THE ECONOMIC RECORD, VOL. 81, NO. 255, AUGUST, 2005, S96–S103

Explaining House Prices in Australia: 1970–2003*

PETER ABELSON, ROSELYNE JOYEUX, GEORGE MILUNOVICH and DEMI CHUNG
Macquarie University, Sydney, Australia

This paper aims to explain changes in real house prices in Australia from 1970 to 2003. We develop and estimate a long-run equilibrium model that shows the real long-run economic determinants of house prices and a short-run asymmetric error correction model to represent house price changes in the short run. We find that, in the long run, real house prices are determined significantly and positively by real disposable income and the consumer price index. They are also determined significantly and negatively by the unemployment rate, real mortgage rates, equity prices and the housing stock. Employing our short-run asymmetric error correction model, we find that there are significant lags in adjustment to equilibrium. When real house prices are rising at more than 2 per cent per annum, the housing market adjusts to equilibrium in approximately four quarters. When real house prices are static or falling, the adjustment process takes six quarters.

I Introduction

In this paper we aim to explain the changes in real house prices in Australia from 1970 to 2003. The following section provides a brief outline of house prices over this period. In section III we discuss how house prices are determined. Section IV describes the basis for our econometric modelling. Sections V and VI give the results of our modelling work. A final section summarises the main points. We find that real house prices are significantly determined by income, unemployment rate, real mortgage rates, equity prices, consumer price index (CPI) (as a reflection of inflation) and the supply of housing. However, house prices typically take 1–2 years to adjust to equilibrium values.

II House Prices in Australia: 1970–2003

Figure 1 shows a real quarterly house price index for Australian capital cities from 1970 to 2003 using an Australian Treasury estimated series.[1] This series is based on estimated nominal median house prices for the capital cities. The city house prices are weighted according to the city weight in the national CPI and the weighted nominal index is converted into a real index using the CPI deflator. In a detailed description of house prices in Australia, Abelson and Chung (2004) show that, although house price *levels* differ considerably between the cities, *changes* in city house prices have been similar, especially since 1990. They also show that changes in median unit prices in cities and changes in non-metropolitan house prices were both broadly similar to changes in median house prices in the cities. Accordingly, a capital city house price index is representative of all residential prices in Australia.

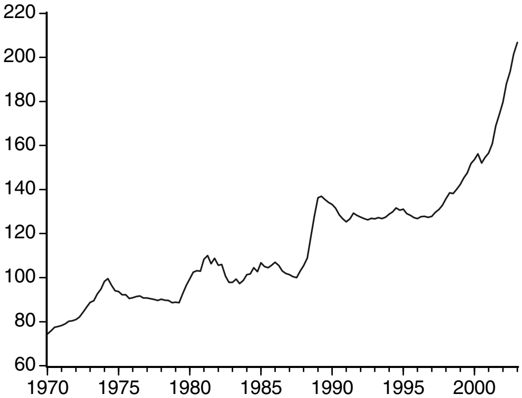
The national real house price index has two significant features. First, there has been a substantial

* We wish to thank two referees for their comments on the draft paper.
JEL classifications: G12, R31
Correspondence: Peter Abelson, Department of Economics, Macquarie University, Sydney, NSW 2109, Australia. Email: pabelson@efs.mq.edu.au.

[1] The Productivity Commission provided this unpublished index to the authors. Abelson and Chung (2004) show that the estimated Treasury national house price index is similar to one that they estimated using their preferred city figures and weights based on city housing stocks. In this paper, we use the Treasury index because it is a quarterly index, whereas the Abelson–Chung index is an annual one.

© 2005. The Economic Society of Australia. ISSN 0013–0249

FIGURE 1
Australian Real House Price Index



long-run increase in observed real house prices. Real house prices rose by approximately 180 per cent between 1970 and 2003, equal to a compound rate of approximately 3.3 per cent per annum. However, two caveats are necessary: (i) these figures do not allow for quality changes in housing. Drawing on data for expenditure on alterations and additions and on new house sizes, Abelson and Chung (2004) estimate that the quality of the housing stock increased by 1 per cent per annum on average. On a quality-adjusted basis, real house prices rose by approximately 2.3 per cent per annum and approximately doubled between 1970 and 2003. (ii) The year 2003 was almost certainly a peak house price year, whereas 1970 was not. Despite these caveats, there has been a striking long-run increase in real house prices across Australia.

