Sam Houston State University
Department of Economics and International Business
Working Paper Series

City Relative Price Dynamics in Australia: Are Structural Breaks Important?

Hiranya K. Nath
Sam Houston State University

Jayanta Sarkar
Queensland University of Technology

SHSU Economics & Intl. Business Working Paper No. 13-01
February 2013

Abstract:
This paper examines the dynamic behavior of relative prices across seven Australian cities by applying panel unit root test procedures with structural breaks to quarterly CPI data for 1972Q1-2011Q4. We find overwhelming evidence of convergence in city relative prices. Three common structural breaks are endogenously determined at 1985, 1995, and 2007. Further, correcting for two potential biases, namely Nickell bias and time aggregation bias, we obtain half-life estimates of 2.3-3.8 quarters that are much shorter than those reported by previous research. Thus, we conclude that both structural breaks and bias corrections are important to obtain shorter half-life estimates.

City Relative Price Dynamics in Australia: Are Structural Breaks Important?

Hiranya K. Nath*

Jayanta Sarkar†

Abstract: This paper examines the dynamic behavior of relative prices across seven Australian cities by applying panel unit root test procedures with structural breaks to quarterly CPI data for 1972Q1-2011Q4. We find overwhelming evidence of convergence in city relative prices. Three common structural breaks are endogenously determined at 1985, 1995, and 2007. Further, correcting for two potential biases, namely Nickell bias and time aggregation bias, we obtain half-life estimates of 2.3-3.8 quarters that are much shorter than those reported by previous research. Thus, we conclude that both structural breaks and bias corrections are important to obtain shorter half-life estimates.

JEL Codes: C33; E31; R19

Keywords: Relative price convergence; Structural break; Panel unit root test; Half-life; Time aggregation bias

* Department of Economics and International Business, Sam Houston State University, Huntsville, TX 77341-2118, USA; Phone: +1-936-294-4760; E-mail: eco_hkn@shsu.edu

† School of Economics and Finance, QUT Business School, Queensland University of Technology, Brisbane, QLD, Australia; Phone: +61-7-3138-4252; E-mail: jayanta.sarkar@qut.edu.au

City Relative Price Dynamics in Australia: Are Structural Breaks Important?

1. **Introduction**

There is a growing literature on relative price convergence across different cities or regions within a country.[1] This literature was originally motivated by a desire to develop a better understanding of the deviations from the international purchasing power parity (PPP). The PPP hypothesis, based on the *law of one price*, implies that, "once converted to a common currency, national price levels should be equal" (Rogoff, 1996). The test of this hypothesis essentially involves finding evidence of mean-reversion in real exchange rate. Numerous empirical studies using a variety of datasets and empirical methods emerged over last several decades. The results have been mixed: while some studies find evidence of mean reversion, others do not. Even in some cases of mean reversion, the speed of convergence has been found to be extremely slow. The slow mean reversion is known as the "PPP puzzle" (*a la* Rogoff, 1996). The literature has explored various factors that help explain the rejection of the PPP hypothesis as well as the puzzle. Some of these factors, such as tariff and non-tariff barriers, fluctuations in nominal exchange rates, heterogeneity of consumption baskets that are used to construct consumer price indices in different countries, are not relevant while considering price indices for different cities or regions within the same country. Thus, intra-national price data provide an opportunity to conduct a natural experiment to understand the deviations from PPP.

The study of within-country relative price movements is also important because of their implications for resource allocations and regional growth. For example, inflation differentials

[1] The existing literature primarily focuses on convergence of both aggregate as well as individual prices across different cities in the U.S. For example, Culver and Papell (1999), Cecchetti *et al.* (2002), Chen and Devereux (2003), Nath and Sarkar (2009), Sonora (2009), Chmelarova and Nath (2010), Basher and Carrion-i-Silvestre (2009, 2011), Huang *et al.* (2011), Huang *et al.* (2012), Hegwood and Nath (2012) examine price index convergence across major U.S. cities. While Chen and Devereux (2003) and Basher and Carrion-i-Silvestre (2009) consider absolute price convergence, others investigate relative price convergence. All these studies use aggregate price indices. However, they were preceded by two influential studies – Engel and Rogers (1996) and Parsley and Wei (1996) – that look into the disaggregate prices of various commodities across the U.S. cities. Recently, Crucini and Shintani (2008) have used micro-level price data to investigate persistence of PPP deviations across 13 U.S. cities. The studies that use city level CPI data from other countries include Carrion-i-Silvestre *et al.* (2004) for Spain, Chaudhuri and Sheen (2004) for Australia, Sonora (2005) for Mexico, and Busetti *et al.* (2006) for Italy.

across cities – implied by relative price movements – determine the differences in real wages and real interest rates, which in turn influence the directions of labour and capital flows, respectively. In particular, persistent differences (in case of PPP violation) may lead to severe misallocations of resources with consequent suboptimal regional patterns of growth and overall inefficiency.

In this paper, we examine the dynamics of relative prices across seven capital cities in Australia using quarterly data on CPI from 1972Q1 to 2011Q4. In particular, we investigate the importance of structural changes in relative price convergence. To the best of our knowledge, the only study on the relative prices across Australian cities is by Chaudhuri and Sheen (2004) (hereafter, CS). They apply both univariate and panel data techniques to aggregate as well as disaggregate CPI data for seven cities from 1972Q3 to 1999Q1. The current study uses an extended dataset for the same seven cities. Also, it focuses on the dynamics of aggregate CPI and uses panel data methods only. Most importantly, we incorporate structural breaks that are determined endogenously, and examine their implications for relative price behaviour.[2] While CS recognize the potential significance of structural changes, their investigation relies on analyses of different sample periods selected on the basis of exogenous policy changes. Like most studies in this literature, we estimate half-life – a measure of the speed of convergence. In addition to correcting for small-sample bias, as in CS, we correct for an additional bias that arises due to time aggregation of price data in the panel estimation of half-life.

It is important to recognize the significance of including structural breaks in the unit root test in the context of the PPP theory. In the literature, a rejection of the unit root null is typically interpreted as evidence in support of PPP. Thus, it implies that after a temporary deviation, the real exchange rate returns to a constant mean. However, when structural breaks are included, a rejection of the unit root null indicates mean reversion to shifting means. In that case, the long-

[2] Following the seminal contribution of Perron (1989), inclusion of mean shifts (representing structural breaks) has been a way of finding support for the PPP hypothesis and resolving the PPP puzzle. Some prominent examples are: Dornbusch and Vogelsang (1991), Perron and Vogelsang (1992), Culver and Papell (1995), and Hegwood and Papell (1998, 2002) with only one structural break; and Lumsdaine and Papell (1997) and Papell and Prodan (2006) with two breaks. The studies that use structural breaks in the panel context include Papell (2002), Im *et al.* (2005), Breitung and Candelon (2005), Narayan (2008), and Lin and Lee (2010).

run PPP may not always hold.[3] The reversion to mean that is subject to occasional structural changes is called Quasi Purchasing Power Parity (QPPP) by Hegwood and Papell (1998) and Qualified Purchasing Power Parity (QPPP) by Papell and Prodan (2006).

