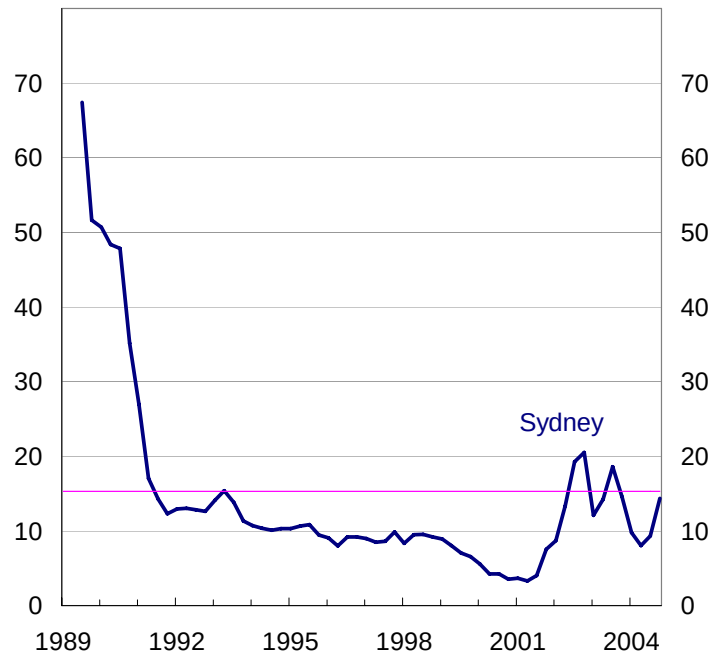


Table 5 – Structural Break Statistics

	Syd	Mel	Bri	Ade	Per	Hob	Dar	Can	Aus
sup-F	67.39	62.38	85.38	66.37	84.75	83.58	45.33	22.38	33.05
Period	1989:3	1993:3	1990:4	2003:4	1989:3	2003:1	2002:1	2002:4	1989:3
Inf-AEG	-5.319	-4.831	-4.19	-3.587	-5.382	-3.13	-3.728	-3.547	-5.189
Period	1999:4	1999:1	1994:4	1995:3	1994:3	1998:3	1994:3	1998:3	1995:1

The 10% critical value for the sup-F test is 15.3 (Hansen, 1992). The 10% critical value for the inf-AEG test is -5.23 (Gregory and Hansen, 1996a).

Figure 7 plots the statistic for Sydney with the values of m plotted on the x-axis. The maximum value of this statistic is 67.39, so this becomes the test statistic. The 10% critical value, marked by a pink line, is 15.3 (Hansen, 1992). Since the test statistic is greater than the critical value, I reject the null hypothesis of no structural break in the relationship for Sydney. Table 5 reports the test statistics for all the cities in the row sup-

F. These statistics range from 22.38 to 85.38, which are greater than the 10% critical value, so I reject the null hypothesis of no structural break in the regressions for all cities.

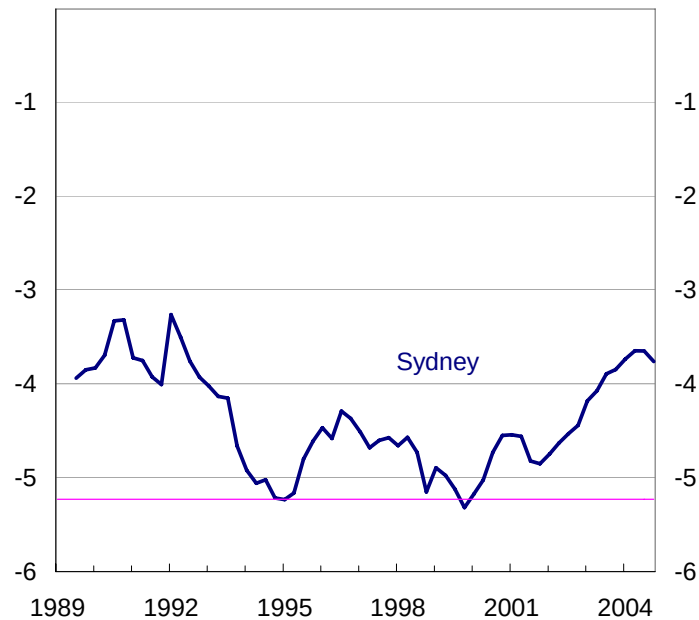
Due to the structural breaks, the AEG tests in Tables 2, 3 and 4 have low power, so I employ the inf-AEG test from Gregory and Hansen (1996a,b), which tests for cointegration in regressions with a structural break. This involves a sequence of the regressions of the form,

