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# Modelling the Balassa-Samuelson Effect in Australia

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## Abstract

This paper examines the Balassa-Samuelson hypothesis in Australia using the ARDL cointegration framework. Evidence was found of a significant long-run relationship between real exchange rate and Australia-US productivity differential during the period of 1950-2003. The results indicate that a one per cent increase in labour productivity in Australia relative to the US will lead to 5.6 per cent appreciation in the real exchange rate of Australia. The estimated coefficient for the error correction term is -0.1983 and is highly significant, indicating that the deviation from the long term real exchange rate equilibrium path is corrected by nearly 20 per cent over the following year. The Author suspects that the elasticity coefficient is "over-estimated" due to the exclusion of relevant explanatory variables in the analytical model. The real exchange rate movements are affected by real fundamentals and policy-induced shifts in its real fundamentals. The fundamentals include the terms of trade, government expenditure, real interest rate differentials, net foreign liabilities among others along with labour productivity differential.

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**JEL Classification:** C22, F11, F31.

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## 1. INTRODUCTION

The Balassa-Samuelson (B-S) effect[2] model was developed simultaneously by Balassa (1964) and Samuelson (1964), working independently. In broad terms, the B-S effect can be construed as either of two related things: (1) that consumer price levels are systematically higher[3] in wealthier countries than in poorer ones (the "Penn effect"), (2) a model predicting (1), based on the assumption that productivity or productivity growth-rates vary more across countries in the traded goods' sectors than in non-traded goods' sectors (the Balassa-Samuelson hypothesis). In this paper we specifically attempt to empirically test proposition (2) above.

The Purchasing Power Parity (PPP) in its absolute form can be expressed as $e = (P / P^*)$, where $e$ is the amount of domestic currency per unit of foreign currency, and $P$ and $P^*$ are the domestic and foreign price levels (* denotes foreign, say US). Thus, PPP theory predicts that, in the long run, relative prices determine the exchange rate; and any deviation of relative prices from the equilibrium exchange rate will be transient and ultimately mean-reverting in the long run. However, according to Balassa (1964) and Samuelson (1964), the persistence of real exchange rate (RER) changes can be attributable to productivity differential in the two economies. Rapid economic growth is accompanied by RER appreciation because of differential productivity growth between traded (T) and nontraded (NT) sectors. Since the differences in productivity increases are expected to be larger in high growth countries, the B-S effect should be more pronounced among fast growing economies[4]. Alternatively, increased productivity induced by technological progress increases factor availability. By reducing the cost and price of tradables, increased productivity makes the tradable sector more competitive and tends to depreciate the RER of the sector. In this situation supply effects of technological progress offset the demand effects according to the *Rybczynski* principle (Edward, 1989, p 48).

RER – the ratio of price of tradables to price of nontradables - is a price that ensures internal and external equilibrium simultaneously. Internal equilibrium implies that the nontradable goods market clears in the current period and is expected to be in equilibrium in future periods. According to this definition, it is assumed that this equilibrium exists where unemployment is at its "natural" level. External equilibrium, however, is attained when the intertemporal budget constraint, that states that the discounted sum of a country's current account has to be equal to zero, is satisfied. The policy issue of "overvaluation/undervaluation" and the resultant existence and magnitude of distortions is discussed in terms of the RER movements. The motivation and benefit of this study can be understood from the role and effects of the exchange rate on the overall economy. The RER:

- Influences macro price ratios such as those between tradable and non-tradable goods, capital goods and labour, and exports and imports.
- Serves as an asset price and thus determines capital flows.
- Partially determines inflation rates through the cost side and as a monetary transmission vector.
- Significantly affects aggregate demand, in both the short and long-run.

[2] Earlier, outlines of the explanation of the effect were provided by Harrod (1933) and Ricardo (1911).
[3] Bhagwati (1984) and Kravis and Lipsey (1983) provide an alternative theory to explain lower price levels in poorer countries.
[4] Japan is a classic example of the B-S effect.

Correspondingly the exchange rate can be targeted toward achieving many policy objectives in the real economy. Five have been of primary importance in recent times:

1. Resource allocation (including employment).
2. Economic development (often in conjunction with commercial and industrial policies).
3. Finance: Control expectations and behavior in financial markets. Exchange rate policy "mistakes" can lead to highly destabilising consequences.
4. External balance, via both "substitution" responses and shifts it can cause in effective demand.
5. Inflation: The exchange rate can serve as a nominal anchor, holding down price increases via real appreciation and/or maintenance by the authorities of a consistently strong rate. It can also serve as an important transmission vector of monetary policy.

