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Abstract

This article models the dynamics of house prices in Australian capital cities using cointegration and causality testing. The approach employed in the article follows fairly closely that used in previous studies by MacDonald and Taylor (1993) and Alexander and Barrow (1994) to analyze regional house prices in the United Kingdom. The main result is that although there is some weak evidence of market segmentation in house prices in the South East and East Coast of Australia, there are few cointegrating relationships between capital cities. In terms of causality, the results suggest that housing prices in Melbourne and Adelaide are Granger causing housing prices in Canberra and housing prices in Perth and Sydney are Granger causing housing prices in Brisbane.

Introduction

The introduction of the Goods and Services Tax, together with the first home owners grant and successive interest rate cuts have made movements in house prices a topic of major public interest (see for example Gordon 2001, 2001a, 2001b). Previous Australian studies have modelled the demand for house ownership (Bourassa 1995), investment in rental housing (Beer 1999; Berry 2000) and the determinants of house prices (Abelson 1994; Bourassa & Hendershott 1995; Reed 2001). This article seeks to address the gap in studies that have examined the dynamics of house prices in Australian capital cities using the techniques of cointegration and causality testing.

The approach used in this article follows fairly closely that employed in previous research by MacDonald and Taylor (1993) and Alexander and Barrow (1994) in their studies of regional house prices in the United Kingdom (UK). A more recent study by O'Donovan and Rae (1997) also addresses some of the issues in this article for regional house prices in New Zealand. This article addresses the questions: (a) is there a long-run equilibrium relationship in housing prices in Australian capital cities? (b) Is there a segmentation of housing prices in Australia, such as North-South or East-West? (c) If there are stable long-run relationships between housing prices between cities, which capital cities are the price setters (in the sense of leading price movements in other cities)?

Movements in house prices are important for several reasons. First, the pattern of house prices is important for the efficient operation of the labour market. The operation of the labour market might be impaired if labour mobility is hampered by high housing costs in one or more capital cities (Bover et al. 1989). Second, housing prices have a significant effect on the distribution of wealth, since housing is the major asset of most households (Alexander & Barrow 1994). Third, stamp duty, charged as a percentage of the house price, is an important source of state government revenue (Baker 2002). Fourth, movements in housing prices have tax implications for individuals, with housing being a tax-effective form of saving. This reflects the fact that investment growth on owner occupied housing receives favourable income tax treatment (Yates 1997; Freebairn 1999).

Housing is immobile; accordingly economic theory suggests that regional prices will not necessarily move together over time. Migration may be one means through which prices converge. However, as Alexander and Barrow (1994) note, migration occurs in response to employment opportunities and other reasons, rather than relative house prices. Therefore, it is unclear whether migration flows alone are sufficient to explain relative house price movements. To foreshadow the main results, we find relatively few cointegrating relationships between capital cities, with some, albeit weak, evidence of market segmentation in house prices in the South East and on the East Coast.

Data and Methodology

Quarterly data on established real house prices compiled by the Australian Bureau of Statistics (ABS) for Canberra, Darwin and each of the state capital cities from 1986 Q2 to 2001 Q2, with base 1989-1990=100 (ABS various) was used. Following MacDonald and Taylor (1993) and Alexander and Barrow (1994), the analysis below was conducted using the logarithm of each housing price series. Figures 1 and 2 show movements in house prices for each of the cities over this period. Figure 1 uses 1989-90 as the base, which is consistent with the ABS. The data in figure 2 has been converted so that the start of the period for the study (1986 Q2) is treated as the base. The largest price increases between 1986 Q2 and 2001 Q2 were experienced in Sydney (215 percent), followed by Melbourne (173 percent), Perth (142 percent) and Brisbane (135 percent).

