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Synchronisation and Commonalities in Metropolitan Housing Market Cycles

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Abstract

This paper examines the degree of commonalities present in the cyclical behavior of the eight largest metropolitan housing markets in Australia. Using two techniques originally in the business cycle literature we consider the degree of synchronization present and secondly decompose the series' into their permanent and cyclical components. Both empirical approaches reveal similar results. Sydney and Melbourne are closely related to each other and are relatively segmented from the smaller metropolitan areas. In contrast, there is substantial evidence of commonalities in the cyclical behavior of the remaining cities, especially those on the Eastern and Southern coasts of Australia.

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This paper examines the degree of commonalities present in the cyclical behavior of the eight largest metropolitan housing markets in Australia. Using two techniques originally in the business cycle literature we consider the degree of synchronization present and secondly decompose the series' into their permanent and cyclical components. Both empirical approaches reveal similar results. Sydney and Melbourne are closely related to each other and are relatively segmented from the smaller metropolitan areas. In contrast, there is substantial evidence of commonalities in the cyclical behavior of the remaining cities, especially those on the Eastern and Southern coasts of Australia.

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1: Introduction

Over the course of the last two decades a large literature has developed to have considered the interaction and relationships present amongst either metropolitan or regional housing markets. In the main this has considered the issue from the perspective of house price diffusion and the analysis of whether causal relationships exist. This literature is particularly prevalent in the UK where considerable research has been conducted examining the *ripple effect* which considers whether house prices movements in London and South East of England impact upon subsequent market behavior in the rest of the UK (e.g. Meen, 1999; Cook, 2003; Holly et al., 2011). This paper contributes to the literature by complementing the existing work on house price diffusion through the adoption of an alternative methodological framework in the context of eight metropolitan areas in Australia. We consider the capitals of Australia's six states, namely; Adelaide (South Australia), Brisbane (Queensland), Hobart (Tasmania), Melbourne (Victoria), Perth (Western Australia) and Sydney (New South Wales). In addition to the six state capitals we also analyse Canberra (Australian Capital Territory) and Darwin (Northern Territory). The case of Australia provides an interesting counterpoint to the studies of the UK and US. Whilst smaller in population than the UK, the geographic size of Australia is similar to the US. The net result is a small number of metropolitan areas that are separated by considerable distances. Therefore, the degree to which they display similarities in their cyclical behaviour is of interest.

The paper considers the degree to which the primary metropolitan markets display characteristics that indicate the presence of common cycles. Two alternative methodological approaches are utilized in this study. The first considers the degree of synchronization between the metropolitan markets using the modified Concordance Indicator of Harding & Pagan (2006). This approach estimates the degree to which two markets are synchronised in terms of the phase of their cycle, i.e. house price appreciation or depreciation. This approach therefore provides a compliment to the conventional comparative analysis of markets. The second approach is also based upon the business cycle literature and decomposes the housing data examined into their trend and cyclical components. Two alternative decomposition approaches are considered, namely those of Beveridge-Nelson (1981) and Hodrick-Prescott (1997). The remainder of the paper is structured as follows. Section 2 discusses the relevant literature pertaining to the inter-linkages between housing markets. Section 3 provides information concerning the data utilized in the paper. Sections 4 and 5 present and report upon the empirical findings, whilst concluding comments are made in Section 6.

2: Literature Review

The literature to have considered the interactions amongst housing markets has largely done so from the context of examining house price diffusion. A large proportion of this literature has investigated either the UK or US and to some degree, and of obvious interest in the context of the current paper, Australia[1]. The UK literature has often specifically considered the *ripple effect*. Meen & Andrew (1998) highlight five factors that may contribute to the presence of a ripple effect in the UK, namely; migration, transaction and search costs, equity transfer, spatial arbitrage and leads and lags in house prices. The majority of the earlier studies relied heavily upon a causality framework. For example, Giussani & Hadijmatheou (1991) and MacDonald & Taylor (1993) both report evidence supportive of the ripple effect with London as the base region. Whilst reporting broadly similar findings, the paper of Alexander & Barrow (1994) extends the analysis in two respects. Firstly, it uses the more robust Vector Error Correction Mechanism (VECM) framework. Secondly, rather than base their analysis on the premise of London being the base region, the paper considers the surrounding South East as an alternative, finding that it is actually a more appropriate base[2]. Muellbauer & Murphy (1994) report complementary evidence in this report, noting that regions contiguous to the South East of England are affected not only by house price movements but also by income in the region. This can be taken as supportive of the role of spatial lags in the ripple effect.

In addition to the tests for causality a number of papers have considered whether UK regions are cointegrated, i.e., if they share a common long-term trend. MacDonald & Taylor (1993) use the bivariate Engle-Granger cointegration test, reporting significant results with respect to pairings of southern and non-southern regions[3]. Cook (2005a) expands upon these tests through the adoption of cointegration tests that allow for asymmetric adjustment. The findings reported indicate that when house prices in the South of England decline relative to other regions then reversion to equilibrium occurs quite rapidly. However, when the reverse scenario is considered, i.e. prices in the south increase on a relative basis, the degree of reversion to equilibrium observed is slower[4].