The recent appreciation in the Australian bilateral exchange rate and RER has raised questions on the sustainable value of the exchange rate and its long-run impacts. RER appreciation reflects an increase in the domestic cost of producing tradable goods. If there are no changes in relative prices in the rest of the world, this appreciation in RER represents a deterioration of the country's degree of international competitiveness: the country now produces tradable goods in a relatively less efficient way than before. RER depreciation likewise represents an improvement in the country's international competitiveness.

Empirical results on the B-S effect are mixed. Although some negative results were returned, there has been some support for the predictions of the BS-hypothesis in the literature, for instance, Bahmani-Oskooe and Rhee (1996) did find a statistically significant correlation between real exchange rates and relative productivities. Lafrance and Schembri (2000) suggest that the Balassa-Samuelson mechanism may be evident in the productivity and exchange rate changes between the United States and Canada during 1979 to 1999. Bahmani-Oskooe and Nasir (2004), using cointegration and error correction modelling in a sample of 44 countries, found evidence of B-S hypothesis in 32 countries (developed and developing) while the B-S hypothesis failed in 12 less developed economies riddled with trade restrictions, capital controls and other trade barriers.

Drine and Rault (2002) argue that the difficulties of confirming the hypothesis have partly been due to testing particular components of it, and that even where the varying-productivity-real exchange rate (RER) link is established it does not necessarily confirm the BS-hypothesis. The purpose of this paper is to bridge the gap in the time series literature on B-S hypothesis in general and Australia in particular. This paper is organised as follows: The analytical framework is outlined in section 2. In section 3 the time series properties of the variables in the presence of endogenous structural break in data are tested. This is done since the traditional unit root tests suffer from power deficiency when structural break is present in the data. In section 4 the model is estimated by using the Auto Regressive Distributed Lag modelling approach which allows an estimation of the model regardless of whether the variables are I(0) or I(1).

## 2. THE ANALYTICAL FRAMEWORK: THE BALASSA-SAMUELSON HYPOTHESIS REVISITED

Let us consider two small open economies (the foreign country is denoted with an asterisk) producing two goods: a tradable commodity (T) for the world market and a non-tradable commodity (NT) for domestic demand. They use labour ($L$) as input and production is subject to constant returns to scale. The Balassa-Samuelson (BS) effect asserts that if the traded goods' marginal productivity relative to non-traded goods' marginal productivity is increasing faster in the domestic economy than in the rest of the world, then the domestic economy will register an appreciation of its real exchange rate. The B-S theory assumes that the international productivity differences in non-tradeables are negligible. Due to constant returns to scale the marginal productivity of labour is proportional to the average product of labour. In this case, the real exchange rate $\rho$ can be written in terms of the average productivity of labour:

$$\rho = \frac{(Y_T / L_T)}{(Y_T^* / L_T^*)} \qquad (1)$$

The real exchange rate is defined as: $\rho = \frac{P}{eP^*} = \frac{P}{P^*}$, where P and P* are the domestic and foreign price level respectively.

If traded goods' average productivity relative to non-traded goods' average productivity grows faster in the domestic economy than the foreign economy, the domestic economy will experience a RER appreciation.

According to the above discussion, the testable reduced form log-linear specification of the Balassa-Samuelson model can be expressed as follows:

$$LnR_t = \alpha + \beta LnPR_t \qquad (2)$$

where, $R = (P_{Aus} / P_{US}) / e$ and $PR = PROD_{Aus} / PROD_{US}$. $R$ denotes the amount of US dollars per one unit of Australian dollar in real terms, while $PR$ denotes the Australia-US labour productivity differential.

## 3. TESTS FOR TIME SERIES PROPERTIES IN THE PRESENCE OF STRUCTURAL BREAK

In this study, annual data for all series from 1950 to the last available data was used until 2003 from Heston, Summers and Aten (2006) Penn World Table Version 6.2. For $PR$, the real GDP per worker was used (in 2000 international prices) of each country treating the US as the reference country. Equation (2) can be analysed using a cointegration test. Prior to conducting the cointegration test, it is essential to check each time series for stationarity. If a time series is nonstationary, the traditional regression analysis will produce spurious results. Therefore, the unit root test is conducted first. Hence it is imperative to review some of the recently developed models and tests for unit roots which are going to be examined in this paper. A succinct review is given in Appendix 1.

