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Abstract

This paper tests empirically the Balassa-Samuelson (BS) hypothesis using annual data for Australia. We applied the ARDL cointegration technique developed by Pesaran et al. (2001) and found evidence of a significant long-run relationship between real exchange rate and Australia-US productivity differential during the period of 1950-2003. We found 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. We suspect that the elasticity coefficient is "over-estimated" due to the exclusion of relevant explanatory variables. The dynamics and the determinants of the real exchange rate movements are numerous; they include terms of trade, interest rate differentials, net foreign liabilities among others along with labour productivity differential.

Keywords: Real Exchange Rate, Balassa-Samuelson hypothesis, Unit-root, Structural break and ARDL.

JEL Classification: C22, F11, F31.

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Introduction

The Balassa-Samuelson (B-S) effect[1] model was surprisingly 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[2] 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 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 $e = \frac{P}{P^*}$, where e is the amount of domestic currency per unit of foreign currency, 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 changes can be attributable to productivity differential in the two economies. Rapid economic growth is accompanied by real exchange rate appreciation because of differential productivity growth between traded (T) and nontraded (NT) sectors. Since the differences in

[1] Earlier, outlines of the explanation of the effect were provided by Harrod (1933) and Ricardo.
[2] Bhagwati (1984) and Kravis and Lipsey (1983) provide an alternative theory to explain lower price levels in poorer countries.

productivity increases are expected to be larger in high growth countries, the B-S prediction should be more visible among fast growing economies[3].

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 II. In section III we test the time series properties of the variables in the presence of endogenous structural break in data. This is done since the traditional unit root tests suffer from power deficiency when structural break is present in the data. In section IV we estimate the model by using the Auto Regressive Distributed Lag modelling

[3] Post war Japan is a classic example of the B-S effect.

approach which allows us to estimate the model regardless whether the variables are I(0) or I(1).

II 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 production functions of the goods in both countries can be specified as follows:

$$Y_T = f(L_T) \qquad Y_{NT} = g(L_{NT}) \qquad (1)$$

$$Y_T^* = f^*(L_T^*) \qquad Y_{NT}^* = g^*(L_{NT}^*) \qquad (2)$$

The Balassa-Samuelson model assumes that the labour market is competitive and labour is perfectly mobile within each country but not between countries. As a consequence, the nominal wage is equal in both sectors for each country as in equations (3) and (4).

$$P_T f'(L_T) = w = P_{NT} f'(L_{NT}) \qquad (3)$$

$$P_T^* f'^*(L_T^*) = w^* = P_{NT}^* f'^*(L_{NT}^*) \qquad (4)$$

where, the prime after a function denotes marginal product of labour. The second assumption of the Balassa-Samuelson model is that the Purchasing Power Parity (PPP) holds for tradable goods only which implies:

$$P_T = eP_T^* \qquad (5)$$

where e denotes the bilateral nominal exchange rate.

The price levels in the two countries are defined as weighted geometric averages of prices in both sectors with weights j and $(1-j)$ reflecting the shares of these goods in the consumption basket:

$$P = P_T^{1-i} P_{NT}^{i} \qquad (6)$$

$$P^* = P_T^{*1-j} P_{NT}^{*j} \qquad (7)$$

To simplify matters we can make the price of tradable goods equal to unity in both countries, i.e., $P_T = P_T^* = 1$. Hence, equation (5) implies that the nominal exchange rate is also equal to unity. The price equations (6) and (7) can thus be re-written as follows:

$$P = P_{NT}^{i} \qquad (6a)$$

$$P^* = P_{NT}^{*j} \qquad (7a)$$

Similarly, from equations (3) and (4) we have:

$$P_{NT} = \frac{f'(L_T)}{g'(L_{NT})} \qquad (3a)$$

$$P_{NT}^* = \frac{f'^*(L_T)}{g'^*(L_{NT})} \qquad (4a)$$

The real exchange rate is defined as:

$$\rho = \frac{P}{eP^*} = \frac{P}{P^*} \qquad (8)$$

Substituting equations (3a) and (4a) into equations (6a) and (7a) and them into equation (8) yields:

$$\rho = \frac{P}{P*} = \frac{\{f'(L_T)/g'(L_{NT})\}^{i}}{\{f'^*(L_T^*)/g'^*(L_{NT}^T)\}^{j}} \qquad (9)$$

Equation (9) expresses the Balassa-Samuelson (BS) effect. It asserts that if the traded goods marginal productivity relative to non-traded goods marginal productivity is increasing faster in 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 right hand side of equation (9) can be rewritten in terms of the average productivity of labour such as.

$$\rho = \frac{\{(Y_T / L_T)^i\}}{\{(Y_T^* / L_T^*)\}} \qquad (10)$$

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 real exchange rate appreciation.