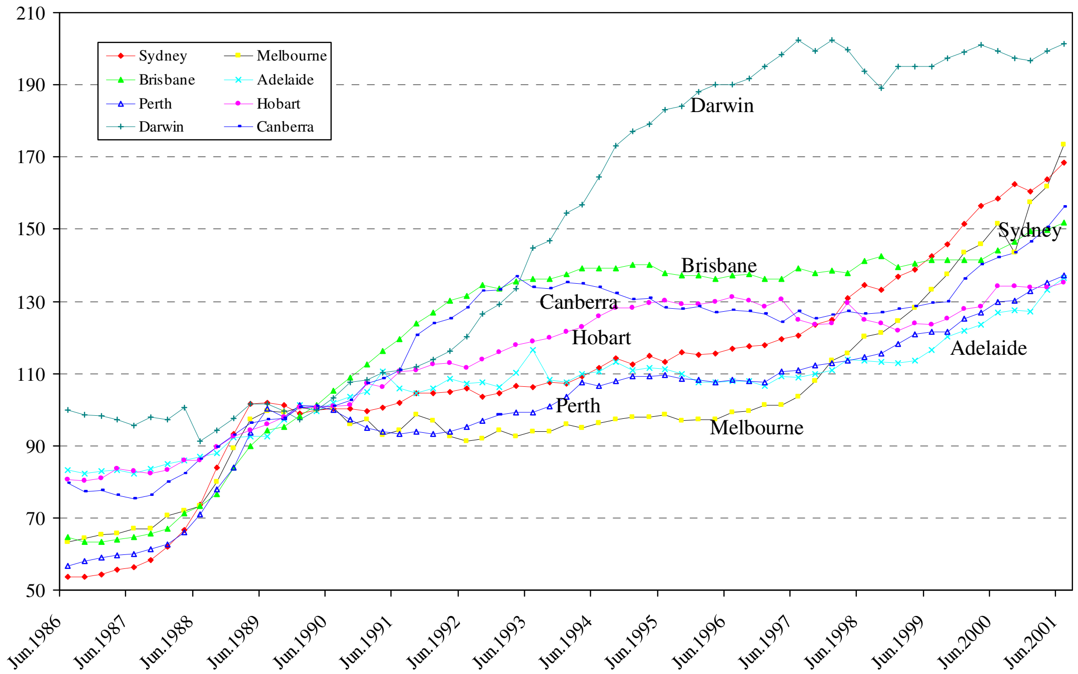

Figure 1: Australian Capital City House Price Index Numbers (Base 1989-90 = 100)

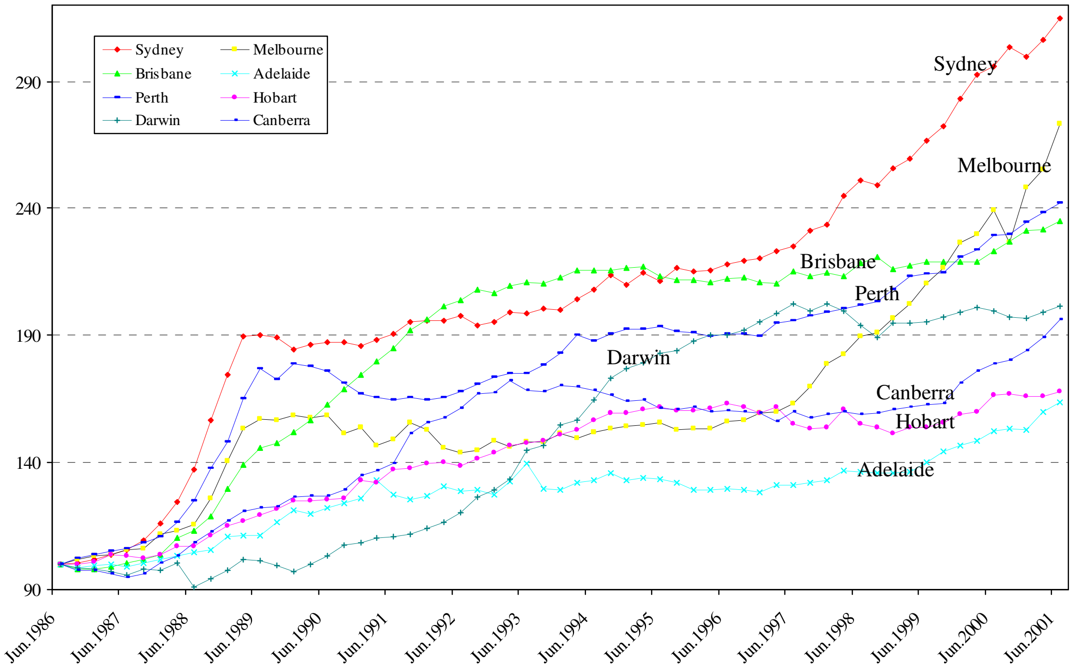

Figure 2: Australian Capital City House Price Index Number (Base June 1986 = 100)