Our paper is closely related to Sonora (2009), Basher and Carrion-i-Silvestre (2009, 2011), Huang *et al.* (2011) and Hegwood and Nath (2012). Applying unit root test procedures that incorporate structural breaks to annual city CPIs across major cities in the U.S., these studies find overwhelming evidence of relative price convergence. Sonora (2009), Basher and Carrion-i-Silvestre (2011), and Hegwood and Nath (2012) also find that inclusion of structural breaks reduces the half-life estimates.

We apply panel unit root test procedures with and without structural breaks to quarterly CPI data for seven Australian cities between 1972Q1 and 2011Q4. We find overwhelming evidence of convergence in city relative prices. Three common structural breaks are endogenously identified in 1985, 1995, and 2007. Further, correcting for two potential biases, namely Nickell bias and time aggregation bias, we obtain half-life estimates of 2.3 - 3.8 quarters that are much shorter than those reported by CS. Thus, we conclude that both structural breaks and bias corrections are important to obtain shorter half-life estimates. These results are robust to shorter sample periods as well as lower frequency data (annual in our case) with longer sample period.

The rest of the paper is organised as follows. Section II discusses the data and methodology used in the analysis. In Section III, we present various panel unit test results and half-life estimates for relative prices across Australian cities. The results are reported for cases with and without structural breaks. Section IV discusses of the economic significance of the structural breaks identified in the models of relative prices. In the final section, we summarise and conclude.

[3] As Papell and Prodan (2006) discuss, if there is a one-time shift in the mean and the series reverts to shifting means, the long-run PPP does not hold in the sense that it does not return to the long-run mean. With multiple breaks, say two breaks, the long-run PPP holds only if the shifts are offsetting and the series returns to a constant mean.

2. Data and Methodology

2.1 Data

We obtain quarterly CPI data for seven Australian cities for the period: 1972Q1 - 2011Q4, from the Australian Bureau of Statistics (ABS). The cities in our sample are the capital cities of the Australian states and the Australian Capital Territory: Adelaide, Brisbane, Hobart, Melbourne, Perth, Sydney, and Canberra.[4] Although quarterly data are available since September 1948, a single CPI value was duplicated for several consecutive quarters for each city prior to 1972. This seems to indicate that data were collected at infrequent and irregular intervals before 1972, and therefore we exclude them from our sample of quarterly data. We use the following equation to construct the relative price series for each city:

$$r_{i,t} = 100 \times \left(\ln P_{i,t} - \ln P_t \right) \tag{1}$$

where $r_{i,t}$ is the logarithm of relative price and $P_{i,t}$ is the CPI in city i in year t, and P_t is the simple average of city CPIs.[5] Note that this relative price represents the percentage deviation of CPI in a city from the national average CPI.[6] The city average CPI captures the common time effect component of relative prices.

2.2 Methodology

We use panel unit root test procedures to examine mean-reversion in relative prices. To evaluate the importance of structural changes, we first conduct these tests with no structural breaks.[7]

Panel Unit Root Test with No Structural Breaks

The test is in the Augmented Dickey-Fuller (ADF) framework and involves estimating the following regression:

---

[4] We exclude Darwin for which CPI data are available only since 1980.

[5] In Australia, national CPI is calculated as the weighted average of the city CPIs. If we use the weighted instead of simple city average, the results do not change qualitatively.

[6] In the international PPP literature, this relative price would be equivalent to real exchange rate. Although a numeraire currency is chosen for calculating real exchange rate, we use the average city CPI, an approach previously adopted by Cecchetti *et al.* (2002), Chen and Devereux (2003), Nath and Sarkar (2009), and Basher and Carrion-i-Silvestre (2011).

[7] Our methodology and empirical strategy are similar to those in Hegwood and Nath (2012).

$$\Delta r_{i,t} = \mu_i + \rho_i \, r_{i,t-1} + \sum_{j=1}^{k_i} c_{i,j} \Delta r_{i,t-j} + \varepsilon_{i,t} \tag{2}$$

The subscript i = 1, …, N indexes the cities in the panel and t indexes time period. In this specification, the intercepts, μ_i, and lag lengths, k_i, are allowed to vary across cities to capture the city-specific idiosyncracies. We use feasible generalised least squares (FGLS) seemingly unrelated regression (SUR) to estimate (2). This method accounts for contemporaneous and serial correlation, both of which are likely to be present in city relative prices.[8] The number of lagged differences, k_i, is determined using the general-to-specific method suggested by Campbell and Perron (1991) and Ng and Perron (1995). This method involves setting a maximum lag length, k_i^{max}, and paring it down to the number of lags where the lagged difference is significant at the 10% level. We start with a maximum lag of 8 quarters for each city.

The null hypothesis is that each series contains a unit root, H_0: $\rho_i = 0$ for all i. The alternative hypothesis is H_1: $\rho_i = \rho < 0$, that is, all of the series are stationary. This alternative hypothesis requires a homogenous ρ, as in Levin *et al.* (2002).[9] Note that the distribution of the panel unit root test statistic is not standard. Therefore, we use Monte Carlo methods involving 5000 replications to calculate critical values that reflect the structure of our panel, accounting for both serial and contemporaneous correlations.[10]

Panel Unit Root Test with Structural Breaks

We now discuss a panel unit root test procedure that uses an Additive Outlier (AO) model framework where structural changes take place instantaneously. This model has been adapted for non-trending data incorporating one or more shifts of the intercept.[11] This panel unit root test

[8] See Murray and Papell (2000).

[9] A less restrictive alternative hypothesis that at least one of the series is stationary, which allows ρ to be heterogeneous, as in Im *et al.* (2003), would not be any more informative if we do reject the null, which is the case in this paper.

[10] For the details on this method, see Hegwood and Papell (2007).

[11] This is a panel adaptation of the univariate test in Perron and Vogelsang (1992). They include an additional set of 'crash' dummies. For a discussion on the panel adaptation, see Murray and Papell (2000).

comprises two stages. In the first stage, we estimate the following equation using the panel of city relative prices:

$$r_{i,t} = \mu_i + \sum_{l=1}^{m} \delta_{l,i} DI_{l,t} + u_{i,t} \tag{3}$$

where $DI_{l,t,}$ is the intercept break dummy and $DI_{l,t,} = 1$ for all $t > SB_l$ (the structural break date) and 0 otherwise; and m is the number of breaks in (3).[12,13] The second stage involves regressing the first difference of the residuals, $u_{i,t}$'s on their lagged value and lagged differences as follows:

$$\Delta u_{i,t} = \rho u_{i,t-1} + \sum_{j=1}^{k_i} c_{i,j} \Delta u_{i,t-j} + \varepsilon_{i,t} \tag{4}$$

As in the panel unit root test with no structural breaks, the number of lagged differences, k_i, is determined by the general-to-specific method. (3) and (4) are estimated sequentially for each possible break date, $SB_l = k^{max}+(8\times m)$, … T-{$8\times(m^{max}-m)$}-1, where k^{max} is the maximum lag length, m^{max} is the maximum number of breaks, and T is the number of observations. The period that minimizes the t-statistic on ρ is chosen to be the break date. The null hypothesis of a unit root is rejected if the absolute value of the (minimum) t-statistic on ρ is greater than the appropriate critical value. As before, the critical values are calculated using Monte Carlo methods.