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

$$LnR_t = \alpha + \beta LnPR_t \qquad (11)$$

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.

III Tests for Time Series Properties in the Presence of Structural Break

In this study, we used annual data for all series from 1950 to the last available data until 2003 from Heston *et al.* (2006) Penn World Table Version 6.2. For PR, we used the real GDP per worker (in 2000 international prices) of each country treating the US as the reference country.

Equation (11) can be analysed by 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 we are going to use in this paper. A succinct review is given in Appendix 1.

To ascertain the order of integration, we applied the traditional Augmented Dickey-Fuller (ADF) and Phillips-Perron (PP) unit root test. 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 structural break[4], we applied the most comprehensive models of Perron (1997) along with Zivot and Andrews (1992) model. 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 Shrestha-Chowdhury (2005) procedure. In Table 1, the unit root hypotheses are rejected at the 5 per cent level of significance for *LnR* by all the tests except IO2 model. The estimated break date corresponds to 1979 for *LnR* and 1985 for *LnPR*. The endogenously determined break dates are plausible with the events occurring in the Australian economy. After a

[4]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 devaluations, deregulation of both real and financial sectors and policy regime shifts.

sustained period of appreciation, depreciations of the real exchange rate occurred during 1974-1978 which had an impact on the Australian economy 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	S	***IO1***	*1*	*1985*	*-6.18*	*NS*
IO2	8	1971	-6.20	NS	**IO2**	1	1984	-6.30	NS
AO	*8*	*1979*	*-4.15*	*S*	**AO**	1	1951	-3.32	S
Zivot-Andrews	1	1977	-4.27	S	**Zivot-Andrews**	2	1976	-3.01	S

Note: 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.

IV 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 the Johansen-Juselius (1990) require that all the variables be of equal degree of integration, i.e., I(1). Therefore, these methods of cointegration are not appropriate and cannot be employed. Hence, we adopt the ARDL modelling approach for cointegration analysis in this study.

The main advantage of ARDL modelling lies in its flexibility that it can be applied when the variables are of different order of integration (Pesaran and Pesaran 1997). 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 and Chai 2003). Moreover, a dynamic error correction model (ECM) can be derived from ARDL through a simple linear transformation (Banerjee *et al.* 1993). The ECM integrates the short run dynamics with the long run equilibrium without losing long run information. It is also argued that using the ARDL approach avoids problems resulting from nonstationary time series data (Laurenceson and Chai 2003).

Thus, the error correction specification of the ARDL model pertaining to equation (11) is given in equation (12) and can be expressed as:

$$\Delta LnR_t = \alpha_0 + \delta_1 LnR_{t-1} + \delta_2 PR_{t-1} + \sum_{i=1}^{p} b_i \Delta LnR_{t-i} + \sum_{i=0}^{q} c_i \Delta LnPR_{t-i} + \varepsilon_t \qquad (12)$$

The parameter δ_i, i =1, 2, are the long run multipliers. The parameters b_i, c_i, are the short run multipliers. ε_t represents residuals.

To select the appropriate model in equation (12), several specifications with different lags were tested for statistical significance and for consistency with the cointegration method. The specification we used here is the restricted intercept with no trend (Case III in Pesaran *et al.*, 2001:296). We have estimated the model given in equation (12) and found the optimal model to be $\text{ARDL}[2,0]$ based on the AIC and SBC model selection criteria[5]. The estimated ARDL model is given in Appendix 2, Table A2.1.

Estimation of Long Run Coefficients

We investigated the long run relationship between the Australian real exchange rate (R_t) and the Australia-US labour productivity differential (PR_t) by the using the

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

‘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 (12). 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 mutually cointegrated categories[6]; and these critical values are generated for sample sizes of 500 and 1000 observations with replications of 20,000 and 40,000 respectively.