$$\log\left(\frac{P}{Y}\right)_t = \beta_0 + \beta_1 \kappa_t + \beta_2 R_t + \beta_3 D_t + \beta_4 (\kappa_t \times D_t) + \beta_5 (R_t \times D_t) + \varepsilon_t \quad (16)$$

for each city, using the sequence of dummy's in (15). Calculating the Augmented Engle-Granger statistic on the residuals, ε_t , for each regression provides a sequence of statistics for each city. The inf-AEG test statistics are the minimums of these statistics for each city, and they test the null hypothesis of no cointegrating relationship between the variables against the alternative of a cointegrating relationship.

Figure 8 plots the inf-AEG statistic over the various values of m for Sydney, and Table 5 reports the test statistics for all cities in the row inf-AEG. In Figure 8, the minimum statistic is -5.319, at 1999:4, which is the test statistic in Table 5. This is less than the critical value of -5.23 (Gregory and Hansen, 1996). Thus, I reject the null hypothesis of no cointegrating relationship for Sydney.

Figure 8 – Inf-AEG Statistic



Based on the statistics in Table 5, I reject the null hypothesis of no cointegrating relationship in Sydney and Perth only. The test statistic for Australia is close to this value, and the Augmented Engle Granger statistic for the test in Table 2, which has low power, is able to reject the null hypothesis of no cointegration, so I maintain the conclusion that there is a cointegrating relationship between these variables at the national level. There is still not enough evidence to reject a null hypothesis of no cointegrating relationship between the variables in Melbourne, Brisbane, Adelaide, Hobart, Darwin, and Canberra. Thus, allowing for structural breaks does not alter the conclusions on whether there is cointegration between the variables.

5.5 Additional variables

Two other possible causes of a lack of evidence of a cointegrating relationship between the variables in these cities are considered. The first possible cause is the assumption that

the primary influence on housing demand is the amount that the individual can borrow. I relax this assumption and include a number of additional variables. The variables added to the model include a measure of inflation, the log of net financial wealth, the log of equity prices, the log of State final demand, and the working age population. These particular variables are included as they are theoretically linked to the demand for housing, and they are occasionally included in empirical house price models, with mixed results on whether they play an important role. See Section 2.1 for more on this.

Two variables that I choose not to include, despite their theoretical link to house prices and their use in the estimates of past authors, are construction costs and the quantity of housing. I exclude construction costs as past evidence does not suggest that they play an important role in the determination of house prices. The quantity of housing would be appropriate in an inverse demand equation, however as this model is a reduced form equation (see Section 3), it is not appropriate to include this variable.

A detailed description of the creation of the additional variables can be found in the Data Appendix. I do not report Augmented Dickey Fuller statistics for these variables, however the tests reveal that the log of net financial wealth is $I(2)$, the log of equity prices, the log of State final demand, and working age population are $I(1)$, and inflation is $I(0)$. Data on some of these variables are not available from 1986:2, so the regressions and subsequent Augmented Engle-Granger statistics employ 79 observations, from 1988:2 to 2007:4.