Papers in the last decade have however taken different methodological approaches in the examination of diffusion and the inter-linkages across markets. Following the observation of Meen (1999), that if the ratio of regional house prices to the overall national figure exhibits evidence of stationarity then this implies long-term convergence, a number of papers have used unit root tests to consider the issue of convergence. Two papers by Cook (2003, 2005b) test for stationarity using a variety of unit root approaches. Cook (2003) considers an asymmetric unit root specification, whilst Cook (2005b) uses the Generalised Least Squares variation of the Augmented Dickey-Fuller unit root test, as proposed by Elliot et al. (1996). The results in both papers provide evidence of convergence. In the case of Cook (2005b) significant results are reported with respect to six UK regions (North, North West, East Anglia, South East, Wales and Northern Ireland). Holmes (2007) considers the issue of stationarity in a panel setting, this complementing the work of Cook (2005b). The results indicate that converging behaviour is present in the UK regional markets[5]. Holmes & Grimes (2008) also consider stationarity but in a slightly different context in that they firstly use principal components analysis to identify the linear combination of the regional house price series that captures the highest degree of variation across the series. They then test for stationarity in this first principal component. Holmes & Grimes (2008) find evidence of stationarity, indicating that UK regional house prices have a single common stochastic trend. Holly et al. (2011) show that London's global role adds an international element to house price diffusion in the UK. Whilst the results support the previously observed ripple effect, it is also noted that London is significantly linked to other global cities, in this case New York. The modeling approach adopted by Holy et al. (2011) allows it to be observed that whilst a shock to London dissipates relatively quickly (two years), the impact of such a shock to other UK regions is not only extended in a temporal sense but varies depending upon the spatial distance of the region to London.

In contrast to the UK, where the literature has largely been concerned with regional housing markets, much of the international literature has studied either metropolitan or sub-market data. In the US the early house price diffusion literature generally concentrated on diffusion between neighbouring markets, often findings results highlighting the importance of geographic proximity. (e.g. Clapp & Tirtiroglu, 1994 and Pollakowski & Ray, 1997)[6]. The divergence in findings between contiguous and non-contiguous markets is often attributed to factors such as the transfer of information and a positive feedback effect, whereby positive or negative movements in one market have a knock-on effect in neighbouring markets. A recent paper by Gupta & Miller (2012) consider the issue of diffusion in the case of eight metropolitan markets in Southern California, reporting substantial evidence of cointegration and causal relations across the various metropolitan markets[7].

As with many global housing markets a number of papers have recently examined the dynamics of the Australian market and in particularly the degree to which speculative behavior has possibly developed (e.g. Hatzvi & Otto 2008; Fry et al., 2010) [8]. Costello et al. (2011) not only consider the degree of divergence from prices that can be justified according to fundamentals, but also the regional variation in such behaviour. Costello et al. (2011) note that the degree of divergence from fundamentals differs across Australian states, for example finding that whilst some states, such as Victoria, have largely seen prices in line with fundamentals since 2005, others have not. In addition, the paper considers the spillover effect of 'non-fundamental prices'. As with their initial analysis they report differences across states, with house prices in New South Wales most vulnerable to non-fundamental, or speculative, spillover effects. In more conventional tests both Tu (2000) and Luo et al. (2007) consider the degree of house price diffusion present. Both papers note a number of significant results with respect to pairings of Australian markets being cointegrated. In addition, evidence of diffusion in a Granger Causality sense is also noted. This is especially evident when Sydney and Melbourne are considered. Luo et al. (2007) provide evidence that there is a distinct diffusion impact, with house price changes originating in Sydney then descending through Melbourne and subsequently to other markets. Evidence of cointegration, in a bilateral context, between a large number of Australian markets is reported. However, it would appear that Sydney, and to a lesser degree Melbourne, are again separated from the other metropolitan markets. Whilst a large number of significant results were noted, there was a marked reduction in the number when Sydney and Melbourne were examined. Sydney was only found to be cointegrated with Melbourne, whilst Melbourne added Adelaide and Perth. This can be taken as being supportive of a diffusion effect, similar to that observed in the UK, with Sydney, and then Melbourne, as the base regions[9].

3: Data

The data used in this study consists of the quarterly Australian Bureau of Statistics (ABS) indices for the eight Australian capital cities, namely; Adelaide (South Australia), Brisbane

(Queensland), Canberra (ACT), Darwin (Northern Territory), Hobart (Tasmania), Melbourne (Victoria), Perth (Western Australia) and Sydney (New South Wales). The indices are weighted averages based upon a stratified clustering approach. The data analysed covers the period June 1986 to December 2010. The indices were re-estimated in 2005 using updated weights. These weights were used to backdate the series' to 2002. The original series is then incorporated into the new series based upon the quarterly percentage changes to backdate the data to 1986. Figure 1 displays the constructed index series for the different markets and for the overall index, whilst the summary statistics for reported in Table 1.