The first step is to check for the presence of unit roots in the house price series for each of the capital cities. For those cities which have house prices that are I(1) processes, the next step is to determine if their house price series are cointegrated. We will use the Johansen method of cointegration. There are two Johansen cointegration tests. The trace test, tests the null hypothesis that there are at most, *r* cointegrating vectors. The maximum eigenvalue test, tests the null hypothesis of *r* cointegrating vectors against the alternative hypothesis of (*r*+1) cointegrating vectors. If a vector of house prices in capital cities are cointegrated, a vector error correction model can be used to investigate how shocks to house prices in one capital city affect house prices in other capital cities, with which it is cointegrated. However, if house prices in two capital cities are each I(1), but not cointegrated, the difference between house prices in these cities will tend to grow larger over time with no tendency towards equilibrium (MacDonald & Taylor 1993).

Empirical Results

Time Series Properties of House Prices in Australian Capital Cities

Table 1 reports the results of the unit root tests for housing prices in each of the capital cities. Both the Augmented Dickey-Fuller (ADF) and the Phillips-Perron (PP) tests were conducted. With the exception of Sydney and Brisbane, both of these tests suggest that housing prices are non-stationary in levels, but stationary in first differences at the 5 percent level. The ADF test fails to reject the null of non-stationarity in first differences for Brisbane and only rejects the null of non-stationarity in first differences for Sydney at 10 percent. Where there is ambiguity, as in the case of Brisbane and Sydney, the PP tests are preferred, because of their generally greater power (see Banerjee et al. 1993, p. 113). Thus we proceed on the basis that housing prices in each of the capital cities are I(1).

Table 1: Tests for Stationarity

CITY	ADF[1]		PP[2]		Order of Integration
	Level	*First Difference*	Level	*First Difference*	
Hobart	-1.83 (3)	-3.85* (2)	-1.06	-8.46*	I(1)
Melbourne	-0.94 (1)	-3.32** (2)	-1.08	-5.72*	I(1)
Sydney	-3.14 (1)	-2.74***(6)	-2.17	-3.51**	I(1)
Brisbane	-1.94 (1)	-1.65 (5)	-2.73	-4.48*	I(1)
Darwin	-0.91 (3)	-3.79* (2)	-1.18	-6.58*	I(1)
Perth	-2.72 (1)	-2.93** (4)	-2.21	-3.73**	I(1)
Adelaide	-1.87 (4)	-3.38** (3)	-1.41	-8.14*	I(1)
Canberra	-1.66 (1)	-3.24** (1)	-1.29	-5.09*	I(1)

Notes:

* Indicates calculated test statistic is significant at 1 percent
** Indicates calculated test statistic is significant at 5 percent
*** Indicates calculated test statistic is significant at 10 percent

1. Constant and trend were included in level, but only constant in first difference (see Baghestani & Mott 1997). The number of lags, which are given in parenthesis, were chosen so as to minimize Akaike's Information Criterion.
2. The truncation lag is 3, based on the Newey-West Suggestion.
3. Critical values are from MacKinnon (1991).

Tests for the Long-run Stability in Capital City Housing Prices

Once it is established that housing prices are I(1) series, Johansen's multivariate cointegration tests can be used to establish the number of distinct cointegrating vectors. Table 2 presents estimates of the number of cointegrating vector for the capital cities. The lag length, which was 3, was chosen so as to minimize Akaike's Information Criterion (AIC). The AIC is a standard method of selecting lag length. Consider a regression model with k parameters, estimated using T observations. Let l be the log of the likelihood function using the k estimated parameters. The AIC is $-2(l/T) + 2(k/T)$.