We then use two simple information-based criteria to determine the appropriate number of breaks.[14] These criteria are: Bayesian Information Criterion (BIC) as suggested by Yao (1988) and a modified Schwarz' criterion (LWZ) as suggested by Liu *et al.* (1997). According to Bai and

---

[12] To be consistent with the PPP theory, we do not include a time trend and only allow for breaks in the intercept. However, Obstfeld (1993) develops a model in which real exchange rates contain a deterministic trend. His model draws on ideas from Balassa (1964) and Samuelson (1964) and argues that productivity differentials between countries determine the domestic relative prices of nontradables, leading to trend deviations from PPP in the long-run.

[13] We impose a restriction by forcing the break date(s) to be the same for all cities. Thus, we are focusing on structural breaks that are common to all cities in the sample. Allowing different breaks for each city also reduces the degrees of freedom and the power of the test. Furthermore, since common breaks usually reflect the effects of national and international events, it is relatively easier to discuss their economic significance.

[14] There are several aspects we need to consider in choosing the number of breaks. In case of multiple shifts, if two break dates are sufficiently close, they may reflect a temporary shock rather than structural breaks in the true sense of the term. Most importantly, we should be able to explain the economic significance of the break dates that are identified endogenously by this procedure.

Perron (2003), these two criteria perform reasonably well when there is no serial correlation in the errors. The specification of (4) does not allow serial correlation in the error term and it seems to be appropriate to use these information-based criteria.[15]

Unbiased Half-Life Estimates with Structural Breaks

As mentioned earlier, half-life is commonly used as a measure of the speed of convergence. It is the time required for any deviation from PPP to dissipate by one half. In an AR(1) case, half-life is calculated as follows:

$$h(\rho) = \frac{-\ln(2)}{\ln(\rho)} \tag{5}$$

where *h*( ) is the half-life and ρ is the AR coefficient.[16] As Choi *et al.* (2006) discuss, there are three potential sources of bias in panel data estimation of the half-life. *First*, if the autoregressive coefficients are significantly different across cities (that is, there is sufficient heterogeneity in the dynamic behavior of relative prices across cities), then panel estimation of a common autoregressive coefficient will be biased upward and so will be the implied half-life. *Second*, if the sample is small, the inclusion of a constant in the estimation of a dynamic regression introduces a downward bias and this is true in the panel context as well.[17] *Finally*, the time-averaging (also referred to as *time aggregation*) of some price data introduces a moving average structure into the regression error.[18] Since this is often ignored in the panel estimation of the autoregressive models of city relative prices, it introduces an additional upward bias in the AR coefficient.

Under the alternative hypothesis of the panel unit root test procedure discussed above, the AR coefficients (ρ_i s) are homogeneous. Thus, if the null hypothesis is rejected (which is the case

---

[15] There are several criteria and/or formal tests suggested in the literature to determine the number of structural breaks. These test procedures have their own strengths and weaknesses. Prodan (2008) discusses the potential pitfalls of the widely-used multiple structural change tests suggested by Bai and Perron (1998, 2003, 2006).

[16] For higher-order autoregressive models, (5) would approximate the true half-life. For such models, the exact half-life can be computed by impulse response analysis. The problem with general ARMA models is that the impulse response may be non-monotonic, resulting in multiple half lives.

[17] Nickell (1981) first discusses this small-sample bias in the panel context and therefore it has been known as the 'Nickell bias'.

[18] For a discussion on the time aggregation bias, see Imbs *et al.* (2005).

for our dataset), then the upward bias due to heterogeneity in dynamic behaviour of prices is no longer a matter of concern. Thus, the panel estimates of autoregressive coefficient and half-life for city relative prices would involve only two biases: a downward bias due to small sample size and an upward bias due to the moving average error term introduced by time aggregation of data. We use a fixed effects panel generalised least squares (GLS) estimation technique that incorporates structural breaks and also controls for cross-sectional dependence.[19] To sketch an outline of the procedure, suppose relative price in city i follows an AR(1) process:

$$r_{i,t} = \mu_i + \sum_{l=1}^{m} \partial_{l,i} DI_{l,t} + \rho_i r_{i,t-1} + u_{i,t} \tag{6}$$

where μ_i is a city-specific constant; $i = 1, 2, \ldots, N$; and $t = 1, 2, \ldots, T$. $DI_{l,t}$ is a dummy variable for the structural break, where $DI_{l,t} = 1$ if $t > SB_l$ and 0 otherwise. In the presence of time aggregation, the regression error has a moving average (MA) structure. Suppose $u_{i,t}$ follows an MA(1) process:

$$u_{i,t} = \upsilon_{i,t} + \lambda \upsilon_{i,t-1} \text{ and } \upsilon_{i,t} = \gamma_i \theta_t + \zeta_{i,t} \tag{7}$$

where γ_is are factor loadings, θ_t is the common shock, and $\zeta_{i,t}$s are serially and mutually independent. The factor loadings and the error covariance matrix are estimated by iterative method of moments, and then the estimated covariance matrix is used to obtain the feasible GLS estimate of ρ. Note that this estimated covariance matrix includes both the contemporaneous and the long-run covariance. We then adjust the estimated autoregressive coefficient for the Nickell bias, the time aggregation bias, and the combined Nickell and time aggregation bias as discussed in Choi *et al.* (2006) and use these bias-corrected estimates of autoregressive coefficient in (5) to obtain various unbiased estimates of the half-life to price index convergence among Australian cities.[20]

---

[19] This technique has been adapted from Phillips and Sul (2004) to include structural breaks.

[20] Time aggregation of the data introduces an interaction between the Nickell bias and the time aggregation bias, which requires additional adjustment in the estimation of the autoregressive coefficient. For a discussion, see Choi *et al.* (2006). The combined Nickell and time aggregation bias correction incorporates this adjustment.