To ascertain the order of integration, the traditional Augmented Dickey-Fuller (ADF) and Phillips-Perron (1988) (PP) unit root test were applied. These tests suggest that all the variables in the model are nonstationary (refer to Table 1). Since the ADF and PP tests suffer from power deficiency in the presence of a structural

break[5], the most comprehensive models of Perron (1997) along with the Zivot and Andrews (1992) model were applied. Perron (1997) includes both *t* (time trend) and $DT_b$ (time at which structural change occurs) in his Innovational Outlier (IO1 and IO2) and Additive Outlier (AO) models. The distinction between the two is worth noting. The IO2 model represents the change that is *gradual* whereas AO model represents the change that is *rapid*.

A summary of the unit root test results is given in Table 1. Of the four models in this category, the Additive Outlier Model (AO) and the Innovational Outlier (IO1) Model are found optimal for *LnR* and *LnPR* on the basis of the Shrestha-Chowdhury (2005) model selection procedure. The results in Table 1 indicate that the unit root hypotheses are rejected at the 5 per cent level of significance for *LnR* by all the tests. The estimated break date corresponds to 1979 for *LnR* and 1984 for *LnPR*. The endogenously determined break dates are plausible with the events occurring in the Australian economy. After a sustained period of appreciation, depreciations of the real exchange rate occurred during 1974-1978 so that the break date for the real exchange rate is picked up in 1979. The recession of the early 1980s in Australia as well as in the US also impacted the productivities in the two countries. The recessionary effect on productivity is captured by the break date of 1985.

Table 1: Unit Root Tests in the Absence and Presence of a Structural Break

| *LnR* | | | | | *LnPR* | | | | |
|---|---|---|---|---|---|---|---|---|---|
| Test | k | TB | Tα=1 | Decision | Test | k | TB | Tα=1 | Decision |
| ADF | 1 | NC | -1.84 | NS | ADF | 0 | NC | -3.44 | NS |
| PP | 1 | NC | -1.25 | NS | PP | 1 | NC | -3.58 | NS |
| IO1 | 8 | 1970 | -4.08 | NS | *IO1* | *1* | *1985* | *-6.18* | *S* |
| IO2 | 8 | 1971 | -6.20 | S | IO2 | 1 | 1984 | -6.30 | S |
| *AO* | *8* | *1979* | *-4.15* | *NS* | AO | 1 | 1951 | -3.32 | NS |
| Zivot-Andrews | 1 | 1977 | -4.27 | NS | Zivot-Andrews | 2 | 1976 | -3.01 | NS |

Notes: S = stationary; NS = nonstationary; NC = not calculated.
The critical values for IO1 for 60 observations are -5.92 and -5.23 and at 1% and 5% respectively.
The critical values for IO2 for 70 observations are -6.32 and -5.59 and at 1% and 5% respectively.
The critical values for AO for 100 observations are -5.45 and -4.83 at 1% and 5% respectively.
The critical values for Zivot-Andrews are -4.93 and -4.42 at 1% and 5% respectively.
The critical values for ADF and PP are -4.14 and -3.49 at 1% and 5% respectively.

## 4. EMPIRICAL FINDINGS

The variables considered in this study are a mix of I(0) ($LnR_t$) and I(1) ($LnPR_t$) series. The cointegration test methods based on Johansen (1991; 1995) and Johansen-Juselius (1990) require that all the variables be of equal degree of integration, i.e., I(1) to yield estimators which are super-consistent, symmetrically distributed, and median-unbiased asymptotically. Theoretically, the mixture of I(0 and I(1) variables does not pose any problems for estimation but the likelihood testing procedure for the cointegrating rank can be sensitive to the presence of stationary variables (Rahbek &

[5]It is widely known that macroeconomic series often experience various breaks in their realisations. This is especially true for transition and emerging market economies, which often experience shocks due to radical policy changes or crises. The examples of policies with break consequences include frequent and significant devaluations, deregulation of both real and financial sectors and policy regime shifts.

Mosconi, 1999). Hence, the ARDL modelling approach was adopted for cointegration analysis in this study.