The VARs included an unrestricted constant and seasonal dummies as in Johansen and Juselius (1990) and McDonald and Taylor (1993). We used centred (orthogonalized) seasonal dummies. Using standard 0-1 seasonal dummies in the test VAR will affect both the mean and the trend of the level series. Johansen (1995, p. 84) suggests that using centred seasonal dummies avoids this problem through shifting the mean without contributing to the trend. The critical values are taken from MacKinnon et al. (1999) who provide approximate critical values which take into account exogenous variables.

Table 2: Multivariate Johanson's Cointegration Tests

Hypothesized No. of CE(s)	λ_{trace} Trace Statistic	5 percent CV	λ_{max} Max-Eigen Statistic	5 percent CV
None	376.44	214.15	116.77	62.50
At most 1	259.67	174.26	84.75	56.38
At most 2	174.92	138.30	57.30	50.20
At most 3	117.62	106.36	47.99	43.98
At most 4	69.64	78.34	32.62	37.67
At most 5	37.02	54.27	19.62	31.24
At most 6	17.39	33.98	16.83	24.54
At most 7	0.56	17.23	0.56	17.23

Notes:

The vector autoregressions included an unrestricted constant and orthogonalized seasonal dummies. The lag length was 3, which was chosen so as to minimize Akaike's Information Criterion. The critical values are from MacKinnon et al. (1999). Both trace and maximum eigenvalue tests indicate 4 cointegrating vectors at the 5 percent level.

Table 2 shows that both trace and maximum eigenvalue tests indicate four cointegrating vectors at the 5 percent level. These results are in sharp contrast with those of MacDonald and Taylor (1993), who find up to nine cointegrating relations amongst the 11 regions of the UK including Wales for the period 1969 to 1987. However, MacDonald and Taylor (1993) use Engle and Granger's (1987) methodology, which is impaired by small sample bias. The findings in this study are more consistent with Alexander and Barrow (1994) and O'Donovan and Rae (1997) who also use the Johansen procedure for cointegration. O'Donovan and Rae (1997) found no clear evidence that house prices were cointegrated across 14 regions of New Zealand. Alexander and Barrow (1994) found four cointegrating vectors amongst nine regions in the UK over a slightly longer period (1968 to 1993) than MacDonald and Taylor (1993) and also using a shorter lag in the VAR.

With eight capital cities and several cointegrating vectors it is difficult to distil meaningful causal flows in an error correction model. Thus, Tables 3 and 4 present the results of bivariate trace and maximum eigenvalue tests for each pair of cities. As in the multivariate cointegration tests, the lag length was chosen to minimize the AIC. The bivariate cointegration trace tests indicate the presence of only two cointegration relations (Brisbane-Sydney and Brisbane-Perth). The maximum eigenvalue tests, however, suggest the presence of four cointegrating relations (Melbourne-Canberra, Brisbane-Sydney, Adelaide-Canberra and Brisbane-Perth). Of the two tests, the maximum eigenvalue test is a more powerful likelihood ratio test. Johansen and Juselius (1990) suggest that where the tests give conflicting results the maximum eigenvalue test should be preferred. Maddala and Kim (1998) proffer a similar view. Thus the analysis below proceeds on the basis that the results for the bivariate maximum eigenvalue tests are to be preferred over those of the bivariate trace tests.

Table 3: Bivariate Cointegration Trace Tests

	MELBOURNE	SYDNEY	CANBERRA	BRISBANE	DARWIN	PERTH	ADELAIDE	HOBART
SYDNEY	12.96 (2)	_						
CANBERRA	24.66 (4)	21.24 (5)	_					
BRISBANE	24.03 (2)	35.26* (5)	21.37 (3)	_				
DARWIN	3.40 (3)	20.42 (5)	25.08 (4)	23.81 (2)	_			
PERTH	11.82 (2)	16.55 (4)	22.18 (5)	36.22* (5)	14.69 (3)	_		
ADELAIDE	14.19 (1)	18.17 (2)	27.95 (4)	18.44 (3)	6.94 (1)	15.69 (3)	_	

HOBART	9.24 (1)	25.84 (4)	14.69 (1)	18.97 (2)	23.41 (1)	13.69 (2)	12.00 (1)	_

Notes:

* Indicates calculated statistic is significant at 5 percent or better

Test assumptions: The vector autoregressions included an unrestricted constant and orthogonalized seasonal dummies. The number of lags, which are given in parenthesis, were chosen so as to minimize Akaike's Information Criterion. The critical value at 5 percent is 33.98 according to MacKinnon et al. (1999).