3. Empirical Results

3.1 Panel Unit Test Results

Table 1 first presents the panel unit test results for the cases with and without structural breaks for the full sample: 1972Q1 – 2011Q4. Note that we conduct the panel unit root tests with 1, 2, 3, and 4 structural breaks and calculate respective values of BIC and LWZ to determine the optimal number of breaks. Both these criteria indicate that the model with three breaks is the best and therefore we report the results for this case only.[21, 22] The breaks are identified at 1985Q1, 1995Q2, and 2007Q2. The null hypothesis of a unit root is rejected at the 5% level for both cases.[23] These results are consistent with the panel results reported in other studies, particularly with those reported for Australian cities by CS. However, for appropriate comparison with their study, we also report the results for the sample period: 1972Q3 – 1999Q1, in Table 1. Again, the unit root null is overwhelmingly rejected for the cases with no breaks and two breaks. If we include more than two breaks, some of the identified break dates are too close and therefore we decide to go with a maximum of two breaks.[24] Thus, these panel unit root test procedures - with and without structural breaks - find overwhelming evidence in support of convergence in Australian city relative prices.

[Insert Table 1]

3.2 Unbiased Half-life Estimates

In Table 2, we report the autoregressive coefficients and implied half-lives with no bias corrections (col. 1 – 2) and combined Nickell and time aggregation bias corrections (col. 3 – 8). We first present the estimates for the full sample with and without structural breaks. In Australia, price data on some items (such as milk, bread, fresh meat and seafood, fresh fruit and vegetables,

[21] That is, it has the lowest value of BIC and LWZ. However, we do not report the BIC and LWZ values for different models to conserve space. Interested reader may obtain the results from the authors.
[22] We choose the maximum possible number of breaks (4 in our case) in an *ad hoc* manner. Given the length of our sample period, this choice seems reasonable.
[23] For the case with three structural breaks, computing the critical values using the Monte Carlo method takes enormous amount of time and therefore we conduct only 500 replications.
[24] With three breaks, the dates are identified at 1988Q4, 1991Q3, and 1996Q3. Between one break and two breaks, two breaks have smaller BIC and LWZ.

alcohol, tobacco, women's outerwear, project homes, motor vehicles, petrol and holiday travel and accommodation) are collected monthly while for others they are collected quarterly. There are a few items such as education, for which price data are collected at much longer intervals. Therefore, we present three alternative scenarios under which time aggregation bias is corrected: (1) data collection frequency of 3 assuming that data are collected monthly; (2) data collection frequency of 1 assuming that data are collected quarterly; and (3) an intermediate frequency of 2.[25]

[Insert Table 2]

As the results show, with no bias correction and no structural breaks, the estimated half-life is 8.2 quarters. With structural breaks alone, the estimated half-life decreases by about 55% to 3.7 quarters. Corrected for Nickell and time aggregation biases, it further decreases to 2.3 to 2.5 quarters (about 72 – 70% lower). To compare our half-life estimates with those reported by CS, we also present the autoregressive coefficients and implied half-life estimates for the sample period 1972Q3 – 1999Q1 in Table 2. They report bias-corrected autoregressive coefficients in the range of 0.89 - 0.90 and implied approximate half-lives of 5.8 - 6.7 quarters.[26] Our results with Nickell bias correction for the case with no structural breaks reported in col. 7 – 8 (0.893, 6.1) are consistent with those results. However, when we include structural breaks and correct for both Nickell and time aggregation bias, the estimated half-lives are in the range of 2.2 - 2.4 quarters.

3.3 Panel Unit Root Test and Hal-life Results for Annual Data

Although most studies applying panel test procedures to annual data on city CPI (e.g. Cecchetti *et al.* 2002) reject the null hypothesis of a unit root, they often report a somewhat puzzling result of extremely slow speed of convergence in relative prices. However, for Australia, CS cannot

[25] Note that since we are using quarterly data, there will be no correction for time aggregation with frequency 1. That is, the autoregressive coefficients and half-life estimates reported in col. 7 - 8 are essentially corrected for Nickell bias only.

[26] See row 1 of Table 4 in CS.

reject the null hypothesis of a unit root while using annual data from 1972 to 1998. We, therefore, examine if we can find evidence of mean reversion in city relative prices using annual data for a longer period of time. Furthermore, we investigate if inclusion of structural breaks and correction of various biases help understand these puzzling results. Taking averages of quarterly data, we construct annual CPI data for the period from 1949 to 2011. We then conduct the panel unit root tests with no structural breaks and up to 4 structural breaks as before. Table 3 reports the panel unit test results along with the autoregressive coefficients and implied half-lives. We include the results for three structural breaks in 1964, 1984, and 2005.[27]

[Insert Table 3]

The first two rows indicate that the null hypothesis of a unit root is rejected at the 5% level for both cases: with no and three structural breaks. However, the half-life estimates of 9 years for the no-break case (without and with bias corrections) are similar to those reported for the U.S. cities by Cecchetti *et al.* (2002). The fact that bias corrected autoregressive coefficients (and, therefore, half-lives) are larger suggests that the Nickell bias outweighs the time aggregation bias. When we include structural breaks, the half-life estimates with bias corrections range between 2.7 - 3.2 years. When we consider annual data for the period between 1972 and 2011, the sample period for which we examine relative price dynamics using quarterly data, the estimated half-lives are much shorter than for the longer sample, indicating faster convergence. However, with no structural breaks, they are still much longer compared to those obtained with quarterly data. This result accords well with the general conclusion of CS. Interestingly, when we include three structural breaks and correct for Nickell and time aggregation biases, we obtain half-life estimates in the range of 0.7 - 0.8 year. They compare reasonably well with the corresponding

[27] BIC and LWZ values provide conflicting choices as for the optimal number of structural breaks. While BIC indicates that the model with 1 break is the best, LWZ indicates that the model with 4 breaks is the best. However, in case of the model with 4 breaks, the first two break dates are identified only 4 years apart. Therefore, we decide not to go with the 4 breaks model. Among others, the model with 3 breaks has the lowest LWZ value and visually fits the data well. Moreover, the second and the third breaks almost coincide with the breaks identified in our model with quarterly data.

bias-corrected half-life estimates of 3.8 quarters (0.95 year) and 2.3 quarters (0.6 year) using quarterly data.

We further examine annual data for the sample period: 1972 – 1998, for which CS are unable to reject the null of a unit root for city CPI. Nevertheless, they report half-life estimates that range from 2 to 3.9 years. In contrast, we reject the unit root null for relative prices using panel unit root test with no structural breaks. The half-life estimates are 2.6 to 2.7 years. However, if we correct for Nickell bias only (akin to CS), the half-life estimate is 4.7 years. With two structural breaks, the unit root null is rejected and the half-life estimates decrease to 0.8 – 1.3 years. These half-life estimates are similar to those with quarterly data for the sample period: 1972Q1 – 2011Q4.

Overall, panel unit root tests using annual data overwhelmingly reject the null hypothesis of a unit root in relative prices irrespective of whether we include structural breaks or not. Furthermore, the half-life estimates, even after correcting for Nickell bias and time aggregation bias, are larger with annual data than quarterly data when no structural breaks are included. This conclusion about half-life accords well with the findings by CS. However, when we include structural breaks, the half-life estimates decrease substantially and they are quantitatively similar for quarterly and annual data.