The second major feature of real house prices was the boom and stagnation pattern within this long-run trend. During this 33-year period, there were four significant house price booms: from 1971 to 1974, from 1979 to 1981, from 1987 to 1989 and from 1996 to 2003 (the most prolonged price boom). After each of the first three booms, real house prices tended to be constant or fall for a few years. These are the major events that this paper seeks to explain.

III Possible Explanations of Australian House Prices

Real house price changes can be expressed generally as:

$$P_t - P_{t-1} = \alpha(D_t - S_t), \tag{1}$$

where P is the real price of housing, D is the estimated demand for housing (which includes consumption and investment demand), S is the supply of housing and the subscript t refers to the period. The difference between estimated demand and supply picks up the disequilibrium influence on prices.

Following standard demand theory, the demand for housing can be viewed as a function of disposable income (Y), user cost (UC) and demographic factors (DEM) such as population growth or household formation.

$$D_t = f(Y, UC, DEM)_t \tag{2}$$

and

$$UC_t = g(P_t, r, P_{t+1}, TS_t), \tag{3}$$

where r is the rate of interest and TS is the taxes or subsidies (as relevant). The future house price represents capital gain, which drives investment demand. Depreciation and maintenance costs are not included in this user-cost equation. Depreciation is subsumed in the expected real house price. Maintenance expenditures do not vary much from 1 year to another.

To incorporate investment motivation, we may also allow for the relative return on housing and other assets insofar as they are not captured in equations 2 and 3. The user cost of housing includes part of the rate of return on housing through the expected capital gain (P_{t+1}). Also, the borrowing rate of interest (r) reflects the (pre-tax) return on alternative cash investments. A full investment model could include rents from housing, the return on other assets such as equities and taxation effects.

Taxation (subsidy) effects may reflect specific housing taxes or subsidies (such as grants for first-home owners or the exemption of owner-occupied homes from capital gains tax) or general taxes. When house buyers purchase a home with their own funds, the user cost is the *after-tax* interest that is foregone. This falls and the demand for housing rises with high marginal tax rates. Also, when nominal income is taxed and capital gains are not taxed, a combination of high inflation and taxation increases the demand for housing.

Turning to the supply of housing, the stock does not change much from year to year. In 2001, there were 7.1 million residential dwellings in Australia. Dwelling completions average approximately 2 per cent of the housing stock per annum. Net increases in stock are lower after allowing for losses of housing stock. The inelastic nature of housing supply is doubtless one reason why housing stock is rarely included in short-run house price models. Another reason may be the difficulty of estimating housing stock on a quarterly basis. However, omitting housing stock from explanatory models may be a major error. We find that the housing stock per capita has a substantial long-run impact on house prices and hence on the short-run model as well.

Some models of house prices include the costs of building new houses, for example, Bourassa and Hendershott (1995) and Bodman and Crosby (2003). However, there is little justification for this. In the short run, the prices of new houses are determined by the value of the existing housing stock. The costs of new houses can affect the price of existing housing only if new house supply significantly affects the size of the housing stock (which is better included as a separate variable). In other circumstances, changes in the cost of new houses (be they changes in construction costs or taxes on developers or GST) reduce the value of land for new housing, but do not affect the price of new or established houses (Abelson, 1999).

Accordingly, in the econometric literature (see Hendry, 1984; Meen, 1990; Muellbauer and Murphy, 1997) the house price model is most often an inverted demand equation of the following general kind:

$$P_t = f(Y, r, P^e, DEM, TS)_t, \tag{4}$$

where P^e is the expected real house price and the other variables are the same as explained earlier. However, economic theory does not provide a finite list of variables. Investment factors, such as housing rents, return on equities and the CPI combined with marginal tax rates, could affect house prices. Indeed, in an open economy or for an international city, such as Sydney or Melbourne, the exchange rate could influence house prices (a low exchange rate increases the attractiveness of housing assets to foreigners). Moreover, care is required with respect to multicollinearity and endogeneity. For example, household formation may be a function of house prices. Importantly, equation 4 says nothing about possible lagged or speculative effects or other disequilibria in the market. We address these issues in our empirical work,which follows.