It can be seen from Figure 1 and Table 1 that while the different markets display broad similarities in terms of their cyclical behaviour there are distinct differences also evident. Adelaide displays both a lower average quarterly return and standard deviation than the other metropolitan markets, whilst at the other extreme the city that displays both the highest return and volatility is Melbourne. In addition, the relative performance of the cities does diverge in the post 2002 period. In particular, Sydney has observed far lower price appreciation than the other markets, indeed the strongest performing markets over the course of the last decade are the smaller secondary markets such as Darwin. Table 1 also reports tests of stationarity, based upon the Augmented Dickey-Fuller test. In each case the first differenced return series is stationary.

4: Synchronisation of Cycles

In order to consider the degree of synchronisation present in the markets considered we adopt the concordance indicator proposed by Harding & Pagan (2001, 2002, 2006) and which has been utilised in a large number of papers that have considered business cycles (e.g. Altavilla, 2004, Harding & Pagan, 2001, 2002) and also in a recent paper considering the commercial office market (Jackson et al., 2008). The methodology defines state variables that consider whether a market is in a state of expansion or contraction. Harding and Pagan (2002) propose a non-parametric approach to estimating the level of concordance between two series. The growth rates are expressed as two binary random variables, S_{it} and S_{jt}, which are the state variables for cycles for markets i and j. The state variables are defined as dummy variables equalling unity when the cycle is on an upward trend and zero otherwise. Using these two state variables, the index of concordance between two cities indicates the proportion of time two cycles spend in the same phase. The concordance index can be estimated as follows:

$$IC = T^{-1}\sum_{t=1}^{T}\left[S_{jt}S_{it} + (1 - S_{jt})(1 - S_{it})\right] \tag{1}$$

This statistic can also be adapted in what has been referred to as the Mean Corrected Index of Concordance. This adaptation, proposed by Harding & Pagan (2001), is designed to adjust the initial indicator for potential biases. Harding & Pagan (2001) noted that the original IC measure might be overstated in the case of two variables that experience prolonged expansion during the period of study. Prolonged growth over a number of consecutive periods is a common feature of real estate and economic cycles' data. Therefore, the Mean Corrected Measure of IC (MCIC) is proposed under the assumption of no relation between two series. In comparison with the original IC statistic, the MCIC measures the proportion of time that two series are expected to share in the same phase under an assumption of independence. The adapted MCIC measure is as follows:

$$MCIC = 2T^{-1}\sum_{t=1}^{T}\left[(S_{it} - \bar{S}_i)(S_{jt} - \bar{S}_j)\right] \tag{2}$$

Where:

$$\bar{S}_i = T^{-1}\sum_{t=1}^{T}S_{it} \tag{3}$$

$$\bar{S}_j = T^{-1}\sum_{t=1}^{T}S_{jt} \tag{4}$$

However, both concordance measures can be difficult to assess and interpret. The Mean Corrected Index of Concordance is unlikely to exceed 0.5, whilst the assumption of independence is a strong assumption to make. The original IC values lie within the interval [0, 1], where 1 implies perfect synchronization. In this case, the value of 0.5 would mean no particular relation between two series. However, the values that exceed 0.5 cannot be interpreted as statistically meaningful based on the index value information. To overcome such limitations, Harding and Pagan (2006) propose an alternative mean-corrected measure of concordance ($\hat{I}_t$), which also allows one to draw inferences about the concordance index values.

Harding and Pagan (2006) show that $\hat{I}_t$ and the empirical correlation between two series ($\hat{\rho}_S$) are monotonically related and the significance of $\hat{\rho}_S$ implies significance of $\hat{I}_t$. They express the revised concordance index as follows:

$$\hat{I}_t = 1 + 2\hat{\rho}_s \sigma_{s_x} \sigma_{s_y} + 2\mu_{s_x}\mu_{s_y} - \mu_{s_x} - \mu_{s_y} \tag{5}$$

where μ_{s_i} and σ_{s_i} are the average and standard deviation of the state variables S_i ($i=x,y$) and $\hat{\rho}_S$ is the correlation between S_{xt} and S_{yt}. The value of $\hat{\rho}_S$ and inferences concerning it can be derived using the following OLS regression:

$$\frac{S_{y_t}}{\sigma_{s_x}\sigma_{s_y}} = \hat{\alpha} + \hat{\rho}_s \frac{S_{x_t}}{\sigma_{s_x}\sigma_{s_y}} + \varepsilon_t \tag{6}$$

In order to control for positive serial correlation in S_{yt}, the $\hat{\rho}_S$ test-statistics are estimated using robust standard errors obtained via the HAC procedure. Harding and Pagan (2006) also note that the alternative estimation of the index, via the $\hat{\rho}_S$, provides an alternative mean-corrected measure of concordance. Since the assumption is that we measure the concordance of two independent series, the regression helps us to identify which relations between two series are significant and validate the information about the degree of their synchronisation. In a case where $\hat{\rho}_S$ is insignificant, the high concordance between two series might be caused by a prolonged expansion phase in both series during the time period under examination. The empirical analysis is conducted on a pairwise basis across all eight markets together with the 8 Capital Cities National Index.