4. Potential Explanations of the Structural Breaks

In the current study, the structural breaks are determined endogenously by data. These are common breaks with heterogeneous mean shifts across seven cities. In this section, we try to understand the common (national level) events that took place around the break dates identified above and had differential effects on the price level in different cities. Note that the structural break dates identified above are not independent of the sample period or of the data frequency. However, for the same sample period, the breaks identified with quarterly and annual data are close to each other. For example, with quarterly data, the break dates are identified at 1985, 1995, and 2007 while, with annual data, they are identified at 1987, 1995, and 2006. In this section, we will focus on the first set of breaks identified for quarterly data from 1972 to 2011.

[Insert Figure 1]

In order to help understand the potential mechanisms through which common events may have caused these mean shifts, we first plot the relative price series along with the shifting means in Figure 1. We make the following observations. First, in 1985, the mean shifted upward for Adelaide, Canberra, and Melbourne and downward for the rest. Adelaide, Brisbane, and Sydney experienced upward mean shifts in 1995. Further, there were upward shifts in Adelaide, Brisbane, Canberra, and Perth in 2007. Second, the mean shifted upward for Adelaide and downward for Hobart at all break dates. For these two cities, the general price level has been continuously moving away from the national average: in positive direction for Adelaide and in negative direction for Hobart. That is, the cost of living is continuously rising in Adelaide while it is continuously falling in Hobart relative to the national average. In other cities, they are meandering around the national average. Third, the year-to-year variations in relative prices have gone down in all cities during almost all intervals between subsequent breaks.[28] We now discuss major national events that occurred around the structural breaks with potential effects on city relative prices for each break date separately.

4.1 First Structural Break in 1985

The first break in 1985 coincided with the aftermath of the recession of 1982-83, that witnessed large fiscal expansion and several important policy reforms. The strong growth during 1984-85 was fuelled by investment in private dwelling, private non-dwelling construction, and increased capital expenditure in public enterprises. However, there was no consensus on the appropriate monetary policy. As the central bank abandoned monetary (M3) targeting, both nominal and real interest rate rose sharply in 1985 and the interest rate volatility increased during the late 1980s.

[28] The standard deviation of relative prices increased slightly between 1985 and 1995 only in Hobart and Melbourne and, after 2007, in Melbourne.

This was particularly harmful for interest rate sensitive industries and, depending on the spatial distribution of these industries, would have differential effects on growth and inflation across regions.

The most important development around this structural break date with potentially differential effects across cities was perhaps the trade liberalisation measures undertaken by Australia. Australia 'opted out' of the post-war rounds of trade liberalization under GATT as its proposal to put full employment before freeing trade as the aim of the prospective International Trade Organization was not accepted by the major trading nations (Capling, 2000). Most Australian industries were protected behind relatively high trade barriers and Australia gained little from free trade unlike other OECD countries. Since around 1985, however, these trends have been reversed. A sharp depreciation of the Australian dollar after its floating in December 1983 and a series of policy measures to improve international competitiveness of the economy spurred the exports of manufacturing goods and services.[29] The states of Victoria (VIC) and South Australia (SA) with large concentration of manufacturing and service enterprises benefitted the most. The upward shift of the mean relative prices after 1985 for Adelaide and Melbourne may have been a reflection of this development. In contrast, Tasmania, despite having a fair share of manufacturing activities, missed out largely due to relatively low labour utilisation rate (about 15 percent more than the national average at the time) that failed to pick up.

4.2 Second Structural Break in 1995

During 1993-98 there was an acceleration of productivity growth in the Australian economy, partly due to a significant increase in business investment.[30] Furthermore, total hours worked as well as employment rose sharply in 1995, in excess of 4% on a year-to-year basis (ABS, 2012a;

[29] These reforms included the deregulation of the financial system, reductions in tariffs and other trade barriers, the removal of most restrictions on foreign investment, etc.

[30] The productivity growth accelerated from a long run average of 1.2% to 2.4% per year during this period of time (Parham, 1999). Business investment as a ratio of GDP rose from a little over 10% in 1992 to more than 14% in 1997 (RBA, 2012).

2012b). The working age population started to show an upward trend for the first time during 1994-95 (ABS, 2012c). This is evident in a strong increase in the labor force participation rate in 1995. Most gains in terms of growth in state share of output percolated to New South Wales (NSW), Queensland (QLD), and Western Australia (WA). A key feature of the mid-1990s is that it marked the secular decline in the importance of manufacturing industries and the rise of services. This structural shift had regional effects because of the spatial distribution of these services. For example, financial and insurance services were concentrated in NSW and VIC while 'other business services' that include information, media and telecommunications; rental, hiring and real estate services; professional scientific and technical services; administrative and support services are concentrated in QLD and SA; and, to some extent, retail and wholesale trade are concentrated in QLD, NSW, and VIC. These sectors attracted large private investments, driving up the capital-labour ratios and real wages in these sectors. This services boom is reflected in rising relative prices in Adelaide, Brisbane, and Sydney.

Another major development around this break date is that the Reserve Bank of Australia (RBA) adopted an inflation target of "2-3 percent per annum, on average over the course of business cycle" in 1993. This target was then formalised in an agreement between the Howard Government and the Governor of RBA in 1996. Furthermore, the federal government confirmed the *de facto* independence of RBA. The change in the stance of monetary policy was immediately visible through a sharp rise in both nominal and real interest rates in 1995 in response to the rising inflationary expectations. The mortgage rate on new housing loans reached a record high of 10.5%. Other interest rates such as small business loans and swap rate with 3-year maturity rose to record high levels in two decades. These developments are likely to have differential effects in different regions of the country. [31]

[31] Using city-level price data for the U.S., Fielding and Shields (2011) show that monetary policy has differential effects on city prices depending on city-specific economic characteristics including the composition of local industry, bank size, house prices, and the age distribution of the population.

Furthermore, the mid-1990s also marked the beginning of an upward trend in the inflow of long term and permanent immigrants to Australia (ABS, 2012d). NSW and VIC already had a substantial stock of immigrants. Since the mid-1990s, QLD became the new destination for overseas immigrants (ABS, 2012e). The consequent rise in the demand for residential dwellings contributed to the rise in housing prices, particularly in Brisbane, which in turn had a positive impact on Brisbane CPI.

4.3 Third Structural Break in 2007

The period 2006-07 marked the last phase of expansion of the world economy before the global economic crisis. Australian economy enjoyed a growth-surge since the mild recession of 2001-02 up until this time, mainly due to exploding commodity prices driven by the growth of China and, to some extent, India. The growth of output was accompanied by rising inflation during this period. The state share of output rose sharply for the mining states of QLD and WA while the shares of all other states went down. The GDP share of mining industry increased from around 4% in 2002-04 to about 10% in 2007-08 (ABS, 2012f). Thus, the external commodity price boom seems to have shifted the mean relative prices upward in Brisbane and Perth. In contrast, mean relative prices either fell (e.g. in Hobart, Melbourne, and Sidney) or changed little (e.g. in Adelaide, Canberra). In general, the volatility of relative prices decreased significantly in almost all cities during the post-2007 period. This seems to reflect the success of inflation targeting of the monetary policy by the central bank.