Based on the ‘bounds test’ (given in Table 3), the computed *F*-statistic is 2.79, is below the lower critical bound (*LCB)* at the 10 per cent significance level. Hence,

Table 3: Bounds Test for Cointegration

Computed *F*-Statistics ($F_{2,44}$)	2.79	
Critical Bounds (10 per cent)	*LCB*: 4.04	*UCB*: 4.78
Critical Bounds (5 per cent)	*LCB*: 4.94	*UCB*: 5.73

Note: Critical Bounds are from Pesaran *et al.* (2001:300) Table CI (iii) Case III.

seemingly, there does not appear to be a long-run relationship between the real exchange rate movement and the productivity differential. 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

[6] 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.* (2001:299).

are calculated for large number of observations (500 and 1000) which may be problematic in small samples as in our case. Therefore, following Kremers *et al.* (1992) we shall rely upon the significance of the error correction term as a useful and efficient way of establishing cointegration.

We tested the robustness of the $F-$test of Pesaran *et al*. (2001) by using the dynamic ordinary least squares (DOLS) estimator proposed by Stock and Watson (1993). The DOLS estimator is based on the modified version of equation (11) that includes past, present and future values of the change in $LnPR_t$:

$$LnR_t = \beta_0 + \theta LnPR_t + \sum_{j=-p}^{p} \delta_j \Delta_{t-j} + \upsilon_t \tag{11a}$$

The DOLS estimator of θ is the OLS estimator in equation (11a), and is efficient in large samples if LnR_t and $LnPR_t$ are cointegrated. Further, the t-statistic constructed using the DOLS estimator with heteroskedastic-and autocorrelation-consistent (HAC) standard errors has a standard normal distribution in large samples. We estimated θ=0.73 with $t = 3.74$ implying there is cointegration between LnR_t and $LnPR_t$.

The estimated long-run coefficients for the ARDL model are given in Table 4. 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 4: Estimated Long Run Coefficients for Equation 12: ARDL (2, 0)
Dependent Variable: LnR_t

Variables	Coefficient	t-ratio	P-value
LnPR	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 (11)[7]. The crux of the B-S hypothesis is premised on the proposition that productivity differential alone is the determinant of the real exchange of a country. However, in recent times, researchers are 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 *et al.* (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

[7] 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 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 (1985:443-46).

rates. In contrast, Bagchi *et al.* (2004:84) find "…the terms of trade prove quantitatively more important in explaining the long-run real exchange rate than the real interest rate differential.[8]" 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 of which Australia is one of them, Bahmani-Oskoee and Nasir (2004) found the productivity differential coefficient to be 0.97 per cent compared to our value of 5.58 per cent. We find this value 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 we are not sure why the results are so vastly different[9]!

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 pass 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 5. The short run adjustment process is measured by the error

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[8] 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 0.04 per cent in real terms.

[9] 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.

correction term (*ECM*). The ECM indicates how quickly variables adjust and return to equilibrium and the coefficient of ECM should carry the negative sign and be statistically significant. As shown in Table 5, 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 current account deficit and the relevant macroeconomic data.

Table 5: Error Correction for the Selected ARDL Model: ARDL (2, 0)
Dependent Variable: ΔLnR_t

Variables	Coefficient	P-value
ΔLnR_{t-1}	0.752*	0.000
$\Delta LnPR_t$	0.134**	0.030
$\Delta Intercept$	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.

V. 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 big void on this topic in Australia but also adds to the limited number of time series study on this subject. We tested the time series properties of the variables in the presence of structural break 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 bounds test of Pesaran *et al.* (2001) failed to reveal any long term association between changes in real exchange rate and productivity differential. However, the DOLS test of Stock and Watson (2003) along with the sign test of the ECM term was sufficient to demonstrate the existence of a long term cointegration between the variables.

We have derived the B-S model based on some simplistic assumptions[10] (single factor of production, constant returns to scale, constancy of terms of trade thus ignoring the demand side of the economy). Our empirical results 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. We 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. We believe 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 terms of trade, interest rate differentials, net foreign liabilities among others.

[10] 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.