The concordance indicators using the modified Harding & Pagan (2006) methodology are reported in Table 2, whilst the corresponding Rho's, together with the relevant p-values, are displayed in Table 3. The results do reveal interesting findings which imply an element of tiers being present in the metropolitan markets of the Australian residential market. It can be seen that whilst Sydney and Melbourne are significantly synchronised in terms of the phase of their cycles, neither of the two largest Australian cities share significant coefficients with respect to many of the other markets. In the case of Sydney it is only significantly synchronised with Adelaide with a concordance indicator of 0.7083 and a reported rho of 0.2561 which is marginally significant, with a p-value of 0.06. For Melbourne a significant result is only reported with respect to Perth, with a Rho of 0.3667. In contrast, neither of the two largest centres are found to be significantly synchronised with any other market. This would indicate that the two largest metropolitan markets, behave in a manner distinct from the rest of the Australian market The findings reported are in many respects similar to the bilateral cointegration results of Luo et al. (2007). Whilst a large number of significant results were noted, there was a marked reduction in the number when Sydney and Melbourne were examined. Sydney was only found to be cointegrated with Melbourne, whilst Melbourne added Adelaide and Perth.

In contrast, with respect to the remaining centres there are a number of pairings that report significant findings, and in every case at least two such results are found. This is particularly so in the case of Adelaide which is significantly synchronised with four markets (Brisbane, Perth, Hobart, Canberra). Three significant pairings are noted for both Hobart (Adelaide, Darwin Canberra) and Canberra (Brisbane, Adelaide and Hobart). Finally, for Brisbane, significant rho's are found with respect to Adelaide and Canberra, Perth with both Adelaide and Darwin and Darwin with Perth and Hobart.

A few issues arise from the analysis. Firstly, it is noticeable that despite the distances involved when examining the Australian market, the importance of contiguous and non-contiguous markets is evident. There is a tendency for markets to be relatively close to each other to be more likely to report evidence of synchronised cycles. This can be illustrated when considering Perth. For example, in terms of geographically dispersed markets, significant results are not reported with Perth and Brisbane, Given the finding with Sydney and Melbourne, it is not that surprising that a significant result is also observed with respect to the two smallest centres, Hobart and Darwin. Whilst geographically dispersed, their housing markets are synchronised. In addition, Hobart significantly related to Canberra. Whilst a larger market than either Darwin or Hobart, Canberra is the smallest mainland city on the east and southern coasts. The majority of the significant findings are between the second tier of cities. One result that warrants further mention is the case of Perth and Darwin. Whilst a significant Rho is reported, it is negative in sign. The modified concordance indicator in this case is also the lowest observed (0.5758). These results indicate that the two markets are actually significantly counter cyclical. With respect to the Eight Capital Cities index it is not too surprising that Sydney and Melbourne report significant degrees of concordance given their relative size and weight in the aggregate index. Whilst Canberra is not significantly synchronised with either of its two large neighbours it is also so with the national index.

5: Decomposition of Housing Cycles

The final section of the paper considers the cyclical behaviour of the eight Australian Metropolitan markets in the context of the decomposition approaches of Beveridge-Nelson (1981) and Hodrick-Prescott (1997). Both of these approaches have been used extensively in the economic cycle's literature to decompose series into their trend and cyclical components. The rationale behind their application in a business cycle context can be easily transferred to a housing market one. By decomposing the series' we can isolate the cyclical element that can be defined as being the deviation from the long-term trend.

The Beveridge-Nelson decomposition separates a time-series (y_t) into permanent (trend) and transitory (cyclical) components as follows:

$$y_t = P_t + T_t \tag{7}$$

Assuming that y_t is an ARIMA (p,1,q) process we can re-write Equation (7) as below:

$$\Delta y_t = \Delta P_t + \Delta T_t \tag{8}$$

Given that the first difference of such a process has a stationary infinite order moving average representation, as displayed in Equation (9) below, we can therefore further define Δy_t as in Equation (10):

$$\Delta y_t = c_0 e_t + c_1 e_{t-1} + ... = C(L) e_t \tag{9}$$

$$\Delta y_t = C(1) e_t + \Psi(L)(1 - L) e_t \tag{10}$$

Where $\Psi(L) = \psi_0 + \psi_1 L + ...$ is a polynomial with $\lim_{j \to \infty} = 0$. The components can therefore be identified as follows:

$$\Delta P_t = C(1)e_t \tag{11}$$

$$\Delta T_t = \Psi(L)(1-L)e_t \tag{12}$$

As $\Delta T_t = (1-L)T_t$, then $T_t = \Psi(L)e_t$. This means that P_t is an I(1) process and T_t is I(0). The Beveridge-Nelson decomposition therefore has two primary characteristics. Firstly, that the shocks in the permanent component are white noise and secondly, that the shocks in the permanent and transitory components are perfectly correlated through the common value (e_t). To empirically decompose the series in question we therefore estimate the permanent component as follows (Newbold, 1990):

$$\bar{y}_t = y_t + \lim_{k \to \infty}\left[\sum_{j=1}^{k} \hat{w}_t(j)\right] = y_t + c_t \tag{13}$$