5. Conclusion

Applying panel unit root test procedures with and without structural breaks to quarterly CPI data for seven Australian cities from 1972Q1 to 2011Q4, we examine the dynamic behaviour of relative prices across these cities. We find overwhelming evidence of convergence in city relative prices. Three common structural break dates are endogenously determined at 1985, 1995, and

2007. Further, correcting for two potential biases, namely Nickell bias and time aggregation bias, we obtain half-life estimates of 2.3 – 3.8 quarters that are much shorter than those reported by CS. Thus, we conclude that both structural breaks and bias corrections are important to obtain shorter half-life estimates. These results are robust to shorter sample period and lower frequency data (annual in our case) with longer sample period.

The fiscal expansion and international trade liberalisation along with the floating of the Australian exchange rate in the mid-1980s; the productivity growth and changes in monetary policy stance during the mid-1990s; and, finally, the international commodity price boom during the mid-2000s seem to have been the triggers of the structural breaks in relative prices identified by our analysis. These events have differential long-term effects on relative prices across cities primarily due to the differences in the mixture of industries and economic activities.

TABLE 1

Panel Unit Root Test Results

	Estimated test statistic	5% critical value
	(1)	(2)
Sample period: 1972Q1 – 2011Q4		
No structural breaks	-5.34	-4.47
Three structural breaks (Break dates: 1985Q1, 1995Q2, and 2007Q2)	-11.21	-10.50
Sample period: 1972Q3 – 1999Q1		
No structural breaks	-6.68	-4.38
Two structural breaks (Break dates: 1988Q1 and 1997Q1)	-9.90	-8.52

Note: The critical values are generated from Monte Carlo simulations.

TABLE 2

Panel Feasible Generalised Least Square (FGLS) Estimation of ϱ and Implied Half-life

	No bias correction		Combined Nickell and time aggregation bias correction					
			Data collection frequency = 3		Data collection frequency = 2		Data collection frequency = 1	
	$\hat{\rho}$	Half-life	$\hat{\rho}$	Half-life	$\hat{\rho}$	Half-life	$\hat{\rho}$	Half-life
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Sample period: 1972Q1 – 2011Q2								
No structural breaks	0.919	8.2	0.896	6.3	0.904	6.8	0.931	9.7
Three structural breaks (Break dates: 1985Q1, 1995Q2, and 2007Q2)	0.829	3.7	0.739	2.3	0.758	2.5	0.835	3.8
Sample period: 1972Q3 – 1999Q1								
No structural breaks	0.877	5.3	0.838	3.9	0.850	4.2	0.893	6.1
Two structural breaks (Break dates: 1988Q1 and 1997Q1)	0.819	3.5	0.730	2.2	0.751	2.4	0.822	3.5