The main advantage of ARDL modelling lies in its flexibility in that it can be applied "irrespective of whether the regressors are purely *I(0)*, purely *I(1)* or mutually cointegrated" (Pesaran Shin & Smith. 2001, p289-290). Another advantage of this approach is that the model takes sufficient numbers of lags to capture the data generating process in a general-to-specific modelling framework (Laurenceson & Chai 2003, p28). Moreover, a dynamic error-correction model (ECM) can be derived from ARDL through a simple linear transformation (Banerjee et al. 1993, p51). The ECM integrates the short-run dynamics with the long-run equilibrium without losing long-run information. Using the ARDL approach avoids problems resulting from non-stationary time-series data, and typically outperforms alternative approaches to cointegration such as the Phililip and Hansen's Fully Modified Least Squares when the sample size is small (Laurenceson & Chai 2003, p28). This is, in particular, true of the size-power performance of the tests on the long-run parameter. Finally, ARDL modelling is robust against simultaneous equation bias and autocorrelation, provided the orders of the ARDL model are adequately selected on the basis of any model selection criterion.

Thus, the error correction specification of the ARDL model pertaining to equation (2) is given below:

$$\Delta LnR_t = \alpha_0 + \delta_1 LnR_{t-1} + \delta_2 LnPR_{t-1} + \sum_{i=1}^{p} b_i \Delta LnR_{t-i} + \sum_{i=0}^{q} c_i \Delta LnPR_{t-i} + \varepsilon_t \quad (3)$$

The parameters $\delta_i$, i =1, 2, are the long run multipliers. The parameters $b_i, c_i$, are the short run multipliers and $\varepsilon_t$ represent the residuals.

To select the appropriate model in equation (3), several specifications with different lags were tested for statistical significance and for consistency with the cointegration method. The specification used here is the unrestricted intercept with no trend (Case III in Pesaran et al. 2001, p296). The model given in equation (2) was estimated and the optimal model was found to be be $\text{ARDL}[2, 0]$ based on the AIC and SBC model selection criteria[6]. The estimated ARDL model is given in Appendix 2, Table A2.1.

*Estimation of Long Run Coefficients*

The long run relationship between the Australian real exchange rate ($LnR_t$) and the Australia-US labour productivity differential ($LnPR_t$) was investigated by using the 'bounds procedure' developed by Pesaran et al. (2001). The bounds test for examining the presence of a long run relationship can be carried out using the $F-$test where the null hypothesis tests the joint significance of $\delta_1 = \delta_2 = 0$ in equation (3). The $F-$test has a non-standard distribution and is contingent upon: (i) whether variables in the ARDL model are I(0) or I(1); (ii) the number of regressors; (iii) whether the model has an intercept and/or a trend; and (iv) the sample size. Pesaran et al. (2001) computes two sets of critical values which classify regressors into pure I(1), I(0) and

[6] All commonly used model selection criteria (AIC, HQ, SBC etc.) are functions of residual sums of squares and are asymptotically equivalent (Judge et al., 1985: 869).

mutually cointegrated categories[7]; and these critical values are generated for sample sizes of 500 and 1000 observations with replications of 20,000 and 40,000 respectively. Narayan (2005) also calculate critical values for small sample sizes ranging from 30-80 observations. These critical upper bound and lower bound values are also reported in Table 2.

Based on the 'bounds test' (given in Table 2), the computed *F*-statistic is 2.79, which is below the lower critical bound (*LCB)* at the 10 per cent significance level.

In applying the *F*-test one must be careful about the number of lags chosen for each of the first differenced variables as the results are sensitive to the lag lengths. Secondly, the LCB and UCB are calculated for large number of observations (500 and 1000) which may be problematic in small samples as intjis study. Therefore, following Kremers et al. (1992) reliance is placed upon the significance of the error correction term as a useful and efficient way of establishing cointegration.

Table 2: Bounds Test for Cointegration

| | | |
|---|---|---|
| Computed *F*-Statistics ( $F_{2,\ 44}$ ) | 2.79 | |
| Critical Bounds (10 per cent) | *LCB*: 4.04♣. 4.16♠ | *UCB*: 4.78♣. 4.93♠ |
| Critical Bounds (5 per cent) | *LCB*: 4.94♣. 5.13♠ | *UCB*: 5.73♣. 6.05♠ |

Notes: ♣Critical Bounds are from Pesaran et al. (2001, p 300) Table CI (iii) Case III.
♠Critical Bounds are from Narayan (2005, p1988) Case III.

The estimated long-run coefficients for the ARDL model are given in Table 3. In the long-run, a one per cent increase in the productivity differential will lead to 5.58 per cent appreciation of Australian real exchange rate. The empirical result shows that the productivity differential has a statistically significant positive effect on the movement of the Australian real exchange rate. Thus, the Balassa-Samuelson proposition is vindicated.

