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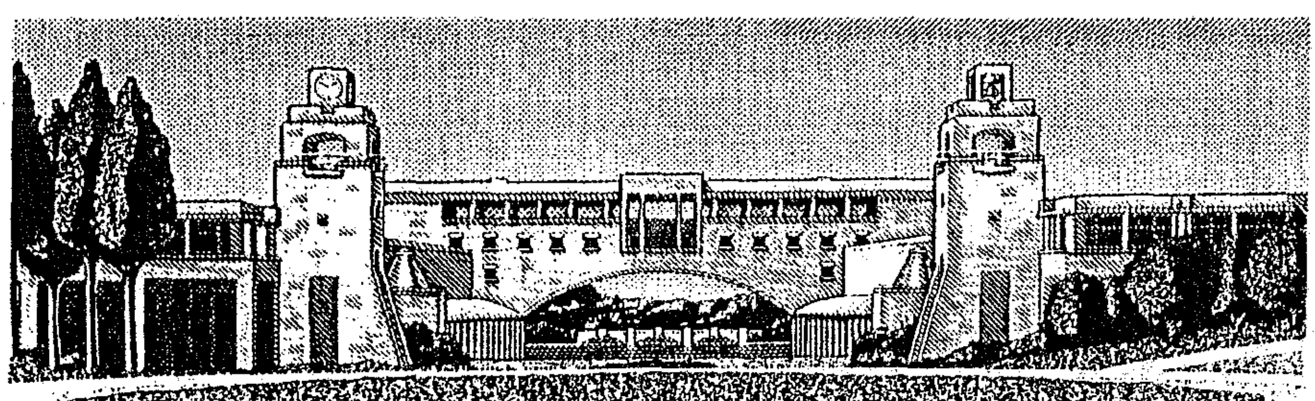


Positive Evidence for Purchasing Power Parity

Christopher M. Adam and Ronald Bewley

Abstract

The absolute deterministic version of purchasing power parity (PPP), in which nominal exchange rates move precisely to offset changes in aggregate price levels between national economies, is not generally supported by empirical evidence. Testing PPP with standard regression models permits some stochastic "slippage" in the relation, but still yields little support for PPP. We argue in this paper that such tests do not correctly appraise in a statistical fashion the strength of the relations among prices and exchange rates. Using a development of Box and Tiao's canonical correlation approach to cointegration, we find a significant stable relation among certain aggregate national price indices and the exchange rate between the US and Australia in a period of floating exchange rates. The sign of the relation is as advocated by PPP, although the size of the adjustment effect is lower. Furthermore, the relation breaks down at the times of known foreign exchange market intervention. The search procedure thus can be used to identify *ex post* the timing of switches in policy regimes.

Paper presented to the Australasian meeting of the Econometrics Society at Armidale, NSW, July 1989. This paper was written while Bewley was a visitor in the Department of Agricultural Economics, University of Sydney and he wishes to thank the department for his support.

The Appropriate Version of PPP to Test

As a theory of exchange rate determination, purchasing power parity (PI. insists on the dominance of trade in real goods and services. Monet. disturbances may temporarily cause exchange rates to deviate from PPP, once the disturbance is eliminated, trade in real goods and services ensu that the purchasing power of the different currencies of the trading countr will be the same in those countries. Real trend deviations among countries, however, will break PPP in the long-run. Hence any measures of PPP over t longer time horizon should allow for differing rates of productivity growth.

The "law of one price" states that the price of a single good "will be the same in all locations when quoted in the same currency" (Dornbusch, 1985 p.2), provided the trading takes place in an integrated, competitive mark If we construct price indices for each of two trading economies and t indices use identical weights, then the existence of the law of one price each traded good produces a law of one price for the aggregate price indices. We then obtain the strong or absolute form of PPP, in which the ratio of t price indices sets a single nominal exchange rate for the two currencies the trading countries. The real exchange rate, taking the ratio of the pri levels of the two countries expressed in a common currency is then constar at unity.

Empirical objections to this are manifold. Any cost associated with t trading in the markets for the goods, be it information costs, tariff transportation costs or other impediments to trade expressed as a cost prevent the law of one price from holding for every single good or service in the indices, and hence threaten to destroy the finding of PPP in comparisons of the relative price levels and the nominal exchange rate.

To overcome these objections, a weak form of PPP allows for a scale difference in the prices of the goods traded between the countries, so that the nominal exchange rate is some constant proportion of the price ratio. This version of PPP still contains an empirical hypothesis: the scale factor is constant over time, so that comparison of the price levels across the countries over time would yield the same multiple of the nominal exchange rate.

If a weak or relative version of PPP holds, then the proportionate change in the nominal exchange rate should equal the difference in the inflation rates (proportionate rates of change of the aggregate price indices) in each country. This measure still requires the researcher to use the same basket of commodities to effect the aggregation of price levels in each country.

Suppose, however, we take away even the use of identical baskets to create the indices. That is, we have now admitted to differences in prices due to transportation costs, and that we are no longer using the same index number formula to aggregate the prices for each economy. Could any version of PPP still hold except by accident? Provided the homogeneity postulate of monetary theory holds, we can yet derive a form of the PPP proposition. As Dornbusch (1985, p.5) notes, "the homogeneity asserts that a purely monetary disturbance, leaving unchanged all equilibrium prices, will lead to an equiproportionate change in money and all prices, including the price of foreign exchange. In this very special experiment PPP holds even if the law of one price does not apply."