IV Long-Run and a Short-Run Model of House Prices

In this paper, we model the quarterly real Australian house price index estimated by the Treasury from 1970 to 2003. Our approach is to model house prices in both the long and short run. We assume that in the long run, real house prices adapt to economic fundamentals. In the short run, house prices may deviate from this equilibrium, but they continually readjust to it in a non-linear fashion through an asymmetric error correction term. The asymmetry arises because we expect that adjustments to equilibrium will be faster when prices are rising than when they are flat or falling. Our observation is that buyers are keen to get into the housing market when prices are rising for fear that delay will mean paying still higher prices. However, sellers are often unwilling to reduce prices when markets are flat. Transaction times are typically shorter when real prices are rising than in flatter periods.

More specifically, we:

- specify the variables to be included in the long-run relationship;
- examine the time-series properties of those variables – specifically unit root and co-integration tests are carried out;
- estimate the long-run equilibrium relationship; and
- we derive and estimate a non-linear error correction model of house prices.

The long-run model in this analysis is

$$\begin{aligned}\log(P_t) = {} & \gamma_1 + \gamma_2 \log(Allords_t) + \gamma_3 R_t \\ & + \gamma_4 \log(GDI_t) + \gamma_5 \log(ER_t) \\ & + \gamma_6 \log(UE_t) + \gamma_7 \log(CPI_t) \\ & + \gamma_8 \log(H_t) + u_t \end{aligned} \tag{5}$$

where P_t is the real house price, R_t is the real mortgage interest rate (quarterly yield), $Allords_t$ is the real All Ordinaries index, GDI_t is the real household disposable income per capita, ER_t is the trade-weighted exchange rate, CPI_t is the consumer price index, UE_t is the unemployment rate, H_t is the detached housing stock per capita and u_t is a stationary error term. The All Ordinaries index is included to capture a possible substitution effect with the stock market. The exchange rate is a potentially important variable as it may influence overseas demand for Australian real estate. The CPI is included to capture the after-tax investment advantages of housing as an asset in conditions of rising prices. The unemployment rate is included as a barometer of economic conditions and is expected to influence the price of houses negatively. The housing stock is included as a supply variable. It is expected to be negatively related to house prices.

Further details of the data are described in Appendix I. We report below the results for the full sample of observations, 1970q1–2003q1 and for a subsample, 1975q1–2003q1. As a result of the concerns with the early 1970s house price data, we comment only on the results for the subsample, which we deem more reliable.

In order to examine the time series properties of the variables, we define the vector z_t as:

$$\begin{aligned} z_t = {} & (\log(P_t), \log(Allords_t), R_t, \\ & \log(GDI_t), \log(ER_t), \log(UE_t), \\ & \log(CPI_t), \log(H_t)). \end{aligned} \tag{6}$$

We test the elements of z_t for unit roots and then test z_t for multivariate co-integration using the Johansen tests. The long-run equilibrium relationship is estimated using Stock and Watson's dynamic ordinary least squares (DOLS) and finally a nonlinear error correction model of house prices is derived and estimated.

(i) Unit Root Tests

Before we test for co-integration between the variables of the price equation, we need to ensure that the variables are integrated of order 1 ($I(1)$). We conduct unit root tests for each variable using the Dickey–Fuller (1979, 1981) tests and Phillips–Perron (1988) tests. The results are summarised in Table I of Appendix I. The unit root tests show that all the variables have a unit root and it is therefore appropriate to test for co-integration among those variables.

(ii) Tests for Multivariate Co-integration

To test for co-integration, we employ the Johansen (1988, 1991) maximum likelihood procedure[2] for z_t and an unrestricted constant term.[3]

Table 1 presents the trace and maximum eigenvalue statistics for two sample periods: 1970q2–2003q1 and 1975q1–2003q1. On the basis of the maximum eigenvalue and the trace tests we conclude that there are five co-integrating vectors in the full sample period and four co-integrating vectors in the period excluding the early 1970s. Therefore, from the Johansen tests for co-integration we conclude that z_t is co-integrated.