Table 4: Bivariate Cointegration Maximum Eigenvalue Tests

	MELBOURNE	SYDNEY	CANBERRA	BRISBANE	DARWIN	PERTH	ADELAIDE	HOBART
SYDNEY	10.87 (2)	_						
CANBERRA	24.62* (4)	16.61 (5)	_					
BRISBANE	13.65 (2)	30.99* (5	18.73 (3)	_				
DARWIN	3.18 (3)	14.31 (5)	20.38 (4)	13.58 (2)	_			
PERTH	8.86 (2)	13.85 (4)	16.55 (5)	30.60* (5)	12.47 (3)	_		
ADELAIDE	14.18 (1)	16.76 (2)	27.88*(4)	16.24 (3)	6.53 (1)	15.57 (3)	_	
HOBART	9.08 (1)	12.52 (4)	13.30 (1)	12.60 (2)	14.8 (1)	8.00 (2)	12.00	_

Notes:

* Indicates calculated statistic is significant at 5 percent or better

Test assumptions: The vector autoregressions included an unrestricted constant and three orthogonalized seasonal dummies. The number of lags, which are given in parenthesis, were chosen so as to minimize Akaike's Information Criterion. The critical value at 5 percent is 24.54 according to MacKinnon et al. (1999).

The results provide some evidence of cointegration between cities on the East Coast and in the South East, with cointegrating relations between Melbourne-Canberra, Brisbane-Sydney and Adelaide-Canberra. However, in the South East, while Melbourne-Canberra and Adelaide-Canberra are cointegrated, Melbourne is not cointegrated with Adelaide. Formally this means that Z_1=ADEL + a_1MELB and Z_2=ADEL +a_2CAN are both stationary, but no linear combination of Z_1 and Z_2 is stationary. This is possible if, for example, Z_1 and Z_2 move counter-cyclically (Alexander & Barrow 1994, p. 1678). Alexander and Barrow (1994) found a similar phenomenon in the south of England, where house prices in East Anglia were cointegrated with both the South East and South West Regions, but these two regions themselves were not cointegrated.

Causality Tests Based on the Error Correction Model

The next step is to establish the direction of causality between housing prices in the capital cities which are cointegrated. To do this, we employ Granger causality. Consider the following equations involving two variables *x* and *y* where *j* represents the lag length:

$$y_t = \alpha_0 + \alpha_1 y_{t-1} + \ldots\ldots + \alpha_j y_{t-j} + \beta_1 x_{t-1} + \ldots\ldots + \beta_j x_{t-j} + \varepsilon_t \quad (1)$$
$$x_t = \alpha_0 + \alpha_1 x_{t-1} + \ldots\ldots + \alpha_j x_{t-j} + \beta_1 y_{t-1} + \ldots\ldots + \beta_j y_{t-j} + \mu_t \quad (2)$$

In the Granger sense, considering whether *x* causes *y* involves seeing how much of the current value of *y* can be explained by past values of *y* and then seeing whether adding lagged values of *x* can improve the explanation. *x* is said to Granger cause *y* if *x* helps in the prediction of *y* or if the coefficients on the lagged *x*'s are statistically significant

(Granger 1969). The null hypothesis in equation (1) is that *x* does not Granger cause *y* and the null hypothesis in equation (2) is that *y* does not Granger cause *x*.

If two series are I(1), Granger causality must exist in one or more directions in, at least, the I(0) variables (Engle & Granger, 1987). However, standard tests for causality (such as Granger/Sims tests) are only valid if the housing price time series are not cointegrated. If house prices in two cities are cointegrated, any causal inference will not be valid (Granger 1988). To make the process valid, the error-correction term obtained from the cointegration regression has to be included in the standard causality tests (Giles et al. 1993, Oxley 1993). Thus, the long-run cointegrating relations found from the bivariate maximum eigenvalue tests were used in Error Correction Models to determine the direction(s) of Granger causality. The results of a Wald statistic are presented in Table 5. For pairs of (*x, y*) series, the Wald test statistic, which follows a chi-square distribution, tests the joint hypothesis that $\beta_1 = \beta_2 = \ldots \beta_j = 0$ in equations (1) and (2).