$where:$

$\bar{y}_t =$ The permanent component of y_t

$\hat{w}_t = \Delta \hat{y}_t - \mu$

μ is the mean of the ARIMA(p,q) process of the permanent component of Δy_t

$\Delta \hat{y}_t$ are the forecasts from a fitted ARMA process of Δy_t

The alternative decomposition model used is that of Hodrick & Prescott (1997). This decomposition is a linear filter that estimates a smoothed trend series. This is achieved by minimizing the variance of the original series (y) around the trend (T), subject to a constraint concerning the second difference of T. Therefore, T is selected such that it minimizes the following:

$$\sum_{t=1}^{T}(y_t - T_t)^2 + \lambda \sum_{t=2}^{T-1}\left[(T_{t+1} - T_t) - (T_t - T_{t-1})\right]^2 \tag{14}$$

The parameter λ controls for the smoothness of the series. For the purposes of this paper we use the frequency power rule of Ravn & Uhlig (2002). This is defined such that the number of periods per annum is divided by 4, squared and multiplied by 1,600. Given that we have quarterly data this provides a figure of 1,600 for our purposes.

The results from the two decompositions are displayed in Figures 2 and 3 and Table 4. Figure 2 displays the trends estimated from the two approaches, whilst the corresponding cyclical estimates are displayed in Figure 3. As would be expected the Hodrick-Prescott Filter provides smoother trends than the corresponding Beveridge-Nelson estimates, as can be clearly seen in Figure 2. This also means that a higher proportion of the variability of the series is captured in the cyclical element of the Hodrick-Prescott decomposition. Therefore, the cyclical elements may display greater variation, a feature that is also captured in the standard deviation figures reported in Table 4 in the case of four of the eight markets. The reason behind this difference is that the Beveridge-Nelson decomposition defines the trend as the random walk component. It would therefore be expected that it capture more variability in comparison to the Hodrick-Prescott approach. Table 4 reports the correlations between the cyclical elements for each of the eight markets, together with the standard deviation and the first order autocorrelation of the cyclical elements. The results illustrate a degree of divergence across the cities in terms of the correlations across the cyclical components. Indeed, the correlations are in many respects supportive of the results from the concordance indicators. As with the previous results the strong relationship between the two largest metropolitan areas, Sydney and Melbourne, is evident. In the case of the Beveridge-Nelson decomposition the cyclical elements for the two markets have a correlation of 0.5789, the highest coefficient reported for either city. The corresponding coefficient when using the Hodrick-Prescott framework is 0.8165, and again the highest noted for either Sydney or Melbourne. Indeed, with the exception of Canberra, the only case where either Sydney or Melbourne report a correlation above 0.50 is with Brisbane in the case of Sydney with the Hodrick-Prescott decomposition.

A key finding earlier in the concordance analysis was that strong relationships were observed amongst the smaller markets, especially those in the east of Australia, a result that is echoed in these tests. Correlations in excess of 0.50 are observed for the pairings of Adelaide-Brisbane, Brisbane-Hobart, Brisbane-Canberra, Adelaide-Hobart and Canberra-Hobart in the case of the Beveridge-Nelson results. Only the two most isolated centres, Perth and Darwin, see no correlation in excess of 0.50 with any other market using either decomposition technique. The results with respect to the correlations do not however reveal parallels in the standard deviations reported. There are also quite distinct differences in the volatility of the cyclical components in either framework. In the Beveridge-Nelson case the market with the highest volatility is Brisbane, whilst with the Hodrick-Prescott data, this is the case with Perth. Broadly speaking the cyclical component tends to be highest across the two methodologies, in Sydney, Hobart and the aforementioned Brisbane and Perth. These four cut across the three broad groupings of Sydney-Melbourne, the remaining eastern cities and the outlying Perth and Hobart. The differences observed in the volatilities are consistent with previous work on business cycles, such as Carlino & Sill (2001) in their analysis of regional income cycles in the US.

6: Concluding Comments

The analysis of interlinkages across metropolitan housing markets has largely considered the issue from the perspective of house price diffusion and convergence. This study has examined the commonalities present in the cyclical behaviour of eight metropolitan centres in Australia using approaches originated in the business cycle literature. Both the measure of concordance of cycles and the decomposition of the price series into their permanent and cyclical elements provide complementary evidence to the existing Australian empirical literature. Sydney and Melbourne, as the two largest markets display high degree of interaction and commonalities using either approach. However, in the vast majority of cases these commonalities are not extended to the remaining six markets. In contrast however, there is widespread evidence of synchronization using either empirical approach, with the remaining markets, and in particular those markets on the eastern and southern seaboards of Australia. The results are consistent with many of the existing work to have considered Australia, and given the different empirical framework adopted provide additional support to the notion that Sydney and Melbourne do have distinct cyclical features in comparison to the remaining metropolitan centres in Australia.