TABLE 1
Tests for the Co-integration Rank

H_0	λ_{trace}	$\lambda_{trace}(0.95)$	λ_{max}	$\lambda_{max}(0.95)$
1975q1–2003q1				
$r = 0$	274.78	156.00	88.73	51.42
$r \leq 1$	186.05	124.24	68.02	45.28
$r \leq 2$	118.04	94.15	41.78	39.37
$r \leq 3$	76.26	68.52	38.10	33.46
$r \leq 4$	38.16	47.21	19.90	27.07
$r \leq 5$	18.26	29.68	11.98	20.97
$r \leq 6$	6.28	15.41	6.28	14.07
$r \leq 7$	0.01	3.76	0.01	3.76
1970q2–2003q1				
$r = 0$	255.00	156.00	70.31	51.42
$r \leq 1$	184.69	124.24	49.89	45.28
$r \leq 2$	134.80	94.15	48.35	39.37
$r \leq 3$	86.45	68.52	34.24	33.46
$r \leq 4$	52.21	47.21	31.53	27.07
$r \leq 5$	20.68	29.68	10.51	20.97
$r \leq 6$	10.17	15.41	8.37	14.07
$r \leq 7$	1.79	3.76	1.79	3.76

Critical values from Osterwald-Lenum (1992) are denoted by $\lambda_{trace}(0.95)$ and $\lambda_{max}(0.95)$.

V Estimation of the Long-Run Price Equation

The co-integration vectors can be estimated using the Johansen estimation method. However, given the large number of variables included in the model relative to the number of observations, we use a more robust method proposed by Stock and Watson (1993).

We partition z_t into $\log(P_t)$ and $x_t = (\log(Allords_t), R_t, \log(HDI_t), \log(ER_t), \log(UE_t), \log(CPI_t), \log(H_t))'$. $\text{Log}(P_t)$ is then regressed on a constant x_t and k leads and lags of x_t.

$$\log(P_t) = a_0 + \theta x_t + \sum_{j=-k}^{k} \delta_j \Delta x_{t-j} + v_t. \quad (7)$$

In equation 7, the DOLS estimator of θ is the OLS estimator of θ in equation 7. If $\log(P_t)$ and x_t are co-integrated, then the DOLS estimator of θ is efficient and consistent in large samples even if x_t includes some endogenous variables. The usual tests of hypothesis on the coefficients of equation 7 are also valid if robust standard errors, such as the Newey–West (1987) standard errors, are used.

The DOLS estimation results are presented in Table 2 for both the sample periods. For both the samples, the value of k was chosen to be 2 as was

[2] It should be noted that the Johansen co-integration tests (and the Engle and Granger (1987) test) are mis-specified if the adjustment is asymmetric. Enders and Siklos (2001) generalise the Engle–Granger test to allow for threshold adjustment towards a co-integrating relationship. They concur with Balke and Fomby (1997) that in the presence of threshold adjustment the Engle–Granger test can be used to determine whether the variables are co-integrated and if non-linearity is suspected a non-linear adjustment process should be estimated. However, the Engle–Granger and the Enders–Siklos tests suffer from low power and the usual problems associated with a single equation approach; consequently, we chose to test for co-integration using the Johansen method.

[3] We include seasonal dummy variables and four dummy variables: one taking the value 1 in the fourth quarter of 1987 (stock market crash), one taking the value 1 in the fourth quarter of 1976, one taking the value 1 in the third quarter of 1986 and one taking the value 1 in the first quarter of 1981. The number of lags in the unrestricted VAR is set to $k + 1 = 3$; it is chosen such that the residuals fulfill the required assumptions and in order to minimise the conventional information criteria. The constant is not restricted to the co-integrating space (such a choice was confirmed by testing the joint hypothesis of both the rank order and the deterministic components).

TABLE 2
Stock–Watson Dynamic Ordinary Least Square (DOLS) Long-Run Coefficients

	Estimated coefficient (SE)	
Variable	1970q2–2003q1	1975q1–2003q1
Constant	−5.0415* (2.6898)	−5.5212 (3.3698)
log(*Allords*)	−0.0804 (0.0528)	−0.1421** (0.0690)
R	−0.0424 (0.0263)	−0.0542* (0.0318)
log(*GDI*)	1.4051** (0.3547)	1.7097** (0.3619)
log(*ER*)	0.0254 (0.1132)	−0.0034 (0.1051)
log(*UE*)	−0.2558** (0.0528)	−0.1895** (0.0582)
log(*CPI*)	0.7172** (0.1738)	0.7632** (0.2113)
log(*H*)	−3.3364** (1.5639)	−3.5980* (1.8836)
R^2 adjusted	0.9616	0.9565

** and * indicate 5 and 10 per cent significance level, respectively. Standard errors are the Newey–West standard errors computed with three lags.
Note: Dependent variable $\log(P_t)$.

the case for the multivariate vector auto regressions (VAR) model. Except for the coefficient of the real interest rate, those estimates are the long-run elasticities of the real house price with respect to the individual variables.