Table 3: Estimated Long Run Coefficients for Equation 2: ARDL (2, 0)
Dependent Variable: $LnR_t$

| Variables | Coefficient | t-ratio | P-value |
|---|---|---|---|
| $LnPR_t$ | 5.58* | 2.15 | 0.037 |
| *Intercept* | 1.696** | 1.90 | 0.064 |

Note: *, ** denote significant at the 5% and 10% respectively.

This high elasticity value is due to the probable misspecification (underfitting) of the model in equation (2)[8]. The crux of the B-S hypothesis is premised on the

[7] If the computed F-statistic is greater than the upper critical bound (UCB), the regressors are I(1); if the F-statistic is less than the lower critical bound (LCB), the regressors are I(0); and if the F-statistic falls within the interval of LCB and UCB, inference is inconclusive and order of integration between the underlying variables are required for a conclusive inference Pesaran et al. (2001, p299).

[8] On average, the estimated coefficient will overestimate the true coefficient which explains the high coefficient estimate obtained here. As an illustration, suppose the true model is: $Y_t = \beta_1 + \beta_2 X_{2t} + \beta_3 X_{3t} + u_t$ but we estimate the following model: $Y_t = \alpha_1 + \alpha_2 X_{2t} + v_t$. It can be shown that, $E(\hat{\alpha}_2) = \beta_2 + b_{32}\beta_3$, where $b_{23}$ is the slope coefficient of regression of $X_3$ on the included

proposition that productivity differential alone is the determinant of the real exchange of a country. However, in recent times, researchers have been trying to explain the long run adjustment of real exchange rates by a host of other factors (called fundamentals) such as real interest rate differentials, productivity differentials, capital accumulation, cumulated current account balances, the level and composition of government spending, saving, trade openness and the terms of trade etc. Blundell-Wignall, Fahrer and Heath (1993) have identified three statistically significant determinants of the Australian real exchange rate. These are: terms of trade; net foreign liabilities; and real long-term interest differentials. This result is also confirmed by the findings of Gruen and Wilkinson (1994). The authors estimate that a real exchange rate appreciation of about 0.3 to 0.5 per cent is associated with a one per cent improvement in the terms of trade, while an appreciation of about 2 to 3.5 per cent is associated with an increase of one percentage point in the differential between Australian and world real interest rates. In contrast, Bagchi Chortareas and Miller (2004, p84) find "…the terms of trade prove quantitatively more important in explaining the long-run real exchange rate than the real interest rate differential.[9]" Tarditi (1996) augmented the Blundell-Wignall et al. (1993) model by including terms of trade, cumulated current account balance (proxy for net foreign liability), yield curve differential (instead of long term interest rate differential) and fiscal deficit as a proportion of GDP and found them to be significantly affecting the Australian trade-weighted real exchange rate.

In testing the B-S effect in 44 countries, including Australia, Bahmani-Oskooee and Nasir (2004) found the productivity differential coefficient to be 0.97 per cent compared to this study's value of 5.58 per cent. This value was found to be low given that the determinants of the Australian real exchange rate are numerous and significant as shown by the discussion above. The result is puzzling and the reason that the results are so vastly different[10]is not immediately apparant.

Various diagnostic analyses for serial correlation, heteroskedasticity, normality of residuals and other tests are reported in Appendix 2, Table A2.1. These tests indicate that the specified model passes all the diagnostic tests. As can be seen, there is no evidence of autocorrelation and the model passes the test of normality. Furthermore, Figure A2.1 of Appendix 2 indicates the stability of both long and short run coefficients since the residuals lie within the upper and lower bounds of the critical values.

### *Short Run Dynamics*

The short run dynamics and the long run equilibrium for the estimated ARDL model is given in Table 4. The short run adjustment process is measured by the error correction term (ECM). The ECM indicates how quickly variables adjust and return to

---

variable $X_2$. The bias due to omission of other variables can be shown in an analogous way. It can also be shown that $Var(\hat{\alpha}_2)$ will be biased as well. Refer to Kmenta (1985, p443-446).

[9] A one per cent increase in terms of trade leads to a real appreciation of nearly 1.4 per cent of the Australian dollar while a one per cent increase in interest rate differential appreciates the Australian dollar by only 0.04 per cent in real terms.

[10] Bahmani-Oskooee and Nasir (2004) use data from Penn World Table (Mark 5) where 1985 international prices are used as opposed to 2000 international prices in Penn World Table Version 6.2.