In order to assess whether there is a cointegrating relationship between the log of the real median house price to real income ratio and the chosen independent variables, I run one OLS regression with the real interest rate, the total housing interest payments to disposable income ratio, and all of the additional variables. Table A1, in the Appendix, reports the coefficient estimates from this regression for each city, and Table 6 reports the Augmented Engle-Granger statistic in the row AEG1. I employ the 10% critical value with five variables in the system and 100 observations from Engle & Yoo (1987), - 4.26, as an approximate critical value. There is insufficient evidence to reject the null hypothesis of no cointegration at the 10% level in all of the regressions, based on these statistics. There are also a number of estimates for which the sign of the estimate does not coincide with *a priori* expectations. It is possible that some of these variables provide benefit, while others distort the results, so regressions with the additional variables included individually are run. Although I do not report these results, the conclusions from both an analysis of the estimates and the Augmented Engle Granger tests run from these regressions does not change from the conclusions in Table 2, regardless of which additional variable is included. Thus, these results suggest that the assumption that factors affecting the amount that the individual can borrow are the primary determinants of the demand for housing does not hinder evidence of a cointegrating relationship between the variables.

Table 6 – Additional Augmented Engle Granger Statistics

Variable	Syd	Mel	Bri	Ade	Per	Hob	Dar	Can	Aus
AEG1	-2.656	-1.819	-1.791	-3.475	-2.925	-2.899	-1.690	-2.733	-2.300
AEG2	-4.068	-3.620	-2.888	-2.964	-3.966	-2.430	-2.419	-2.809	-3.498

AEG1 is the Augmented Engle Granger statistic testing whether there is evidence of a cointegrating relationship between real median house price to real income ratios, real interest rates, housing interest payments to disposable income ratios, a measure of inflation, the log of net financial wealth, the log of equity prices, the log of State final demand, and the working age population. There are only 79 observations in these regressions. An approximate critical value is the Augmented Engle Granger 10% critical value for a sample size of 100, and 5 variables in the system. This is – 4.26 (Engle & Yoo, 1987). AEG2 is the Augmented Engle Granger test statistic testing whether there is evidence of a cointegrating relationship between real median house price to real income ratios and real interest rates, housing interest payments to disposable income ratios, and non-linear terms involving real interest rates. There are 87 observations. An approximate critical value is the 10% critical value with four variables in the system and 100 observations. This is - 3.89 (Engle and Yoo, 1987).

A second possible cause of a lack of evidence of a cointegration relationship in certain cities is the assumption that the coefficient on $\log[1 - (1 + R_t)^{-\tau}]$ is zero. To assess whether this assumption hinders the evidence of a cointegrating relationship, I add this variable to the equation estimated in Table 2 and calculate Augmented Engle Granger statistics. The calculation of this variable requires an estimate of λ , the length of the mortgage. As there are no data on this variable, it is set to 80 quarters. This assumption implies that there has been no change in the length of the average mortgage in the relevant Australian capital city.

I subsequently run OLS regressions of the log of the real median house price to real income ratio on the real interest rate, the total housing interest payments to disposable income ratio, and this additional variable. Table 6 reports the statistics in the row AEG2, and Table A2 reports the coefficient estimates. The 10% critical value with four variables in the system and 100 observations is - 3.89 (Engle and Yoo, 1987). I reject the null hypothesis of no cointegrating relationship for Sydney and Perth only. Thus, the restrictions implemented in Section 3 do not diminish the evidence of a cointegrating relationship between house prices and its determinants.

6. Conclusion

The empirical evidence that I present in this paper suggest that it is plausible that there may be a cointegrating relationship between house prices and determinants of house prices in the Australian data at the capital city level. The evidence is not conclusive, suggesting that there is a cointegrating relationship between the real median house price to real income ratio, real interest rates, and the housing interest payments to disposable income ratio in some Australian State capital cities such as Sydney and Perth, but not in the other State capital cities. Adding certain variables and allowing for structural breaks does not affect this result.

The empirical estimates of the cointegrating relationships suggest that the long run effect of the real interest rate on real median house price to real income ratios varies by city. The sign on a number of the estimates is not as one would expect *a priori*, which is an undesirable feature of the estimates. Additionally, a number of the estimates are

statistically insignificant. These results may be suggesting that a key variable has been omitted from the model.