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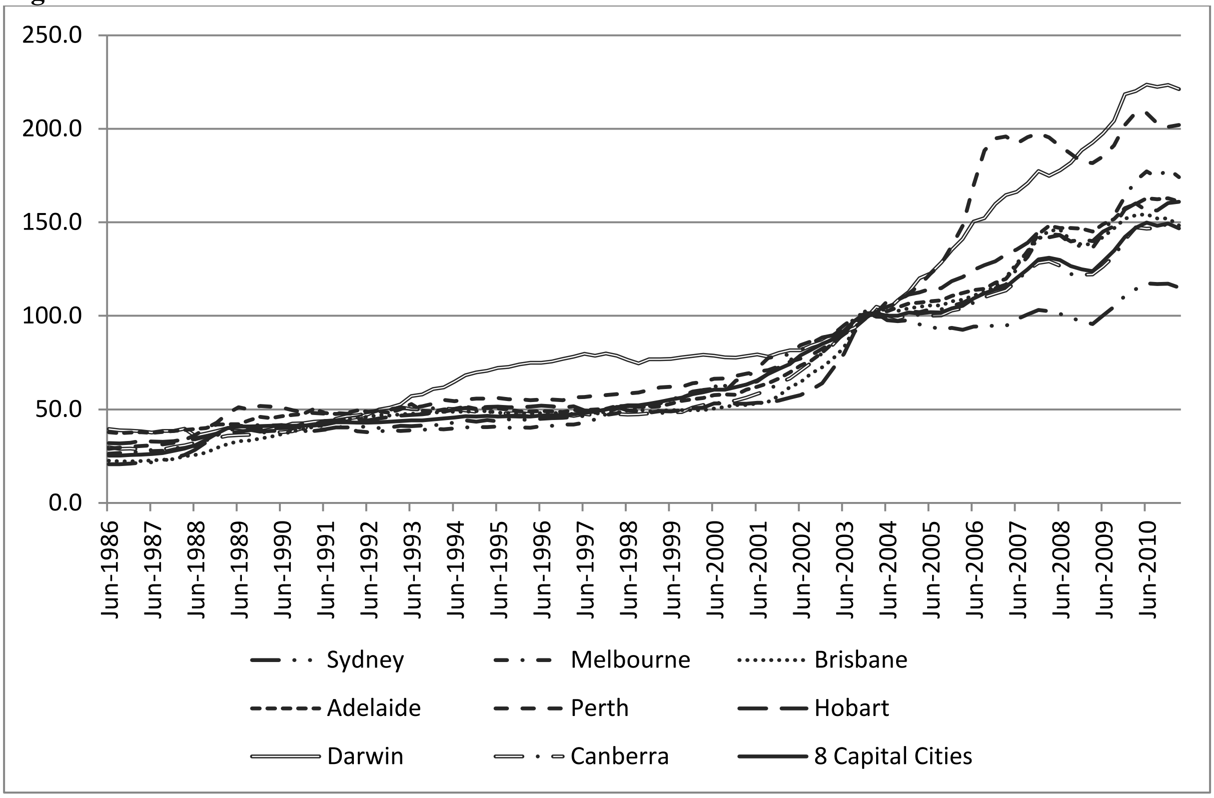
Tables & Figures

Figure 1: ABS House Price Indices



Notes: Figure 1 displays the raw index data for the eight capital cities used in the empirical tests. The indices are displayed in notional terms. The revised SBS weights and indices from 2002 are backdated with the original series.

Table 1: Summary Statistics

	Average Return	Standard Deviation	Augmented Dickey-Fuller Unit Root Tests	
			Levels	First Difference
Sydney	1.7357	2.8827	-1.8711	-4.2313***
Melbourne	1.9104	3.0254	0.7279	-4.5345***
Brisbane	1.8995	2.6287	-1.4258	-3.9443***
Adelaide	1.4637	2.3662	0.8776	-7.0731***
Perth	1.9618	2.9891	-0.6164	-3.9399***
Hobart	1.6353	2.7359	-0.2429	-3.4625**
Darwin	1.7422	2.5424	0.5383	-4.8206***
Canberra	1.6266	2.4625	-0.3992	-5.0733***
8 Capital Cities	1.7793	2.2114	-1.1178	-3.9206***

Notes: Table 1 details the summary statistics for the different markets examined. * indicates significance at the 10% level, ** at the 5% level and *** at the 1% level.

Table 2: Concordance Measures

	Sydney	Melbourne	Brisbane	Adelaide	Perth	Hobart	Darwin	Canberra
Melbourne	0.7561							
Brisbane	0.6479	0.6768						
Adelaide	0.7083	0.6934	0.7506					
Perth	0.6896	0.7738	0.6954	0.7197				
Hobart	0.6372	0.6465	0.6863	0.8157	0.7081			
Darwin	0.6476	0.6366	0.6562	0.6367	0.5758	0.7061		
Canberra	0.6878	0.6934	0.7888	0.7374	0.6754	0.7038	0.6360	
8 Cities	0.8643	0.8990	0.7172	0.7222	0.7611	0.6669	0.6570	0.7222