Table 4: Error Correction for the Selected ARDL Model: ARDL (2, 0)
Dependent Variable: $\Delta LnR_t$

| Variables | Coefficient | P-value |
|---|---|---|
| $\Delta LnR_{t-1}$ | 0.752* | 0.000 |
| $\Delta LnPR_t$ | 0.134** | 0.030 |
| *Intercept* | 0.093 | 0.108 |
| $ECM_{t-1}$ | -0.1983* | 0.007 |
| R-Squared | 0.215 | |
| AIC | 71.996 | |
| SBC | 68.133 | |
| Durbin-Watson | 2.010 | |
| F(3, 47) | 4.031 | 0.012 |

Note: *, **, denote significant at the 1% and 5% respectively.

equilibrium and the coefficient of ECM should carry the negative sign and be statistically significant. As shown in Table 4, the estimated coefficient for ECM is equal to -0.1983 for the specified model and is highly significant, indicating that the deviation from the long term real exchange rate equilibrium path is corrected by nearly 20 per cent over the following year. In other words, the adjustment process is very high. The statistical significance of the ECM further confirms the presence of long run equilibrium between the real exchange rate of Australia and the productivity differential of Australia and the US.

## 5. SUMMARY AND CONCLUSION

The purpose of this paper was to test the B-S productivity bias hypothesis using time series data from Australia. This study not only fills in a void on this topic in Australia but also adds to the limited number of time series studies on this subject. The time series properties of the variables in the presence of a structural break were tested since traditional unit root tests (ADF and PP) suffer from power deficiency and found that the variables are a mixture of I(0) and I(1) variables. A flexible, robust econometric framework called the ARDL modelling was applied to estimate long and short term relationships among variables.

The B-S model was derived based on some simplifying assumptions[11] (single factor of production, constant returns to scale, constancy of terms of trade thus ignoring the demand side of the economy). The empirical results of this study support the B-S proposition that there is a strong, positive link between the real exchange rate and productivity differential in Australia during the period of 1950-2003. This result also confirms the absence of the *Rybczynski* effect in the Australian economy. In this study, it was found that a one per cent increase in labour productivity in Australia relative to the US will lead to 5.6 per cent increase in the real exchange rate of Australia. The Author believes that the elasticity coefficient is "over-estimated" due to the exclusion of relevant explanatory variables since the dynamics and the determinants of the real exchange rate movements are numerous such as the terms of

[11] Obstfeld and Rogoff (1996: 210-216) derive the same result utilising a model with two productive factors (K, L) and perfect capital mobility among economies. In an extension of their basic model, Obstfeld and Rogoff (1996) generalise the B-S result by including (1) a third factor of production, namely skilled labour S, to produce tradables and nontradables; and (2) internationally immobile capital.

trade, interest rate differentials and net foreign liabilities among others. The estimated coefficient for ECM is equal to -0.1983 and is highly significant, indicating that the deviation from the long term real exchange rate equilibrium path is corrected by nearly 20 per cent over the following year. In other words, the adjustment process is very high. The statistical significance of the ECM further confirms the presence of long run equilibrium between the real exchange rate and the productivity differential of Australia and the US.

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## Appendix 1

### A Review of Unit Root Tests with Endogenous Structural Break

Traditional tests for unit roots (such as Dickey-Fuller, Augmented Dickey-Fuller and Phillips-Perron, 1988) have low power in the presence of a structural break. Perron (1989) demonstrated that, in the presence of a structural break in time series, many perceived nonstationary series were in fact stationary. Perron (1989) re-examined Nelson and Plosser (1982) data and found that 11 of the 14 important US macroeconomic variables were stationary when a known exogenous structural break is included[12]. Perron (1989) allows for a one time structural change occurring at a time $T_B$ ($1 < T_B < T$), where T is the number of observations.

The following models were developed by Perron (1989) for three different cases. Notations used in equations A1- A16 are the same as in the papers quoted.

Null Hypothesis:

Model A $y_t = \mu + dD(TB)_t + y_{t-1} + e_t$ (A1)

Model B $y_t = \mu_1 + y_{t-1} + (\mu_2 - \mu_1)DU_t + e_t$ (A2)

Model C $y_t = \mu_1 + y_{t-1} + dD(TB)_t + (\mu_2 - \mu_1)DU_t + e_t$ (A3)

where $D(TB)_t = 1$ if $t = T_B + 1$, 0 otherwise, and

$DU_t = 1$ if $t > T_B$, 0 otherwise.