Table 5: Granger Causality Tests Based on the Error Correction Model

	Wald test	d.f	Causal?
Melbourne to Canberra	14.59*	4	Yes
Canberra to Melbourne	1.56	4	No
Sydney to Brisbane	44.90*	5	Yes
Brisbane to Sydney	7.04	5	No
Adelaide to Canberra	13.43*	4	Yes
Canberra to Adelaide	4.73	4	No
Brisbane to Perth	3.98	5	No
Perth to Brisbane	19.77*	5	Yes

Notes:
* Indicates calculated test statistic is significant at 1 percent.

The conclusions that emerge from Table 5 are that there are strong positive flows (at the 1 percent level) from Melbourne to Canberra, Adelaide to Canberra, Sydney to Brisbane and Perth to Brisbane. These results suggest that Melbourne and Adelaide are price leaders with respect to Canberra and that Perth and Sydney are price leaders with respect to Brisbane. An important reason for these causal relationships could be relative price increases coupled with population flows. Figures 3 to 6 show comparative price increases in each of the cities where prices are cointegrated. Over the period, Melbourne and Sydney experienced the biggest increase in house prices. At the same time, New South Wales, together with South Australia, experienced negative net migration throughout the entire 1990s, while Victoria experienced large negative net migration flows until 1997-98 (ABS 1999-2000). Queensland has experienced the largest net migration gains of any state throughout the period of the study (see ABS 1997). Thomas (1993) examines the effect of migration decisions on house prices in the UK. He found that households moving for job reasons were on account of wage differences rather than house prices, while those moving for other reasons, such as retirement, do consider house prices. Thus, Thomas (1993) argues that retirees are more important in influencing house prices.

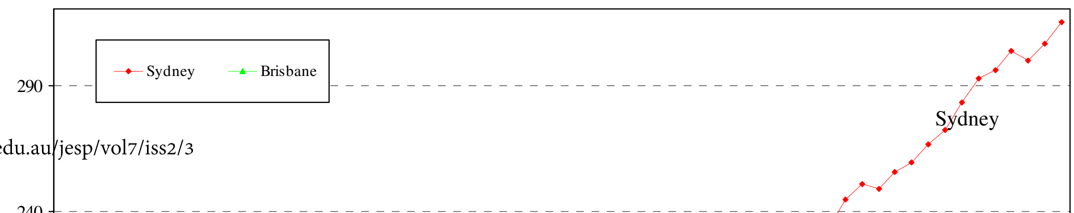

Figure 3: Housing Prices in Sydney and Brisbane (Base June 1986 = 100)