If we look at the subsample period excluding the early 1970s, we see that all signs of the coefficients accord to expectations. The coefficients are significant, except for the coefficients of the exchange rate.[4]

Several results are of considerable interest. Note first the high long-run elasticity of real house prices with respect to real disposable income per capita, which is significantly larger than 1 at the 5 per cent significance level. Real prices would increase by 1.71 per cent on average after an increase of 1 per cent in real disposable income. Second, a 1 per cent increase in the CPI index results in an estimated 0.76 per cent increase in real house prices.[5]

However, an increase of 1 percentage point in the real mortgage rate will lead to a fall in house prices of 5.4 per cent on average. The coefficient of the All Ordinaries is significant and also negative, pointing to an asset substitution effect from stocks to housing after the 1987 and 2000 stock market downturns. The coefficient of unemployment is very significant and negative, suggesting that this is an indicator of economic conditions. Of special interest, because this has not been estimated in other studies and because of its potential policy importance, the coefficient of the housing stock variable is significant at the 5 per cent level. A 1 per cent increase in housing stock per capita leads to an estimated decrease in real house price of 3.6 per cent on average.

VI Asymmetric Error Correction Short-Run Model of House Prices

To estimate the short-run parameters, we estimate an asymmetric error correction price equation. The model is:

$$\begin{aligned}\Delta \log(P_t) = {} & b_0 + \alpha_1 I_{t-1}(\log(P_{t-1}) - \hat{\theta} x_{t-1}) \\ & + \alpha_2(1 - I_{t-1})(\log(P_{t-1}) - \hat{\theta} x_{t-1}) \\ & + \sum_{j=1}^{k} b_j \Delta z_{t-j} + \varepsilon_t, \end{aligned} \quad (8)$$

where $\hat{\theta}$ is the estimated DOLS co-integrating vector and I_t is the Heaviside indicator function which defines 'boom' observations as observations for which the real price growth over the past year has been over 2 per cent. We chose the value of 2 per cent because this is the average annual value of improvements plus selling costs.[6]

$$I_t = \begin{cases} 1 & \text{if } \log(P_t) - \log(P_{t-4}) > 0.02, \\ 0 & \text{otherwise.} \end{cases}$$

The only exceptions are observations 2000q3, 2003q4 and 2001q1, which would be classified as 'non-boom' observations as a result of the introduction of the GST in 2000q3. To redress this distortion, those observations were classified as 'boom' observations.

The existence of co-integration between some variables implies that those variables move together through time, tracing a long-run path from which they are disturbed by temporary shocks, but to which they continually readjust.[7] The significance tests on the

[4] We have kept the exchange rate in the equation even though it is insignificant in the long run because it is significant in the short run. Moreover, we expect the exchange rate to be an important factor in determining Sydney house prices. Bewley *et al.* (2004) show that Sydney house prices lead other capital cities prices, consequently the inclusion of the exchange rate appears warranted.

[5] A recent international study (Tsatsaronis and Zhu, 2004) found that inflation accounted for over half the total variation in house prices in 17 developed countries over the period 1970–2003.

[6] It is also the value chosen by Muellbauer and Murphy (1997) in a different model specification.

[7] Note that, because we have not estimated the full vector error correction model (VECM) system and identified all the co-integrating vectors, it is not meaningful to comment on whether or not the system is in equilibrium at any point in time by considering this single equation estimated error correction term.

TABLE 3
Asymmetric Error Correction Model of Real House Prices

Variable	Estimated coefficient (SE) 1970q2–2003q1	1975q1–2003q1
$\alpha 1$	−0.1446** (0.0492)	−0.2100** (0.0503)
$\alpha 2$	−0.1402* (0.0573)	−0.1396* (0.0575)
R^2 adjusted	0.2958	0.3971
Serial correlation		
B–P–L1 $\chi^2(1)$	0.0069	0.0467
B–P–L2 $\chi^2(2)$	0.1043	0.0806
B–P–L4 $\chi^2(4)$	0.2666	1.7815
Heteroscedasticity $\chi^2(1)$	2.063	0.179
ARCH(1) $\chi^2(1)$	0.629	0.416

** and * indicate 1 and 5 per cent significance level, respectively. B-P-L1, B-P-L2 and B-P-L4 stand for the Box–Pierce–Ljung autocorrelation tests for first-order, second-order and fourth-order autocorrelation, respectively. The heteroscedasticity test is based on a regression of the squared residuals on a constant and the squared fitted values. According to the diagnostic tests there is no problem of heteroscedasticity, autocorrelation or ARCH(1) effects for both the periods.

differenced explanatory variables give us information on the strength of the short-term effects. The coefficients on the lagged error correction terms represent the proportion by which the long-term disequilibrium in the log of real house prices is being corrected in each period. In our model, we posit that those speeds of adjustment are different in boom and non-boom years. Estimation results for equation 8 are presented in Table 3.