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) have low power in the presence of 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 known exogenous structural break is included[11]. 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) $y_t = \mu + dD(TB)_t + y_{t-1} + e_t$ (A 1)

Model (B) $y_t = \mu_1 + y_{t-1} + (\mu_2 - \mu_1)DU_t + e_t$ (A 2)

Model (C) $y_t = \mu_1 + y_{t-1} + dD(TB)_t + (\mu_2 - \mu_1)DU_t + e_t$ (A 3)

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) $y_t = \mu_1 + \beta t + (\mu_2 - \mu_1)DU_t + e_t$ (A 4)

Model (B) $y_t = \mu + \beta_1 t + (\beta_2 - \beta_1)DT_t^* + e_t$ (A 5)

Model (C) $y_t = \mu_1 + \beta_1 t + (\mu_2 - \mu_1)DU_t + (\beta_2 - \beta_1)DT_t + e_t$ (A 6)

[11] 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.

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 known break date which raises the problem of pre-testing and data mining regarding the choice of the break date (Maddala and 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 we are going to discuss 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 we review these models briefly and detailed discussions are found in the cited works. Zivot and Andrews (1992) models are as follows:

Model with Intercept

$$y_t = \hat{\mu}^A + \hat{\theta}^A DU_t(\hat{\lambda}) + \hat{\beta}^A t + \hat{\alpha}^A y_{t-1} + \sum_{j=1}^{k} \hat{c}_j^A \Delta y_{t-j} + \hat{e}_t \qquad \text{(A 7)}$$

Model with Trend

$$y_t = \hat{\mu}^B + \hat{\beta}^B t + \hat{\gamma}^B DT_t^*(\hat{\lambda}) + \hat{\alpha}^B y_{t-1} + \sum_{j=i}^{k} \hat{c}_j^B \Delta y_{t-j} + \hat{e}_t \qquad \text{(A 8)}$$

Model with Both Intercept and Trend

$$y_t = \hat{\mu}^C + \hat{\theta}^C DU_t(\hat{\lambda}) + \hat{\beta}^C t + \hat{\gamma}^C DT_t^*(\hat{\lambda}) + \hat{\alpha}^C y_{t-1} + \sum_{j=1}^{k} \hat{c}_j^C \Delta y_{t-j} + \hat{e}_t \qquad \text{(A 9)}$$

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)

$$y_t = \mu + \delta DU_t + \theta D(T_b)_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t \tag{A 10}$$

Additive Outlier Model (AOM) – Two Steps

$$y_t = \mu + \delta DU_t + \tilde{y}_t \tag{A 11}$$

and

$$\tilde{y}_t = \sum_{i=0}^{k} w_i D(T_b)_{t-i} + \alpha \tilde{y}_{t-1} + \sum_{i=1}^{k} c_i \Delta \tilde{y}_{t-i} + e_t \tag{A 12}$$

$\tilde{y}$ 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):

$$y_t = \mu + \theta DU_t + \beta t + \delta D(T_b)_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t \tag{A 13}$$

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

$$y_t = \mu + \theta DU_t + \beta t + \gamma DT_t + \delta D(T_b)_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + e_t \tag{A 14}$$

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

$$y_t = \mu + \beta t + \delta DT_t^* + \tilde{y}_t \tag{A 15}$$

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

$$\widetilde{y}_t = \alpha\widetilde{y}_{t-1}\sum_{i=1}^{k} c_i \Delta\widetilde{y}_{t-i} + e_t \qquad \text{(A 16)}$$

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 (Lumsdaine and Papell 1997; Bai and Perron 2003) which we are not going to discuss here.

Regarding the power of tests, the Perron and Vogelsang (1992) model is robust. The testing power of Perron (1997) models and Zivot and Andrews models (1992) are almost the same. On the other hand, 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