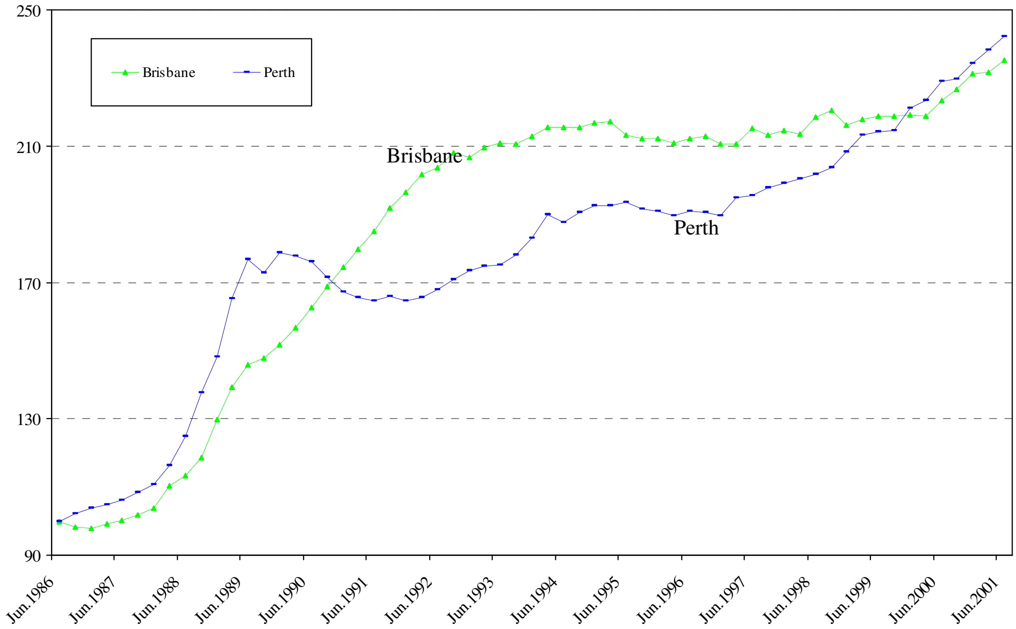

Figure 4: Housing Prices in Perth and Brisbane (Base June 1986 = 100)