For both the samples, the speeds of adjustment α_1 and α_2 are significant at the 1 and 5 per cent levels, respectively (although using the Dickey–Fuller critical values only α_1 is significant at the 5 per cent level). They are both negative, indicating adjustment to equilibrium. For both the periods, the speeds of adjustment are not significantly different during boom and outside boom times. Excluding the early 1970s, if an external shock throws the variables out of equilibrium during 'boom' time and assuming no further shocks, the price adjusts to its long-run equilibrium with approximately 21 per cent of the adjustment taking place in each quarter. Such an adjustment speed is reasonably fast. Outside 'boom' time, the price adjusts to its long-run equilibrium with approximately 14 per cent of the adjustment taking place per quarter.

These results show that if there is an external shock, such as a tax shock for example, prices will adjust rapidly during boom times and approximately 30 per cent slower during flat or falling times. Our model does not provide precisely determined lags, because it does not account for feedback effects. However, when real prices are rising at more than 2 per cent per annum, the housing market adjusts to equilibrium in approximately four quarters. When real prices are static or falling, the adjustment process takes approximately six quarters.

VII Conclusions

Between 1970 and 2003, Australian real house prices rose by over 3 per cent per annum. On a quality-adjusted basis, house prices rose by approximately 2.3 per cent per annum. Over this period, there were four house price booms: 1972–1974, 1979–1981, 1987–1989 and 1996–2003. In between these booms, real house prices tended to fall.

In this paper we estimate a long-run equilibrium model of house prices that shows the real economic determinants of house prices and a short-run asymmetric error correction model to represent house price changes in the short run. Consistent with the economic theory, we find that in the long-run real house prices are determined significantly by real disposable income, unemployment, real interest rates, equity prices, CPI and supply of housing.

Summarising particular results, the estimated long-run elasticity of real house prices is 1.7 with respect to real disposable income per capita and 0.8 with respect to CPI (reflecting a mixture of expected capital gains and tax benefits). However, there appears to be a strong negative relationship between real house prices and real mortgage rates (a 1 per cent rise/fall in the real mortgage rate will lead to a fall/rise in real house prices of 5.4 per cent on average). The estimated long-run elasticity of real house prices is −0.2 with respect to unemployment (a measure of economic confidence) and −0.14 with respect to All Ordinaries index. Also of potential policy interest is the estimated long-run elasticity with respect to housing stock per capita, which is estimated to be -3.6.

Employing our short-run asymmetric error correction model, we find that there are significant lags in adjustment to equilibrium. Although our model does not provide precisely determined lags, when real house prices are rising at more than 2 per cent per annum the housing market adjusts to equilibrium in four quarters. When real prices are flat or falling, the adjustment process takes six quarters.

REFERENCES

Abelson, P. (1999), 'The Real Incidence of Imposts on Residential Land Development and Building', *Economic Papers*, **18**, 85–90.

Abelson, P. and Chung, D. (2004), *Housing Prices in Australia 1970 to 2003*, Macquarie University Economics Research Paper, No. 9, September 2004, Macquarie University.

Balke, N.S. and Fomby, T. (1997), 'Threshold Cointegration', *International Economic Review*, **38**, 627–43.

Bewley, R., Dvornak, N. and Livera, V. (2004), *House Price Spirals: Where the Buck Starts*, CommSec, Commonwealth Bank, Sydney.

Bodman, P.M. and Crosby, M. (2003), *How Far to Fall? Bubbles in Major City House Prices in Australia*, Working Paper, Melbourne Business School.

Bourassa, S.C. and Hendershott, P.H. (1995), 'Australian Capital City Real House Prices, 1979–93', *Australian Economic Review*, **3rd quarter**, 16–26.

Campbell, J. and Perron, P. (1991), 'Pitfalls and Opportunities: What Macroeconomists Should Know About Unit Roots', *NBER Macroeconomics Annual*, **6**, 141–63.