The estimates of the coefficients on the total housing interest payments to annual disposable income ratio suggest that this variable, which is a measure of credit conditions, plays an important role in the determination of house prices in all capital cities. The estimates of the long run effect of credit conditions on house prices are statistically and economically significant in all cities except Darwin. Increases in lending per dollar of disposable income in the early 2000's, in particular, contribute to the large rises in real median house price to real income ratios.

7. Appendices

A1. Data

I employ the following subscripts in this section:

$i = \{\text{Sydney, Melbourne, Brisbane, Adelaide, Perth, Hobart, Darwin, Canberra, Australia}\}$

$s = \left\{ \begin{array}{l} \text{New South Wales, Victoria, Queensland, South Australia, Western Australia, Tasmania,} \\ \text{Northern Territory, Australian Capital Territory, Australia} \end{array} \right\}$

$t = \{\text{quarter}\}$. When not mentioned, the period 1986:2 to 2007:4 is implied.

Consumer price index ($CPI_{i,t}$) and *consumer price index excluding housing* ($CPIH_{i,t}$)

I employ two measures of the consumer price index, standard Consumer Price Indices, and the consumer price indices excluding housing, which are available on the Australian Bureau of Statistics (ABS) website, www.abs.gov.au, for each capital city. To eliminate the effect of the Goods and Services Tax (GST), I alter the observations for both series from 2001:3 onwards, so that the inflation rate in 2001:3 is the average of the inflation rate for 2001:2 and 2001:4, and the inflation rates from 2001:4 onwards are the same. This creates the series $CPI_{i,t}$ and $CPIH_{i,t}$.

Working age population ($WAP_{i,t}$)

Monthly observations for the number of people in the labour force, and those not in the labour force, in each capital city and Australia, excluding Darwin and Canberra, are available from the ABS website. As measures of the population in Darwin and Canberra, I use the population in the Northern Territory and the Australian Capital Territory. Monthly observations for the working age population in each city are the sum of those in

the labour force and those not in the labour force. Taking the average of the three months in the quarter creates quarterly observations.

Real median house prices ($RMHP_{i,t}$)

Using quarterly data on a number of series that the ABS provide allows the creation of a measure of real median house prices from 1986:2 to 2007:4. House price indices for 1986:2 to 2005:2 ($HPI1_{i,t}$) are available from the ABS website, as are observations for nominal median house prices for the period 2002:1 to 2006:4 ($MHP1_{i,t}$). Australian median house price observations are not available. I convert the nominal series into dollar values by multiplying by 1000. The series for median house prices from 1986:2 to 2001:4 ($MHP2_{i,t}$) is,

$$MHP2_{i,t} = \frac{MHP1_{i,2002:1}}{HPI1_{i,2002:1}} \times HPI1_{i,t}$$

These do not include observations for Australia. I create observations for 2007:1 to 2007:4 using another ABS index for median house prices ($HPI2_{i,t}$), for which observations are available for the period 2002:1 to 2007:4, using the formula,

$$MHP3_{i,t} = \frac{MHP1_{i,2006:4}}{HPI2_{i,2006:4}} \times HPI2_{i,t}$$

Thus, the series for nominal median house prices from 1986:2 to 2007:4, ($NMHP_{i,t}$) is,

$$NMHP_{i,t} = \begin{cases} MHP2_{i,t} & \text{for } 1986:2 \leq t \leq 2001:4 \\ MHP1_{i,t} & \text{for } 2002:1 \leq t \leq 2006:4 \\ MHP3_{i,t} & \text{for } 2007:1 \leq t \leq 2007:4 \end{cases}$$

The series $RMHP_{i,t}$ is then,

$$RMHP_{i,t} = NMHP_{i,t} \times \frac{100}{CPIH_{i,t}}$$

A series of Australian real median house prices is:

$$AUSRMHP_t = \sum_{i=1}^8 \left(\frac{WAP_{i,t}}{\sum_{i=1}^8 WAP_{i,t}} \times RMHP_{i,t} \right)$$

Real average yearly earnings ($RAYE_{s,t}$)

The ABS provides quarterly observations of seasonally adjusted average weekly earnings for each State in dollar terms ($AWE_{s,t}$). The quarters differ from the other data. The observation for 1986:2, for example, is June, July and August, rather than May, June and July. Converting the observations into real, annual terms,

$$RAYE_{s,t} = 52 \times \left(AWE_{s,t} \times \frac{100}{CPI_{i,t}} \right)$$

Real median house price to real average yearly earnings ratio ($RHPY_{i,t}$)

Creating this ratio involves dividing real median house prices by real average yearly earnings.