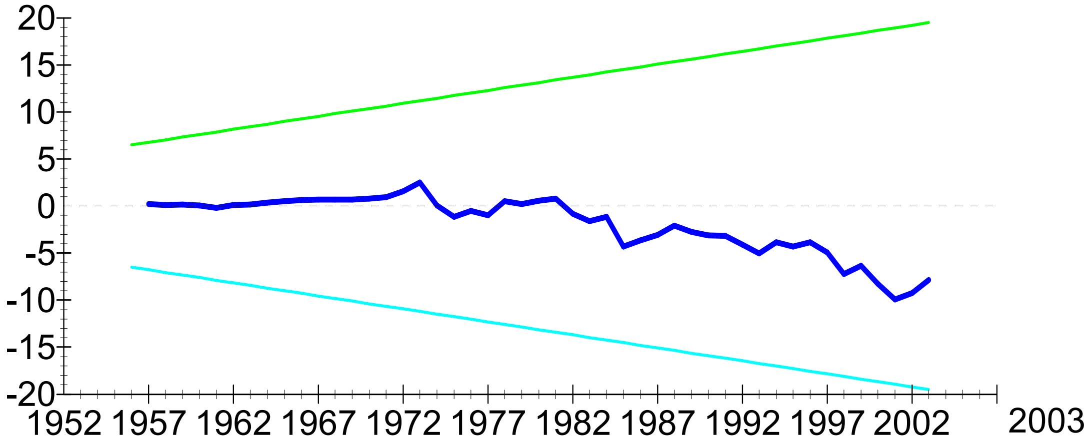
Table A2.1 Autoregressive Distributed Lag Estimates for equation (12) ARDL (2, 0) selected based on Akaike Information Criterion

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-1}	-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

Plot of Cumulative Sum of Recursive Residuals



The straight lines represent critical bounds at 5% significance level

REFERENCES

Asea, P.K. and E.G. Mendoza (1994) "The Balassa-Samuelson Model: A General Equilibrium Appraisal", *Review of International Economics* 2, 244-267.

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

Bagchi, B., G.E. Chortareas and S.M. Miller (2004) The Real Exchange Rate in Small, Open Developed Economies: Evidence from Cointegration Analysis, *Economic Record* 80 (248), 76-88.

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

Balvers, R.J. and J.H. Bergstrand (1997) "Equilibrium Real Exchange Rates: Closed-Form Theoretical Solutions and Some Empirical Evidence", *Journal of International Money and Finance* 16, 345-366.

Banerjee, A.; J. Dolado and R. Mestre (1998) "Error-Correction Mechanism Tests For Cointegration in Single-Equation Framework", *Journal of Time Series Analysis* 19,267-283.

Banerjee, A., J. Dolado, , J. Galbraith, and D. Hendry (1993) *Cointegration, Error Correction, and the Econometric Analysis of Nonstationary Data* (Oxford: Oxford University Press).

Bergstrand, J.H. (1991) "Structural Determinants Of Real Exchange Rates And National Price Levels: Some Empirical Evidence", *American Economic Review* 81, 325-334.

Bhagwati, J.N. (1984) "Why Are Services Cheaper in The Poor Countries?", *Economic Journal* 94, 279-286.

Blundell-Wignall, A., J. Fahrer and A. Heath (1993) "*Major Influences on the Australian Dollar Exchange Rate"*, in A. Blundell-Wignall (ed), The Exchange Rate, International Trade and the Balance of Payments, Proceedings of a Conference, Reserve Bank of Australia: 30-78.

Canzoneri, M.B., R.E. Cumby, and B. Diba (1999) "Relative Labor Productivity And The Real Exchange Rate in The Long Run: Evidence For A Panel Of OECD Countries", *Journal of International Economics* 47, 245-266.

Charemza, W. and D. Deadman, , 1992, *New Directions in Econometric Practice: General to Specific Modelling, Cointegration and Vector Autoregression* (Aldershot: Edward Elgar).

De Gregorio, J., A. Giovannini and T.H. Krueger (1994) "The Behavior Of Nontradable-Goods Prices in Europe: Evidence and Interpretation", *Review of International Economics* 2, 284-305.

De Gregorio, J. and H.C. Wolf (1994) "Terms of Trade, Productivity, and The Real Exchange Rate", NBER working paper No.4807.

Gruen, D.W.R. and J. Wilkinson, (1994), Australia's Real Exchange Rate – Is It Explained by the Terms of Trade or by Real Interest Differentials?. *Economic Record* **70**, 204–19.

Harrod, R.F. (1933) *International Economics* (Cambridge University Press: Cambridge)

Heston, A., D.A. Nuxoll and R. Summers (1994) "The Differential Productivity Hypothesis and Purchasing Power Parties: Some New Evidence", *Review of International Economics* 2, 227-243.

Johansen, S., (1991) "Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models", *Econometrica* **59**, 1551-1580.