DeJong, D.N., Nankervis, J.C., Savin, N.E. and Whiteman, C.H. (1992), 'The Power Problems of Unit Root Tests in Time Series with Autoregressive Errors', *Journal of Econometrics*, **53**, 323–43.

Dickey, D.A. and Fuller, W.A. (1979), 'Distribution of the Estimators for Autoregressive Time Series with a Unit Root', *Journal of the American Statistical Association*, **74**, 427–31.

Dickey, D.A. and Fuller, W.A. (1981), 'Likelihood Ratio Statistics For Autoregressive Time Series with a Unit Root', *Econometrica*, **49**, 1057–72.

Enders, W. and Siklos, P.L. (2001), 'Cointegration and Threshold Adjustment', *Journal of Business and Economic Statistics*, **19**, 166–76.

Engle, R.F. and Granger, C.W.J. (1987), 'Cointegration and Error Correction: Representation, Estimation and Testing', *Econometrica*, **55**, 251–76.

Hendry, D.F. (1984), 'Econometric Modelling of House Prices in the United Kingdom', in Hendry, D.F. and Wallis, K.F. (eds), *Econometrics and Quantitative Economics*, Basil Blackwell, Oxford.

Johansen, S.J. (1988), 'Statistical Analysis of Co-integration Vectors', *Journal of Economic Dynamics and Control*, **12**, 231–54.

Johansen, S.J. (1991), 'Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregression Models', *Econometrica*, **59**, 1551–80.

Kwiatkowski, D., Phillips, P.C.B., Schmidt, P. and Shin, Y. (1992), 'Testing the Null Hypothesis of Stationarity Against the Alternative of a Unit Root: How Sure are We that Economic Time Series Have a Unit Root', *Journal of Econometrics*, **54**, 159–78.

MacKinnon, J. (1991), 'Critical Values for Cointegration Tests', in Engle, R.F. and Granger, C.W.J. (eds), *Long-Run Economic Relationships*, Oxford University Press, Oxford, 267–76.

Meen, G. (1990), 'The Removal of Mortgage Constraints and the Implications for Econometric Modeling of UK House Prices', *Oxford Bulletin of Economic and Statistics*, **32**, 1–23.

Muellbauer, J. and Murphy, A. (1997), 'Booms and Busts in the UK Housing Market', *Economic Journal*, **107**, 1701–27.

Newey, W.K. and West, K.D. (1987), 'A Simple, Positive Semi-definite Heteroscedasticity and Autocorrelation Consistent Covariance Matrix', *Econometrica*, **55**, 703–8.

Osterwald-Lenum, M. (1992), 'A Note with Quantiles of the Asymptotic Distribution of the ML Cointegration Rank Test Statistics', *Oxford Bulletin of Economic and Statistics*, **54**, 461–72.

Phillips, P.C.B. and Perron, P. (1988), 'Testing for a Unit Root in Time Series Regression', *Biometrica*, **75**, 335–46.

Stock, J.H. and Watson, M. (1993), 'Testing for Common Trends', *Journal of the American Statistical Association*, **83**, 1097–107.

Tsatsaronis, K. and Zhu, H. (2004), 'What Drives Housing Price Dynamics: Cross-Country Evidence', *BIS Quarterly Review*, **March**, 65–78.

APPENDIX I

Data Description and Sources

The dataset used in this study consists of quarterly series on eight variables including: real house price index, exchange rate, All Ordinaries share price index, unemployment rate, consumer price index, real mortgage rate, gross disposable income per capita and estimated national stock of detached houses per capita. The quarterly real house price index is an unpublished index estimated by the Australian Department of Treasury. We obtained the series for unemployment rate, All Ordinaries, CPI, gross disposable income and exchange rate from the DX database for the period March 1970–March 2003. The exchange rate variable used is a trade-weighted index value of the Australian dollar. Real gross disposable income per capita is derived by dividing the real gross disposable income by population and the real All Ordinaries share price index by deflating the All Ordinaries index by CPI. The real mortgage rate variable was estimated by adjusting the standard variable mortgage rate (from the Reserve Bank website) for inflation.

We constructed the (detached) housing stock variable from census data 1981 and 1986 and from completion data for the whole period. We used the housing stock data from the 1981 and 1986 censuses (because these censuses used consistent definitions of housing) together with completion data between 1981 and 1986 to compute a quarterly depreciation rate between these two dates: δ. We then constructed a housing stock for the whole period by using the formula: $H_t = (1 - \delta)H_{t-1} + C_t$ where H_t is the housing stock at time t and C_t is the completion at time t. The source for the completion data was ABS, Table 8752.23a.