Alternative Hypothesis:

Model A $y_t = \mu_1 + \beta t + (\mu_2 - \mu_1)DU_t + e_t$ (A4)

Model B $y_t = \mu + \beta_1 t + (\beta_2 - \beta_1)DT_t^* + e_t$ (A5)

Model C $y_t = \mu_1 + \beta_1 t + (\mu_2 - \mu_1)DU_t + (\beta_2 - \beta_1)DT_t + e_t$ (A6)

where $DT_t^* = t - T_B$ , if $t > T_B$, and 0 otherwise.

Model A permits an exogenous change in the level of the series whereas Model B permits an exogenous change in the rate of growth. Model C allows change in both. Perron (1989) models include one known structural break. These models cannot be applied where such breaks are unknown. Therefore, this procedure is criticised for assuming the known break date which raises the problem of pre-testing and data mining regarding the choice of the break date (Maddala & Kim 2003). Further, the choice of the break date can be viewed as being correlated with the data.

### Unit Root Tests in the Presence of a Single Endogenous Structural Break

Despite the limitations of Perron (1989) models, they form the foundation of subsequent studies that will be discussed hereafter. Zivot and Andrews (1992), Perron and Vogelsang (1992), and Perron (1997) among others have developed unit root test methods which include one endogenously determined structural break. Here these models are briefly reviewed and detailed discussions are found in the cited works.

Zivot and Andrews (1992) models are as follows:

Model with Intercept

$$y_t = \hat{\mu}^A + \hat{\theta}^A DU_t(\hat{\lambda}) + \hat{\beta}^A t + \hat{\alpha}^A y_{t-1} + \sum_{j=1}^{k} \hat{c}_j^A \Delta y_{t-j} + \hat{e}_t \quad \text{(A7)}$$

[12] However, subsequent studies using endogenous breaks have countered this finding with Zivot and Andrews (1992) concluding that 7 of these 11 variables are in fact nonstationary.

Model with Trend

$$y_t = \hat{\mu}^B + \hat{\beta}^B t + \hat{\gamma}^B DT_t^*(\hat{\lambda}) + \hat{\alpha}^B y_{t-1} + \sum_{j=i}^{k} \hat{c}_j^B \Delta y_{t-j} + \hat{e}_t \quad \text{(A8)}$$

Model with Both Intercept and Trend

$$y_t = \hat{\mu}^C + \hat{\theta}^C DU_t(\hat{\lambda}) + \hat{\beta}^C t + \hat{\gamma}^C DT_t^*(\hat{\lambda}) + \hat{\alpha}^C y_{t-1} + \sum_{j=1}^{k} \hat{c}_j^C \Delta y_{t-j} + \hat{e}_t \quad \text{(A9)}$$

where, $DU_t(\lambda)$ = 1 if t > $T\lambda$, 0 otherwise;

$DT_t^*(\lambda) = t - T\lambda$ if $t > T\lambda$, 0 otherwise.

The above models are based on the Perron (1989) models. However, these modified models do not include $DT_b$.

On the other hand, Perron and Vogelsang (1992) include $DT_b$ but exclude $t$ in their models. Perron and Vogelsang (1992) models are given below:

Innovational Outlier Model (IOM)

$$y_t = \mu + \delta DU_t + \theta D(T_b)_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t \quad \text{(A10)}$$

Additive Outlier Model (AOM) – Two Steps

$$y_t = \mu + \delta DU_t + \tilde{y}_t \quad \text{(A11)}$$

and

$$\tilde{y}_t = \sum_{i=0}^{k} w_i D(T_b)_{t-i} + \alpha \tilde{y}_{t-1} + \sum_{i=1}^{k} c_i \Delta \tilde{y}_{t-i} + e_t \quad \text{(A12)}$$

$\tilde{y}$ in the above equations represents a detrended series $y$.

Perron (1997) includes both $t$ (time trend) and $DT_b$ (time at which structural change occurs) in his Innovational Outlier (IO1 and IO2) and Additive Outlier (AO) models.

Innovational Outlier Model allowing one time change in intercept only (IO1):

$$y_t = \mu + \theta DU_t + \beta t + \delta D(T_b)_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t \quad \text{(A13)}$$

Innovational Outlier Model allowing one time change in both intercept and slope (IO2):

$$y_t = \mu + \theta DU_t + \beta t + \gamma DT_t + \delta D(T_b)_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t \quad \text{(A14)}$$

Additive Outlier Model allowing one time change in slope (AO):

$$y_t = \mu + \beta t + \delta DT_t^* + \tilde{y}_t \quad \text{(A15)}$$

where $DT_t^*$ = 1(t > $T_b$)(t – $T_b$)

$$\tilde{y}_t = \alpha \tilde{y}_{t-1} \sum_{i=1}^{k} c_i \Delta \tilde{y}_{t-i} + e_t \quad \text{(A16)}$$

The Innovational Outlier models represent the change that is gradual whereas Additive Outlier model represents the change that is rapid. All the models considered above report their asymptotic critical values.