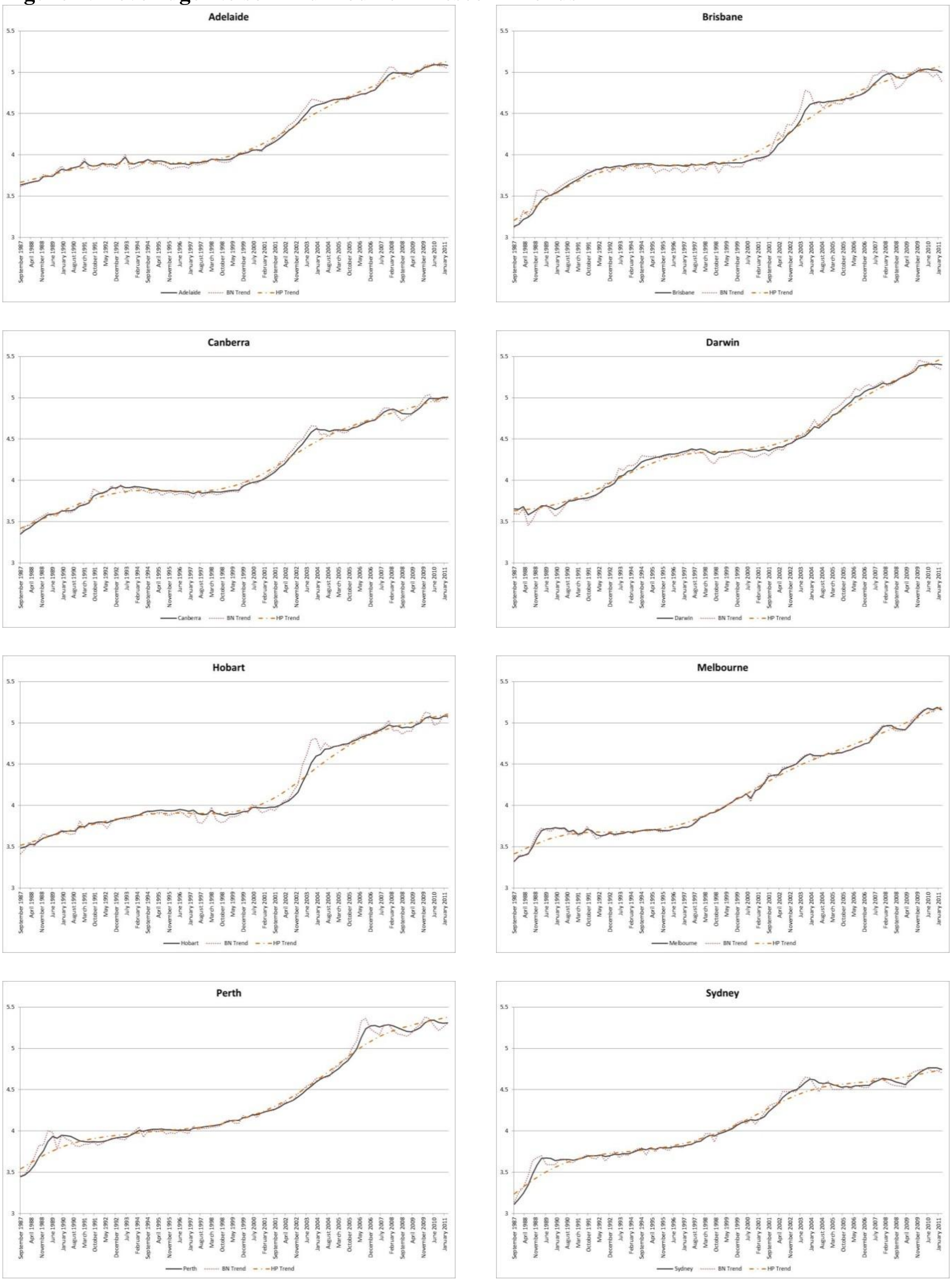
Notes: Table 2 reports the revised concordance indicator of Harding & Pagan (2006), as displayed in Equation (5).

Table 3: Rho's

	Sydney	Melbourne	Brisbane	Adelaide	Perth	Hobart	Darwin	Canberra
Melbourne	**0.3506**							
	(0.0214)							
Brisbane	0.0584	0.0649						
	(0.6237)	(0.5947)						
Adelaide	**0.2561**	0.1681	**0.3246**					
	(0.0620)	(0.2151)	**(0.0027)**					
Perth	0.1900	**0.3667**	0.1467	**0.2583**				
	(0.1602)	**(0.0006)**	(0.1956)	**(0.0349)**				
Hobart	0.0412	-0.0063	0.1070	**0.5074**	0.1939			
	(0.7042)	(0.9446)	(0.2131)	**(0.0000)**	(0.1019)			
Darwin	0.0800	-0.0183	0.0367	0.0383	**-0.1550**	**0.1883**		
	(0.4718)	(0.8604)	(0.7655)	(0.7305)	**(0.0799)**	**(0.0654)**		
Canberra	0.2039	0.1681	**0.4289**	**0.3219**	0.1407	**0.2065**	0.0364	
	(0.1122)	(0.1684)	**(0.0000)**	**(0.0047)**	(0.2491)	**(0.0272)**	(0.7445)	
8 Cities	**0.6429**	**0.7078**	0.1818	0.2468	**0.3312**	0.0519	0.0390	**0.2468**
	(0.0000)	**(0.0000)**	(0.1287)	(0.1226)	**(0.0141)**	(0.5929)	(0.7251)	**(0.0602)**

Notes: Table 3 reports the rho's estimated from Equation (6). P-values are reported in parenthesis. Those estimates that are of significance of at least 10% are displayed in bold.