Unit Root Tests

In order to identify degrees of integration of the above-mentioned series, we perform the Augmented Dickey–Fuller (ADF, 1979) and the Phillips–Perron (1988) tests. Both procedures test the null hypothesis of a unit root in the autoregressive representation of the series. The ADF test constructs a parametric correction for higher-order correlation by assuming that the series follows an $AR(p)$ process and adding lagged difference terms of the dependent variable to the right-hand side of the test regression:

$$\Delta y_t = \alpha_0 + \gamma y_{t-1} + \alpha_1 t + \sum_{i=1}^{p-1} \beta_i \Delta y_{t-i} + \varepsilon_t.$$

The unit root test is carried out by testing the null $\gamma = 0$ using MacKinnon (1991) critical values. The number of lagged difference terms p is determined using the Schwartz Information Criterion. Phillips–Perron (PP) test differs from the ADF test in that it accounts for serial correlation non-parametrically when testing for unit root. This procedure estimates the non-augmented DF specification:

$$\Delta y_t = \alpha_0 + \gamma y_{t-1} + \alpha_1 t + \varepsilon_t$$

and modifies the *t-ratio* of the γ coefficient so that serial correlation does not affect the asymptotic distribution of the test statistic. Thus, the corrected *t-ratio* has the same asymptotic distribution as the ADF statistic. The tests statistics presented are for the model including a trend term. Sequential testing after exclusion of the nuisance parameters when necessary did not change the results. Table I presents the ADF and the PP tests on both the series in levels and first differences.

Except for the real mortgage rate, all variables appear to be $I(1)$ processes according to both the ADF and Phillips–Perron tests. The real mortgage rate appears to be integrated of order one according to the ADF test and stationary according to the PP test. However, both these tests suffer from a lack of power as noted in Campbell and Perron (1991) and DeJong *et al*. (1992). Because the two tests provide contrary conclusions for the real mortgage rate, we further carry out the Kwiatkowski, Phillips, Schmidt and Shin (KPSS, 1992) test. The KPSS test differs from the other unit root tests used here in that the series is assumed to be (trend-) stationary under the null. Rejecting the null would lead us to conclude that the series is non-stationary. Table II presents KPSS test results for the real mortgage rate variable.

As indicated by the KPSS test statistic, we can reject the null hypothesis of stationarity at 1 per cent significance level. In combination with evidence from the ADF test, we conclude that the real mortgage rate is $I(1)$.

TABLE I
Unit Root Tests

Variable	ADF	PP
log(*P*)	−2.609 (0.277)	−1.666 (0.761)
log(*CPI*)	0.621 (0.999)	−0.051 (0.995)
log(*GDI*)	−2.765 (0.213)	−3.362 (0.061)
log(*Allords*)	−3.107 (0.109)	−3.148 (0.099)
log(*UE*)	−2.063 (0.561)	−1.856 (0.672)
log(*ER*)	−1.772 (0.713)	−2.045 (0.571)
log(*H*)	−1.951 (0.622)	−1.026 (0.936)
R	−1.962 (0.616)	−7.939 (0.000)
Variables in first difference		
log(*P*)	−6.605 (0.000)	−6.751 (0.000)
log(*CPI*)	−5.018 (0.000)	−8.492 (0.000)
log(*GDI*)	−10.983 (0.000)	−15.489 (0.000)
log(*Allords*)	−11.384 (0.000)	−11.384 (0.000)
log(*UE*)	−7.251 (0.000)	−7.370 (0.000)
log(*ER*)	−10.202 (0.000)	−10.194 (0.000)
log(*H*)	−3.984 (0.011)	−9.264 (0.000)
R	−11.476 (0.000)	−26.005 (0.000)

Note: Test statistics under the null hypothesis of unit root are presented with their *P*-values in brackets. ADF, Augmented Dickey–Fuller.

TABLE II
Additional Unit Root Test for the Real Mortgage Rate

Variable tested: *rater*		LM statistic
Kwiatkowski–Phillips–Schmidt–Shin test statistic		0.840683
Asymptotic critical values*	1 per cent level	0.739
	5 per cent level	0.463
	10 per cent level	0.347

*KPSS, Table 1, p. 166.