More recently, additional test methods have been proposed for unit root test allowing for multiple structural breaks in the data series (Lee & Starzicich, 2003; Lumsdaine & Papell 1997) which will not be discussed here.

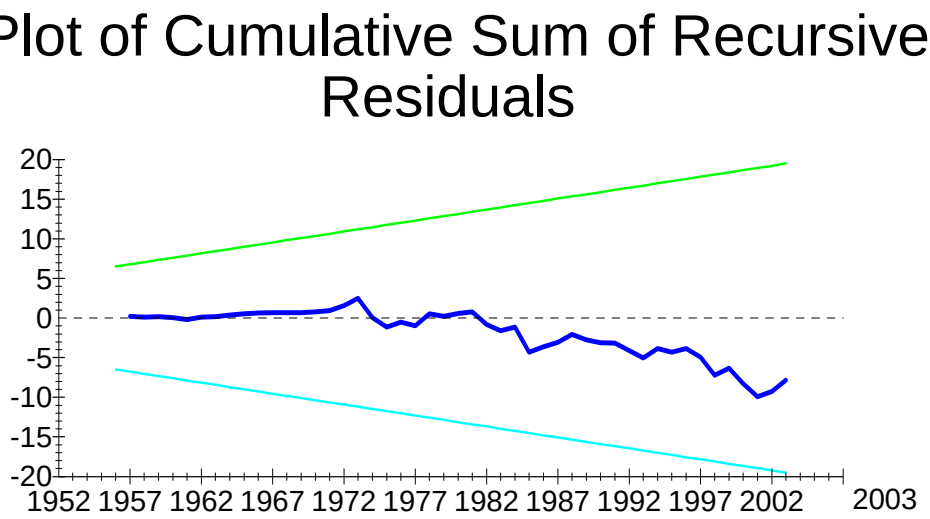
Regarding the power of tests, the Perron and Vogelsang (1992) model is robust. The testing power of Perron (1997) and Zivot and Andrews (1992) models are almost the same. On the other hand, the Perron (1997) model is more comprehensive than Zivot and Andrews (1992) model as the former includes both $t$ and $DT_b$ while the latter includes $t$ only.

## Appendix 2

Table A2.1 Autoregressive Distributed Lag Estimates for equation (2)
ARDL (2, 0) selected based on Akaike Information Criterion

| DEPENDENT VARIABLE IS $LnR_t$ | | | | |
|---|---|---|---|---|
| Regressors | Coefficient | Standard Error | T-Ratio | Probability |
| $LnR_{t-1}$ | 1.224 | 0.141 | 8.693 | 0.000 |
| $LnR_{t-2}$ | -0.422 | 0.148 | -2.848 | 0.007 |
| $LnPR_t$ | 1.107 | 0.617 | 1.793 | 0.079 |
| *Intercept* | 0.093 | 0.057 | 1.640 | 0.108 |
| R-Squared | 0.859 | R-Bar-Squared | 0.850 | |
| S.E. of Regression | 0.057 | F-stat. F(3, 47) | 95.558 [0.000] | |
| Mean of Dependent Variable | -0.056 | S.D. of Dependent Variable | 0.147 | |
| Residual Sum of Squares | 0.152 | Equation Log-likelihood | 75.996 | |
| Akaike Info. Criterion | 71.996 | Schwarz Bayesian Criterion | 68.133 | |
| DW-statistic | 2.010 | | | |
| **Diagnostic Tests** | | | | |
| **Test Statistics** | **LM Version** | | **F Version** | |
| A: Serial Correlation | CHSQ (1) = 0.125 [0.723] | | F(1, 46) = 0.113 [0.738] | |
| B: Functional Form | CHSQ (1) = 0.867 [0.768] | | F(1, 46) = 0.078 [0.781] | |
| C: Normality | CHSQ (2) = 12.736 [0.300] | | Not applicable | |
| D: Heteroscedasticity | CHSQ (1) = 1.347 [0.246] | | F(1, 49) = 0.170 [0.682] | |

A: Lagrange multiplier test of residual serial correlation
B: Ramsey's RESET test using the square of the fitted values
C: Based on a test of skewness and kurtosis of residuals
D: Based on the regression of squared residuals on squared fitted values




The straight lines represent critical bounds at 5% significance level