Real interest rate ($r_{i,t}$)

The nominal interest rate in annual percentage terms (i_t) is the standard variable interest rate provided by banks on housing loans, obtained from the RBA website on a monthly

basis. Taking the average over the quarter creates a quarterly series. Observations for the annual real interest rate in decimal terms ($ar_{i,t}$) are,

$$ar_{i,t} = \frac{i_t}{100} - \left[\left(\frac{CPI_{i,t} - CPI_{i,t-4}}{CPI_{i,t-4}} \right) \right]$$

The interest rate in quarterly percentage terms ($r_{i,t}$),

$$r_{i,t} = 100 \left\{ (1 + ar_{i,t})^{0.25} - 1 \right\}$$

Real households' net wealth per capita ($RNWPC_{i,t}$)

Data on households' financial assets, consumer durables, and total liabilities are available at the national level from the RBA's website for the period 1988:2 to 2007:4. Adding the first two and subtracting the last one provides a measure of net financial wealth (ANW_t).

To create a measure of nominal net financial wealth by city,

$$NW_{i,t} = \frac{WAP_{i,t}}{WAP_{Aus,t}} \times ANW_t$$

where $WAP_{Aus,t}$ is the working age population for Australia. Thus, the weights for the capital cities sum to less than one. Converting into real terms,

$$RNW_{i,t} = NW_{i,t} \times \frac{100}{CPI_{i,t}}$$

Multiplying by one billion, as they are in billions of dollars, and dividing by the working age population in each city provides an estimate of real net wealth per capita in dollars ($RNWPC_{i,t}$).

Real stock prices ($RASX_{i,t}$)

Monthly observations for the S&P/ASX 200 Accumulation Index were available from the ABS website for 1987:2 to April 2008:4. Taking the average over the three observations in the quarter converts these to quarterly observations (ASX_t). Converting into real terms,

$$RASX_{i,t} = ASX_t \times \frac{100}{CPI_{i,t}}$$

State final demand

Quarterly observations of seasonally adjusted State final demand for each State are available on the ABS website. I multiply these values by one million to get the observations in dollar terms, then divide by the working age population in each State (which is also available on the ABS website) to get State final demand per capita.

A2. Tables

Table A1 – OLS Regressions with Additional Variables for Capital Cities									
Variable	Syd	Mel	Bri	Ade	Per	Hob	Dar	Can	Aus
R	0.0046	0.0027	0.1410	-0.0030	-0.0089	0.0259	-0.1347	0.0453	-0.0024
κ	0.0537	0.0515	0.0871	0.1121	0.1058	0.1222	0.0236	0.0996	0.0566
Π	0.0126	0.0110	-0.0183	-0.0143	-0.0055	-0.0168	-0.0375	-0.0067	0.0033
NW	-0.2333	1.2192	0.6458	0.4138	1.0268	0.6652	-0.3617	0.9785	0.3521
RASX	-0.3485	-0.9840	-1.0992	-0.6580	-0.4614	-0.5428	0.7262	-1.1213	-0.5907
SFD	0.4636	2.0816	0.3654	0.7374	0.6091	-0.1051	-0.4828	0.7076	1.1885
WAP	0.0011	0.0008	0.0032	0.0049	0.0004	0.0270	-0.0063	0.0157	0.0002
β_0	-0.486	-24.078	-3.158	-8.577	-11.700	-4.555	4.508	-9.815	-10.289
AEG	-2.656	-1.819	-1.791	-3.475	-2.925	-2.899	-1.690	-2.733	-2.300

The dependent variable is the log of the real median house price to real average yearly income ratio. R is the real interest rate, κ is the housing interest payments to disposable income ratio, π is inflation, NW is the log of net wealth, RASX is the log of a measure of real equity prices, SFD is the log of State Final Demand, WAP is the working age population, and β_0 is the constant. I employ the 10% critical value with five variables in the system and 100 observations from Engle and Yoo (1987), - 4.26, as an approximate critical value for the AEG test.