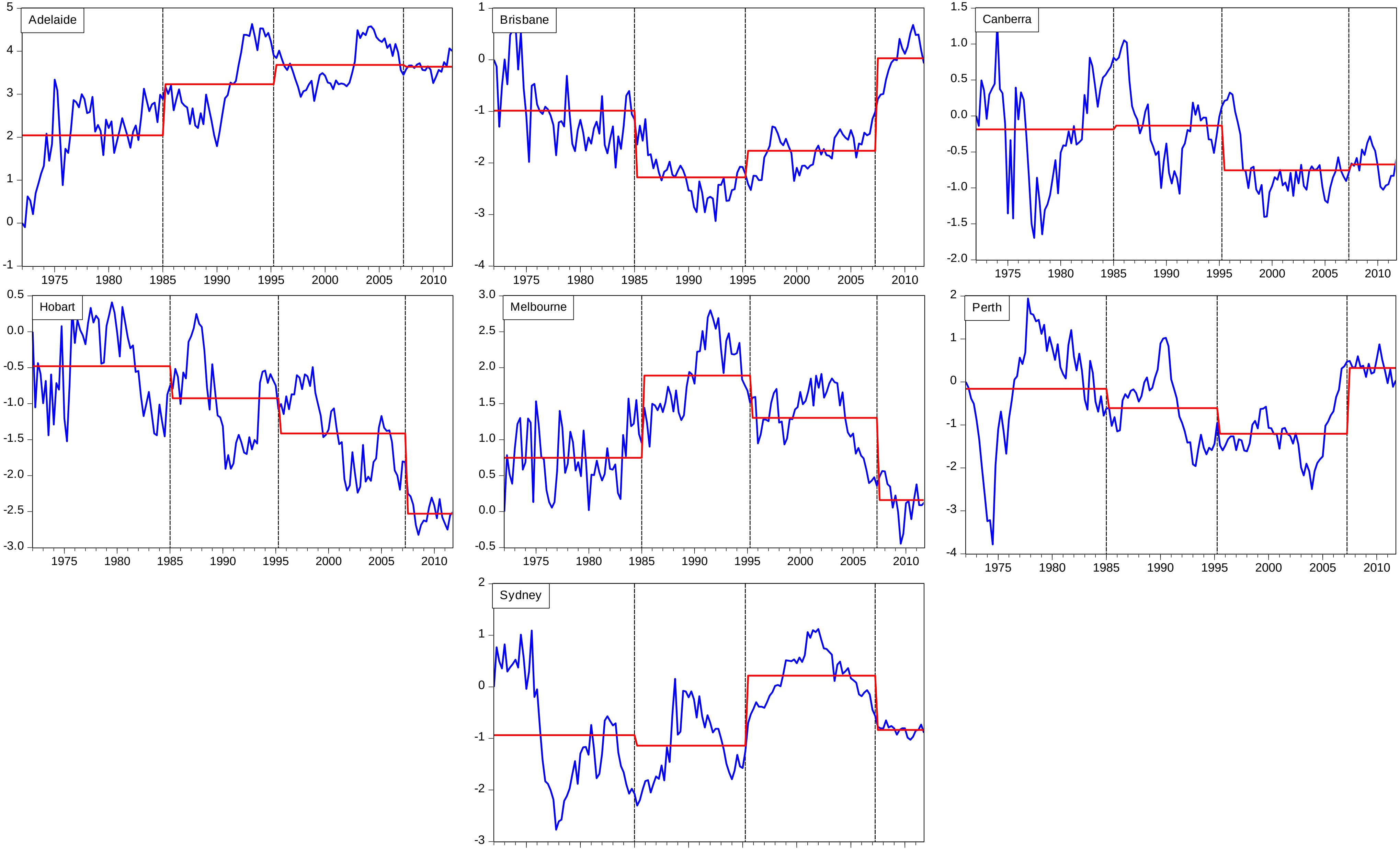
TABLE 3

Panel Unit Root Test Results and Estimated Half-lives with Annual Data

	Estimated test statistic	5% critical value	Autoregressive coefficients and implied half-life					
			No bias correction		Nickell and time aggregation bias correction			
					Data collection frequency = 4		Data collection frequency = 12	
			$\hat{\rho}$	Half-life	$\hat{\rho}$	Half-life	$\hat{\rho}$	Half-life
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Sample period: 1949 – 2011								
No structural breaks	-6.52	-4.52	0.926	9.0	0.931	9.7	0.929	9.4
Three structural breaks (Break dates: 1964, 1984, and 2005)	-11.02	-9.47	0.857	4.5	0.807	3.2	0.774	2.7
Sample period: 1972 – 2011								
No structural breaks	-4.93	-4.10	0.858	4.5	0.872	5.1	0.865	4.8
Three structural breaks (Break dates:1987, 1995, and 2006)	-12.23	-9.73	0.584	1.3	0.417	0.8	0.391	0.7
Sample period: 1972 – 1998								
No structural breaks	-7.02	-4.73	0.775	2.7	0.771	2.7	0.762	2.6
Two structural breaks (Break dates: 1980 and 1987)	-11.26	-9.09	0.591	1.3	0.451	0.9	0.436	0.8

FIGURE 1

Relative Prices in seven Australian Cities with Three Structural Breaks in 1985: Q1, 1995:Q2, and 2007:Q2 (Sample Period: 1972:1 – 2011:4)



References

Australian Bureau of Statistics. 2012a. *Labour force, Australia*, 'Table 12. Labour force status by sex – states and territories', Time series spreadsheet, cat. no. 6202.0, viewed 16, December 2012: http://abs.gov.au/AUSSTATS/abs@.nsf/DetailsPage/6202.0Nov%202012?OpenDocument

Australian Bureau of Statistics, 2012b, *Labour force, Australia*, 'Table 20. Aggregate monthly hours worked by State – trend and seasonally adjusted', Time series spreadsheet, cat. no. 6202.0, viewed 16 December 2012: http://abs.gov.au/AUSSTATS/abs@.nsf/DetailsPage/6202.0Nov%202012?OpenDocument

Australian Bureau of Statistics, 2012c, *Australian Demographic Statistics*, 'Table 59. Estimated resident population by single year of age, Australia', Time series spreadsheet, cat. no. 3101.0, viewed 17 December 2012: http://www.abs.gov.au/AUSSTATS/abs@.nsf/DetailsPage/3101.0Jun%202012?OpenDocument.

Australian Bureau of Statistics, 2012d, *Australian Demographic Statistics*, 'Table 1. Population change, Summary – Australia ('000)', Time series spreadsheet, cat. no. 3101.0, viewed 18 December 2012: http://www.abs.gov.au/AUSSTATS/abs@.nsf/DetailsPage/3101.0Jun%202012?OpenDocument.

Australian Bureau of Statistics, 2012e, *Australian Demographic Statistics*, 'Table 2. Population change, Components – states and territories (number)', Time series spreadsheet, cat. no. 3101.0, viewed 18 December 2012: http://www.abs.gov.au/AUSSTATS/abs@.nsf/DetailsPage/3101.0Jun%202012?OpenDocument.

Australian Bureau of Statistics, 2012f, *Australian National Accounts: State Accounts*, 2011-12, 'Table 10. Expenditure, income and industry components of gross domestic product, Australia, chain volume measures and current prices', Time series spreadsheet, cat. no. 5220.0, viewed 18 December 2012, http://www.abs.gov.au/AUSSTATS/abs@.nsf/DetailsPage/5220.02011-12?OpenDocument

Bai, J. and Perron, P. (1998), 'Estimating and Testing Linear Models with Multiple Structural Changes', *Econometrica*, **66**, 47 – 78.

Bai, J. and Perron, P. (2003), 'Computation and Analysis of Multiple Structural Change Models', *Journal of Applied Econometrics*, **18**, 1-22.

Bai, J. and Perron, P. (2006), 'Multiple Structural Change Models: A Simulation Analysis', in Corbae, D., Darlauf, S., and Hansen, B.E. (eds), *Econometric Theory and Practice: Frontiers of Analysis and Applied Research*. Cambridge University Press, New York, 212-237.

Balassa, B. (1964), 'The Purchasing Power Parity Doctrine: A Reappraisal', *Journal of Political Economy*, **72**, 584 – 596.

Basher, S.A. and Carrion-i-Silvestre, J. L. (2009), 'Price Level Convergence, Purchasing Power Parity and Multiple Structural Breaks in Panel Data Analysis: An Application to U.S. Cities', *Journal of Time Series Econometrics*, **1**, 1-36.

Basher, S.A. and Carrion-i-Silvestre, J. L. (2011), 'Measuring Persistence of U.S. City Prices: New Evidence from Robust Tests', *Empirical Economics*, **41**, 739-745.

Breitung, J. and Candelon, B. (2005), 'Purchasing Power Parity during Currency Crises: A Panel Unit Root Test under Structural Breaks', *Review of World Economics* (*Weltwirtschaftliches Archiv*), **141**, 124-140.

Busetti, F., Fabiani, S., and Harvey, A. (2006), 'Convergence of Prices and Rates of Inflation', *mimeo*, Cambridge University.

Campbell, J. and Perron, P. (1991), 'Pitfalls and Opportunities: What Macroeconomists Should Know About Unit Roots', *NBER Macroeconomic Annual 1990,* 141-201.

Capling, A. (2000), 'The "enfant terrible": Australia and the Reconstruction of the Multilateral Trade System', Australian *Economic History Review,* **40**, 21.

Cecchetti, S. G., Mark, N. C. and Sonora, R. (2002), 'Price Convergence among United States Cities', *International Economic Review*, **43**, 1081-1098.

Chaudhuri, K. and Sheen, J. (2004), 'Purchasing Power Parity across States and Goods Within Australia', *The Economic Record*, **80**, 314 – 329.

Chen, L. L. And Devereux, J. (2003), 'What Can US City Price Tell Us about Purchasing Power Parity?', *Journal of International Money and Finance*, **22**, 213-222.

Chmelarova, V. and Nath, H.K. (2010), 'Relative Price Convergence among US Cities: Does the Choice of Numeraire City Matter?', *Journal of Macroeconomics*, **32**, 405 – 414.

Choi, C.Y., Mark, N. C., and Sul, D. (2006), 'Unbiased Estimation of the Half-Life to PPP Convergence in Panel Data', *Journal of Money, Credit and Banking*, **38**, 921–38.

Carrion-I-Silvestre, J.L., del Barrio, T., Lopez-Bazo, E. (2004), 'Evidence on the Purchasing Power Parity in a Panel of Cities', *Applied Economics*, **36**, 961–966.