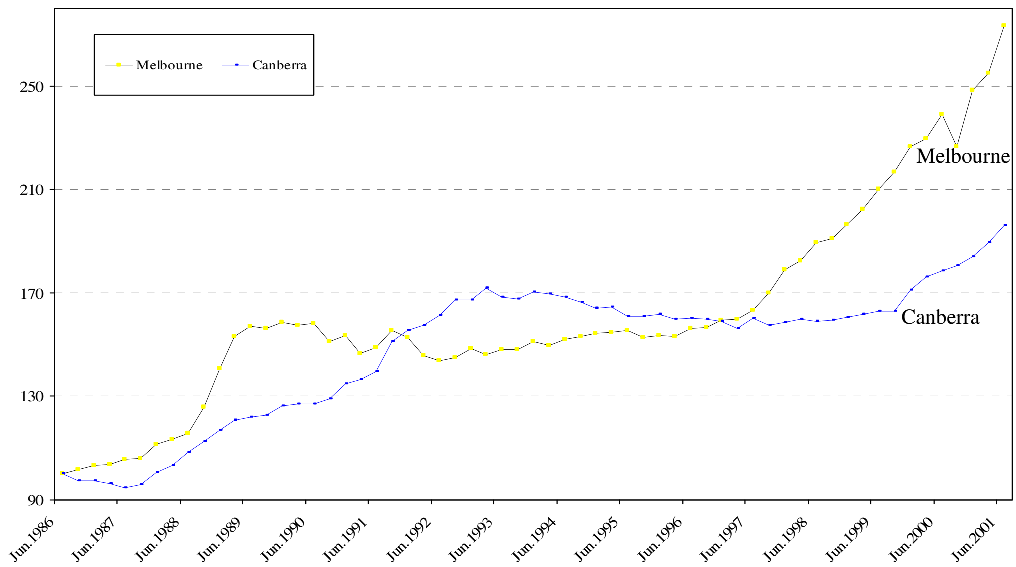

Figure 5: Housing Prices in Melbourne and Canberra (Base June 1986 = 100

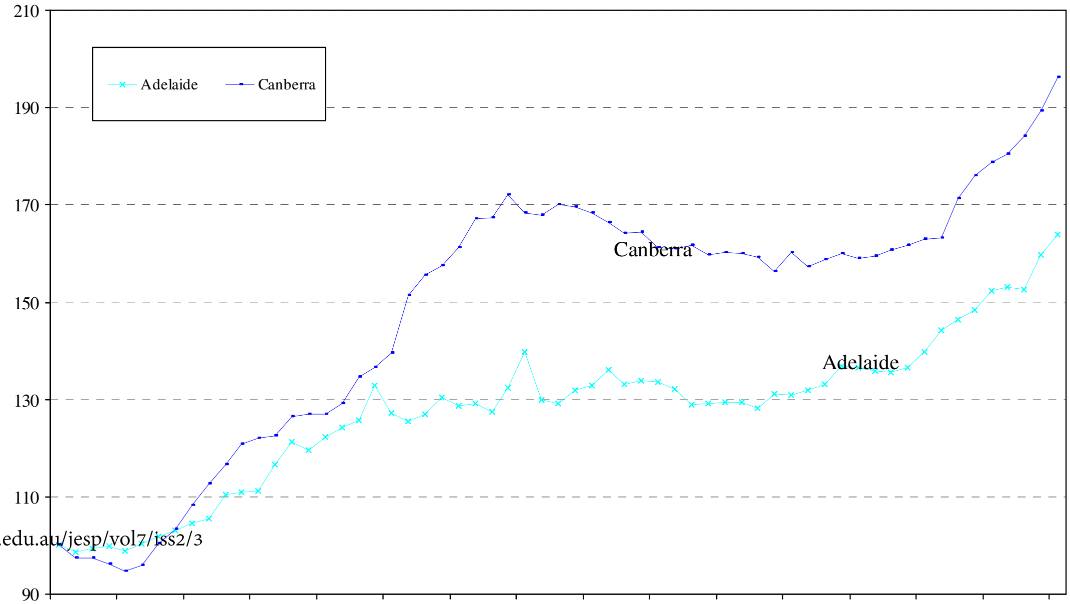

Figure 6: Housing Prices in Adelaide and Canberra with (Base June 1986 = 100)

In the Australian context, we suspect that this is true in explaining the results for Brisbane. Queensland is a popular destination for self-funded retirees from other states, who might see rising house prices in their home states as a catalyst to act. While, perhaps surprisingly, Melbourne and Brisbane house prices are not cointegrated, the highest interstate flows for most of the period were from New South Wales to Queensland, which helps to explain the high value of the Wald test for causation running from Sydney to Brisbane. The second highest population flows at the end of the 1990s were in the reverse direction from Queensland to Sydney (ABS 1999-2000). While there is no evidence of reverse causality from Brisbane to Sydney, high population flows in both directions are consistent with the cointegrating vector between the two cities.

Canberra has experienced relatively large population flows throughout most of the 1990s, but because inflows were roughly the same as outflows net migration was small. Outflows may be due largely to school leavers, who move from Canberra in search of employment opportunities. At the same time, there is an inflow of university graduates from other capital cities, such as Adelaide and Melbourne, primarily to take up positions in the Commonwealth Public Service. Consistent with the findings of Thomas (1993) for the UK, it is likely that job opportunities are the major motive behind migration decisions for these people. However, at least for some who see the Commonwealth Public Service as a long-term career, relatively low house prices in Canberra could be a further incentive.

Conclusions

This article models the dynamics of house prices in Australian capital cities using cointegration and causality testing. The approach follows studies for the UK by MacDonald and Taylor (1993) and Alexander and Barrow (1994). The main result is that there are few cointegrating relationships between capital cities, which is consistent with the findings of Alexander and Barrow (1994) for the UK and O'Donovan and Rae (1997) for New Zealand. There is weak evidence of market segmentation in house prices in the South East and East Coast of Australia. The results suggest that housing prices in Melbourne and Adelaide are Granger causing housing prices in Canberra and housing prices in Perth and Sydney are Granger causing housing prices in Brisbane.

The study has limitations, which need to be considered when interpreting the results. First, in the last section, we have speculated about the reasons for our findings, focusing on population flows between the cities being based on job or retirement purposes. One limitation of the article is that we have analyzed housing prices largely in what MacDonald and Taylor (1993) term 'black box' statistical terms. To obtain a better understanding of the factors driving the relationships between housing prices, a more structural model is required where, for example, housing price differences between cities are regressed on other economic variables such as wage rates. Thus our results should be viewed in conjunction with studies such as Abelson (1994) and Bourassa and Hendershott (1995), which specifically examine the determinants of housing prices.

A second limitation is that we consider linkages between capital cities, but do not consider linkages between other major cities or within states. The dynamics of local house prices within selected metropolitan areas is one potential line of future research. This would be in the spirit of studies such as Roehner (1999) which examine movements in housing prices using district level data for specific cities. Moreover, while we have speculated about the connection between movements in housing prices and population flows we have not modelled this relationship. To shed light on this issue, future research may examine bi-directional causality between house prices and migration as suggested by Alexander and Barrow (1993, p. 1686). A further avenue for future research could be to explicitly explore the relationship between housing prices, job opportunities, wage rates and migration flows in the Australian context similar to the studies which Bover et al. (1989) and Thomas (1993) have conducted for the UK.

Acknowledgements

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