Figure 2: Beveridge-Nelson and Hodrick-Prescott Trends



Notes: Figure 2 displays the original index data together with permanent trends estimated from the Beveridge-Nelson and Hodrick-Prescott decomposition techniques for each of the eight metropolitan markets.

Figure 2: Beveridge-Nelson and Hodrick-Prescott Cycles

Notes: Figure 3 displays the cyclical components for the eight markets as estimated using both the Beveridge-Nelson and Hodrick-Prescott techniques.

Table 4: Correlation Matrix and Standard Deviations for Cycluical Components

	Sydney	Melbourne	Brisbane	Adelaide	Perth	Hobart	Darwin	Canberra	StDev	ρ
Panel A: Beveridge-Nelson Cycles										
Sydney	1.0000								4.3292%	0.5335
Melbourne	0.5789								1.8332%	0.3491
Brisbane	0.3857	0.3368	1.0000						6.6897%	0.7035
Adelaide	0.2047	0.2137	0.6640	1.0000					3.5183%	0.7607
Perth	0.3812	0.3840	0.3147	0.1166	1.0000				5.2818%	0.5928
Hobart	0.1886	0.1979	0.6666	0.6454	0.3361	1.0000			6.4278%	0.7565
Darwin	-0.2698	-0.1033	0.0341	0.0841	0.1693	0.3076	1.0000		5.0182%	0.8477
Canberra	0.5569	0.5021	0.6990	0.4913	0.4111	0.5422	-0.0757	1.0000	3.6877%	0.6205
Panel B: Hodrick-Prescott Cycles										
Sydney	1.0000								5.7602%	0.9067
Melbourne	0.8165	1.0000							4.8484%	0.8365
Brisbane	0.5185	0.3867	1.0000						5.0775%	0.9079
Adelaide	0.4702	0.4125	0.8503	1.0000					3.4841%	0.8221
Perth	0.3547	0.4179	0.1000	-0.0385	1.0000				6.2270%	0.9180
Hobart	0.2715	0.0423	0.7006	0.5639	0.1839	1.0000			5.2038%	0.9055
Darwin	-0.2702	-0.3098	0.0426	-0.0400	0.3284	0.3696	1.0000		3.7771%	0.8320
Canberra	0.5649	0.3747	0.7877	0.6954	0.0218	0.5139	-0.0467	1.0000	4.7168%	0.9001

Notes: This table reports summary data based upon the cyclical series' estimated for each of the eight metropolitan housing markets. Correlations are estimated for each pairing of the cyclical components. The final two columns report the standard deviation of the cyclical components and the first order autocorrelation of each series (ρ) respectively.

Endnotes:

[1] Research has also considered inter-market dynamics and house price diffusion in Canada (Allen et al., 2009), Finland (Oikarinen, (2006), Ireland (Stevenson, 2004) and Taiwan (Chien, 2010). In a Japanese context Sanjuan et al. (2009) find evidence of cointegration between rents and farmland prices in nine Japanese regions.

[2] Munro & Tu (1996) report results largely supportive of the ripple effect. However, the results also indicate that non-English regions appear to be relatively independent to fluctuations, with far weaker evidence of a ripple effect into Scotland, Wales and Northern Ireland.

[3] Papers such as Ashworth & Parker (1997) also undertake tests for cointegration, whilst Drake (1995) uses a Kalman Filter framework to consider similar issues. Holly & Jones (1997) take a long-term perspective, from 1939, to consider whether UK house prices are cointegrated with key drivers such as income and population.

[4] Cook (2006) uses an alternative test of asymmetry, namely threshold autoregressive methods. However, similar results are reported.

[5] Bilgin et al. (2010) use the same panel approach in the context of rental values in three Turkish cities. In this case however no evidence of convergence is noted.

[6] Clapp & Tirtiroglu (1994) found evidence of significant price diffusion between submarkets in Hartford Connecticut, but not however, between markets that were not contiguous. Pollakowski & Ray (1997) consider both a broad analysis of US regions and a specific analysis of the Greater New York metropolitan area. The results reported note that the national results are weaker in terms of spatial diffusion, with no consistent evidence that neighbouring or contiguous regions, as defined by census divisions, are more significant than non-contiguous regions. However, there is broad evidence that diffusion does take place, with price movements in regions significantly affecting subsequent price changes in other areas. The analysis of New York does however support the positive feedback hypothesis and the principle of spatial diffusion. A higher number of significant findings are reported for neighbouring submarkets of the Greater New York region.

[7] A number of recent US papers has considered regional elements, in a number of cases looking at the role of economic shocks on regional house price dynamics (e.g. Fratantoni & Schuh 2003; Del Negro & Otrok, 2007; Clark & Coggin, 2009; Fadiga & Wang, 2009; Holly et al., 2010; Kuethe & Pede, 2011; Riddel, 2011).

[8] An early paper to have consider such issues was Bourassa & Hendershott (1995) who examined the six largest Australian metropolitan markets.

[9] Other studies to have considered aspects of the Australian market include: Yates (2002), Dvornak & Kohler (2007), Ma & Liu (2010) and Lee & Reed (2011).