Crucini, M. J. and Shintani, M. (2008), 'Persistence in Law of One Price Deviations: Evidence from micro-data', *Journal of Monetary Economics*, **55**, 629 – 644.

Culver, S. E. And Papell, D. H. (1995) 'Real Exchange Rates under the Gold Standard: Can They be Explained by the Trend Break Model', *Journal of International Money and Finance*, **14**, 539-548.

Culver, S. E. and Papell, D. H. (1999), 'Panel Evidence of Purchasing Power Parity Using Intranational and International Data', *Mimeo*, University of Houston.

Dornbusch, R. and Vogelsang, T. (1991), 'Real Exchange Rates and Purchasing Power Parity', in de Melo, J. and Sapir, A. (eds), *Trade Theory and Economic Reform: North, South, and East, Essays in Honor of Bela Balassa*. Basil Blackwell, Cambridge MA, 3-24.

Engel, C., and Rogers, J.H. (1996), 'How Wide is the Border?', *American Economic Review*, **86**, 1112–1125.

Hegwood, N. D. and Papell, D.H. (1998), 'Quasi Purchasing Power Parity', *International Journal of Finance & Economics*, **3**, 279-89.

Hegwood, N. D. and Papell, D.H. (2002), 'Purchasing Power Parity under the Gold Standard', *Southern Economic Journal*, **69**, 72-91.

Hegwood, N.D. and Papell, D.H. (2007), 'Are Real GDP Levels Trend, Difference, or Regime-Wise Trend Stationary? Evidence from Panel Data Tests Incorporating Structural Change', *Southern Economic Journal*, **74**, 104-113.

Hegwood, N. H. And Nath, H. K. (2012), 'Structural Breaks and Relative Price Convergence among U.S. Cities', *SHSU Economics & International Business Working Paper*, 12-04.

Huang, H.C, Lin, P.C., and Yeh, C.C. (2011), 'Price Level Convergence across Cities? Evidence from Panel Unit Root Tests', *Applied Economics Letters*, **18**, 87-93.

Huang, H.C., Liu, W.H, and Yeh, C.C. (1912) 'Convergence in Price Levels across US Cities', *Economics Letters*, **114**, 245-248.

Im, K.S., Lee, J., and Tieslau, M. (2005), 'Panel LM Unit-root Tests with Level Shifts', *Oxford Bulletin of Economics and Statistics*, **67**, 393-419.

Im, K.S., Pesaran, M.H., Shin, Y. (2003), 'Testing for Unit Roots in Heterogeneous Panels', *Journal of Econometrics*, **115**, 53–74.

Imbs, J., Haroon M., Ravn, M. O., and Rey, H. (2005), 'PPP Strikes Back: Aggregation and the Real Exchange Rate', *Quarterly Journal of Economics*, **120**, 1-43.

Levin, A., Lin, C., and Chu, J. (2002), 'Unit Root Tests in Panel Data: Asymptotic and Finite-Sample Properties', *Journal of Econometrics*, **108**, 1-24.

Lin, C. and Lee, Y. (2010), 'Understanding which Events Occur in Structural Breaks by Re-examining the PPP Hypothesis for G7 Countries', *International Research Journal of Finance and Economics*, **51**.

Liu, J., Wu, S., and Zidek, J.V. (1997), 'On Segmented Multivariate Regressions', *Statistica Sinica*, **7**, 97 – 525.

Lumsdaine, R. and Papell, D.H. (1997), 'Multiple Trend Breaks and the Unit Root Hypothesis', *Review of Economics and Statistics*, **79**, 212-218.

Murray, C. J. and Papell, D.H. (2000), 'Testing for Unit Roots in Panels in the Presence of Structural Change with an Application to OECD Unemployment', *Nonstationary Panels, Panel Cointegration and Dynamic Panels*, **15**, 223-238.

Narayan, P. K. (2008), 'The Purchasing Power Parity Revisited: New Evidence for 16 OECD Countries from Panel Unit Root Tests with Structural Breaks', *International Financial Markets, Institutions & Money*, **18**, 137-146.

Nath, H.K. and Sarkar, J. (2009), 'Unbiased estimation of the half-life to price index convergence among US cities', *Journal of Money, Credit and Banking*, **41**, 1041-1046.

Ng, S. and Perron, P. (1995), 'Unit Root Tests in ARMA Models with Data Dependent Methods for the Selection of the Truncation Lag', *Journal of the American Statistical Association*, **90**, 268-281.

Nickell, S. (1981), 'Biases in Dynamic Models with Fixed Effects', *Econometrica*, **49**, 1417-1426.

Obstfeld, M. (1993), 'Model Trending Real Exchange Rates', *Center for International and Development Economic Research Working Paper* No. C93-011.

Papell, D.H. (2002), 'The Great Appreciation, the Great Depreciation, and the Purchasing Power Parity Hypothesis', *Journal of International Economics*, **57**, 51-82.

Papell, D. H. and Prodan, R. (2006), 'Additional Evidence of Long-Run Purchasing Power Parity with Restricted Structural Change', *Journal of Money, Credit and Banking*, **38**, 1329 – 1349.

Parsley, D. C. And Wei, S. J. (1996), 'Convergence to the Law of One Price Without Trade Barriers or Currency Fluctuations', *Quarterly Journal of Economics*, **111**, 1211-1236.

Perron, P., (1989), 'The Great Crash, the Oil Price Shock, and the Unit-Root Hypothesis', *Econometrica*, **57**, 1361-1401.

Perron, P. and Vogelsang, T. (1992), 'Nonstationarity and Level Shifts with an Application to Purchasing Power Parity', *Journal of Business and Economic Statistics*, **10**, 301-320.

Phillips, P.C.B. and Sul, D., (2007), 'Transition modeling and econometric convergence tests', *Econometrica*, **75**, 1771–1855.

Prodan, R. (2008), 'Potential Pitfalls in Determining Multiple Structural Changes With an Application to Purchasing Power Parity', *Journal of Business & Economic Statistics*, **26**, 50 – 65.

Reserve Bank of Australia (RBA), (2012), *The Australian economy and financial markets*: Chart Pack, December, 2012.

Rogoff, K. (1996), 'The Purchasing Power Parity Puzzle', *Journal of Economic Literature*, **XXXIV**, 647-668.

Samuelson, P.A. (1964), 'Theoretical Notes on Trade Problems', *Review of Economics and Statistics*, **46**, 145-154.

Sonora, R.J., (2005), 'City CPI convergence in Mexico', *Review of Development Economics* **9**, 359–367.

Sonora, R.J. (2009), 'City relative price convergence in the USA with structural break(s)', *Applied Economics Letters*, **16**, 939-944.

Yao, Y.C. (1988), 'Estimating the Number of Change-Points via Schwarz' Criterion', *Statistics and Probability Letter*, **6**, 181-189.