Johansen, S., (1995) *Likelihood Based Inference in Cointegrated Vector Autoregressive Models* (Oxford University Press: Oxford).

Johansen, S. and Juselius, K., (1990) Maximum Likelihood Estimation and Inference on Cointegration- with Applications to the Demand for Money, *Oxford Bulletin of Economics and Statistics* **52**, 169-210.

Judge, G. G., Griffiths, W. E. Hill, H. Lutkepohl, R. C. and Lee T. C. (1985) *The Theory and Practice of Econometrics*, Second edition, (John Wiley and Sons, New York).

Kmenta, J. (1985) *Elements of Econometrics*, Second Edition, (New York: Macmillan).

Kravis, I.B. and R.E. Lipsey (1988) "National Price Levels And The Prices Of Tradables And Nontradables", *American Economic Review* 78, 474-478.

Laurenceson, J. and J. Chai (2003) *Financial Reform and Economic Development in China* (Edward Elgar: Cheltenham).

Lothian, J.R and M.P. Taylor (1996) "Real Exchange Rate Behavior: The Recent Float From The Perspective of The Last Two Centuries", *Journal of Political Economy* 104, 488-509.

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

Maddala, G.S. and Kim, I.M., (2003) *Unit Roots, Cointegration, and Structural Change* (Cambridge University Press: Cambridge).

Nelson, C. and C. Plosser (1982) Trends and Random Walks in Macroeconomic Time Series: Some Evidence and Implications, *Journal of Monetary Economics* 10, 139-162.

Obstfeld, M. and K. Rogoff (1996) Foundations of International Macroeconomics, (MIT Press: Cambridge).

Perron, P. and T.J. Vogelsang (1992) Nonstationary and Level Shifts with an Application to Purchasing Power Parity, *Journal of Business and Economic Statistics* 10, 301-320.

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

Perron, P., 1997, Further Evidence on Breaking Trend Functions in Macroeconomic Variables, *Journal of Econometrics* 80, 355-385.

Pesaran, M.H. and B. Pesaran 1997, *Working with Microfit 4.0: Interactive Econometric Analysis* (Oxford University Press: Oxford).

Pesaran, M.H. and R.P. Smith (1998) "Structural Analysis of Cointegrating VARs", *Journal of Economic Survey* 12, 471-505.

Pesaran, M.H. and Y. Shin, (1999) "An Autoregressive Distributed Lag Modelling Approach to Cointegration Analysis", in: S. Strom, A. Holly and P. Diamond (eds) *Econometrics and Economic Theory in the 20th Century: The Ragner Frisch Centennial Symposium* (Cambridge University Press: Cambridge).

Pesaran, M.H., Y. Shin and R.J. Smith (2001) "Bounds Testing Approaches to the Analysis of Level Relationships", *Journal of Applied Econometrics* 16, 289-326.

Phillips, P.C.B. and P. Perron, (1988) "Testing for a Unit Root in Time Series Regression", *Biometrica* 75, 335-346.

Rogoff, K. (1996) The Purchasing Power Parity Puzzle, *Journal of Economic Literature* 34, 647-668.

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

Samuelson, P.A. (1994) "Facets of Balassa-Samuelson Thirty Years Later", *Review of*

International Economics 2, 201-226.

Shrestha, M.B. and K. Chowdhury, (2005) "A Sequential Procedure for Testing Unit Roots in the Presence of Structural Break in Time Series Data: An Application to Nepalese Quarterly Data 1970–2003", *International Journal of Applied Econometrics and Quantitative Studies,* 2**,** 1-16.

Stock, J.H. and M.W. Watson (2003) Introduction to Econometrics, (Addison Wesley: New York).

Strauss, J. (1998) "Relative Price Determination In The Medium Run: The Influence Of Wages, Productivity, and International Prices", *Southern Economic Journal* 65, 223-244.

Tarditi, A. (1996) Modelling the Australian Exchange Rate, Long Bond Yield and Inflationary Expectations, Research Discussion Paper 9608, Economic Analysis Department, Reserve Bank of Australia.

Zivot E. and D.W. Andrews (1992), "Further Evidence on the Great Crash, the Oil-Price Shock, and the Unit-Root Hypotheses", *Journal of Business and Economic Statistics* 10: 251-270.