Table A2 – OLS Regressions with the Non-linear Real Interest Rate									
Variable	Syd	Mel	Bri	Ade	Per	Hob	Dar	Can	Aus
R	0.4148	0.3786	0.1558	0.0951	0.1640	-0.1739	-0.5573	0.1764	0.3362
κ	0.0982	0.1171	0.1145	0.1180	0.1263	0.1283	0.0855	0.0825	0.1066
NLR	-1.5530	-1.4976	-0.5403	-0.2547	-0.5120	0.5337	0.9827	-0.5682	-1.2398
β_0	0.2408	-0.1901	0.5605	0.6653	0.4728	1.1848	2.3876	0.7603	0.1981
AEG	-4.068	-3.620	-2.888	-2.964	-3.966	-2.430	-2.419	-2.809	-3.498

The dependent variable is the log of the real median house price to real average yearly income ratio. R is the real interest rate, κ is the housing interest payments to disposable income ratio, and NLR is the log of the non-linear real interest rate term. An approximate critical value for the AEG test is the 10% critical value with four variables in the system and 100 observations. This is - 3.89 (Engle and Yoo, 1987).

A3. Figures

Figure A1 - House Price to Income Ratios for Melbourne and Brisbane

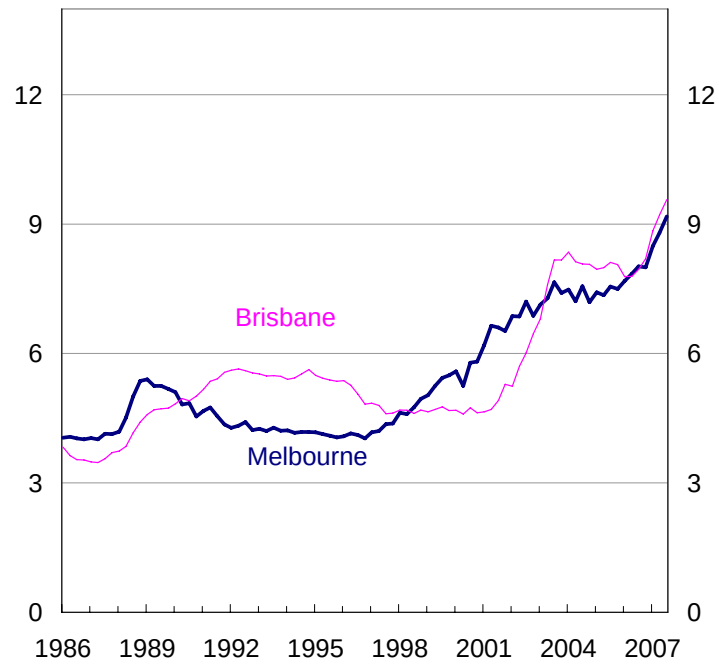


Figure A2 - House Price to Income Ratios for Adelaide and Perth

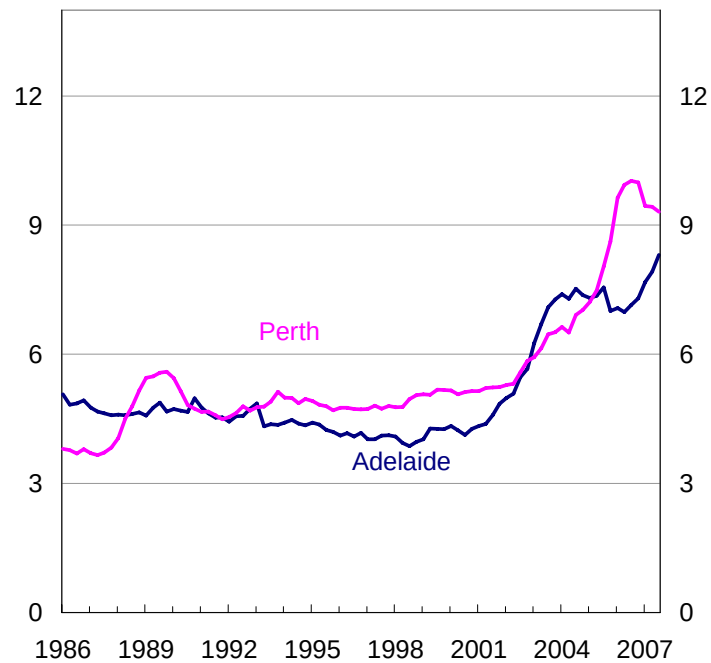


Figure A3 - House Price to Income Ratios for Hobart and Darwin

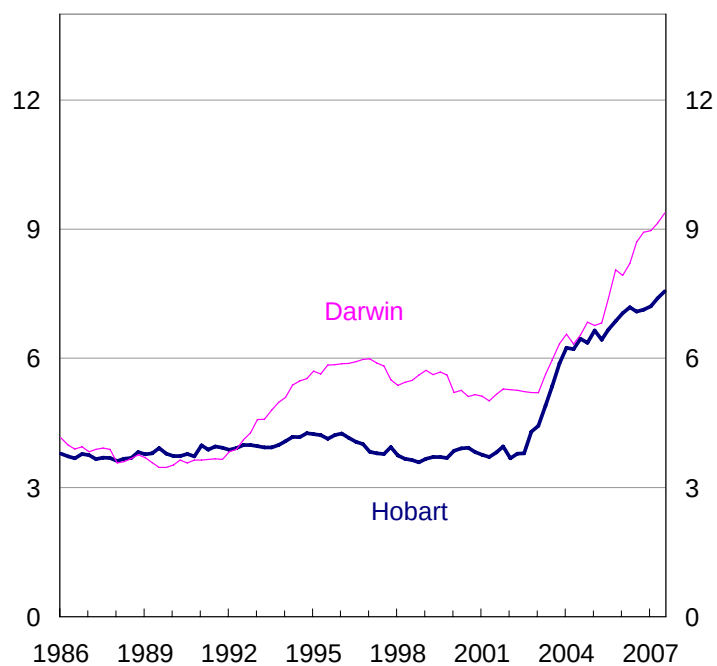
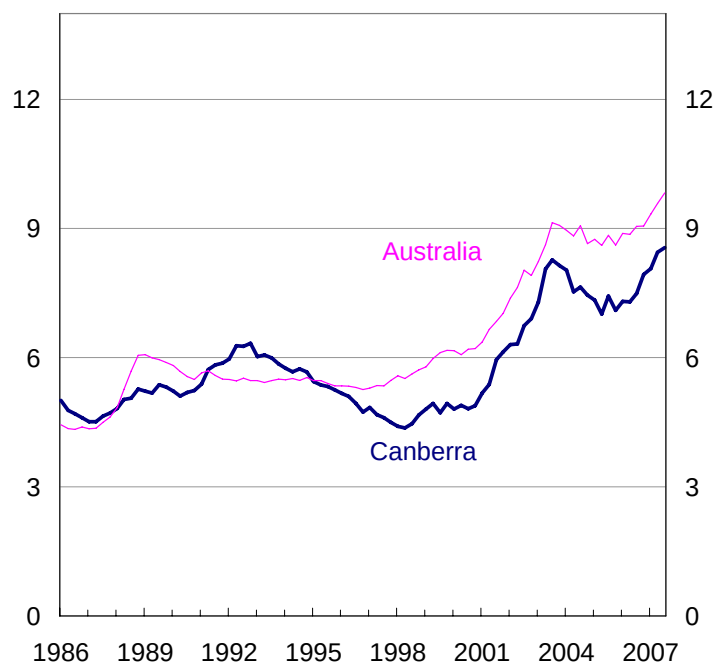


Figure A4 - House Price to Income Ratios for Canberra and Australia



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