The power of this insight is that even differences in price index baskets, or the absence of direct arbitrage pressure for the goods and services (they need not be traded goods only in the indices), can yet produce the finding that proportionate changes in the nominal exchange rate match the differences in inflation rates of the compared economies.

Why should this process hold in the absence of spatial arbitrage? We need to provide an adjustment process linking the monetary changes, prices, and the nominal exchange rate. Note, however, that in such an approach, we need not confine our discussion to a one-way causal link from prices to exchange rates. Indeed, a large section of the current literature on pricing structures in open economies highlights the existence of a causal link from exchange rate movements to prices within each economy (the so-called "pass-through" effect).

In fact we can derive models in which the prices and nominal exchange rates are simultaneously determined, with PPP obtaining as a long-run equilibrium condition but faces disruption in the short-run under the influence of unexpected disturbances to the markets of the trading economies. The models of Dornbusch (1976) and those writing in this style follow precisely this methodology.

We approach the PPP hypothesis from an empirical stance in the first instance. If relevant time series for aggregate price indices in different countries and the appropriate nominal exchange rate move together in a statistically significant way in the "correct" direction for extended periods of time, we can claim that the evidence for some weak PPP relation is well grounded. Without that relationship, we are unable to establish even the grounds for a long-run equilibrium version of the PPP model which allows for measurement error and transportation costs.

Our method uses the recently developed technique of cointegration of time series. In essence, the method permits us to assert that while two or more time series may be individually nonstationary, some linear combination of them might be stationary. We can interpret PPP as a linear combination with particular parameter values. All linear combinations of the time series, except the cointegrating ones, have asymptotically infinite variances, and we can test for this in our modelling of the PPP relation. Non-cointegrated time series tend to drift apart without bound.

This paper is not the first to use cointegration methods for studying afresh the PPP relationship. Work by Baillie and Selover (1987), Johnson (1987), and Taylor (1988) have adapted the cointegration approach to evaluating evidence for PPP. It is interesting to note even this short selection of papers that the results are equivocal. Taylor conducts tests across the nominal exchange rates and relative manufacturing prices against the US for the UK, West Germany, France, Canada, and Japan for the most recent period of floating exchange rates. He finds no strong evidence in favour of PPP in his conclusions.

For the same five countries against the US, Baillie and Selover find cointegration for France, but not for the other nations. Johnson, however, does find good support for long-run equilibrium relationships between Canadian prices, American prices and the Canada-US dollar exchange rate. His findings cover both fixed and floating rate periods.

The Test Method

One problem that has hampered previous research in this area is that, until very recently, econometric methods have only been able to accommodate a single cointegrating relationship and, typically, this equation has been estimated by OLS with the usual normalisation problem that exists whenever the direction of causality is not clear. Bewley, Fisher and Parry (1988) developed the Box-Tiao (1977) approach to modelling near-nonstationary time series to a multi-equation cointegrating representation using an error correction framework.[1] The essence of the Box-Tiao approach is to reparameterize the VAR (Vector Autoregressive) model

$$A(L)y_t = u_t . \tag{1}$$

where $A(L) = I - A_1L \ldots - A_pL^p$, a p'th degree $n \times n$ matrix polynomial in the lag operator L, and u_t is an $n \times 1$ vector of independently and identically distributed multivariate normal disturbances.

On defining

$$\Gamma^{-1}\Xi M = M\Lambda \tag{2}$$

where Γ is the variance-covariance matrix of y_t, Ξ is the variance-covariance matrix of $\hat{y}_t$ obtained from (1), M is the matrix of eigenvectors and Λ the eigenvalues, the original VAR model (1) can be transformed to an equivalent

[1] In this paper we make no distinction between a time series with a root less than by close to unity and a unit root. Any difference is likely to be minimal in an empirical sense with finite time series.

form

$$B(L)Y_t = v_t \tag{3}$$

where $B(L) = M'A(L)$, $Y_t = M'y_t$, and $v_t = M'u_t$. Not only is (3) a system with orthogonal disturbances and orthogonal time series, Y_t, but the canona have a cointegration interpretation. Time series in Y_t with roots near unity are likely to be nonstationary and should be first-differenced. On the other hand, stationary Y_t series are stationary linear combinations of y_t or, in the Engle and Granger (1987) sense, cointegrating relationships. If any of the latter group are indeed white noise, these linear combinations correspond to standard static regression equations.

There are three basic ways of distinguishing between stationary and non-stationary series in Y_t. Johansen (1988) performs a direct test on the elements of Λ while Bewley, Fisher and Parry (1988) and Bewley and Elliott (1989a, 1989b) favour testing each series for the presence of a unit root in the Dickey-Fuller (1979) tradition or by using the error correction version of (3) developed in Bewley, Fisher and Parry

$$C(L)\Delta y_t = \Psi Y^*_{t-1} + w_t \tag{4}$$

where C(L) is of order p-1 and Y^*_t includes only stationary canona. Following Engle and Granger's single equation version of testing for cointegration, both ADF (augmented Dickey-Fuller) tests on Y^*_t, and tests on the joint significance of the elements in each column of Ψ in (4) were conducted. These system-wide tests are facilitated by noting that The series in Y^*_t are orthogonal.

Figure 3: Scatter Plot
(Data in Logarithms)

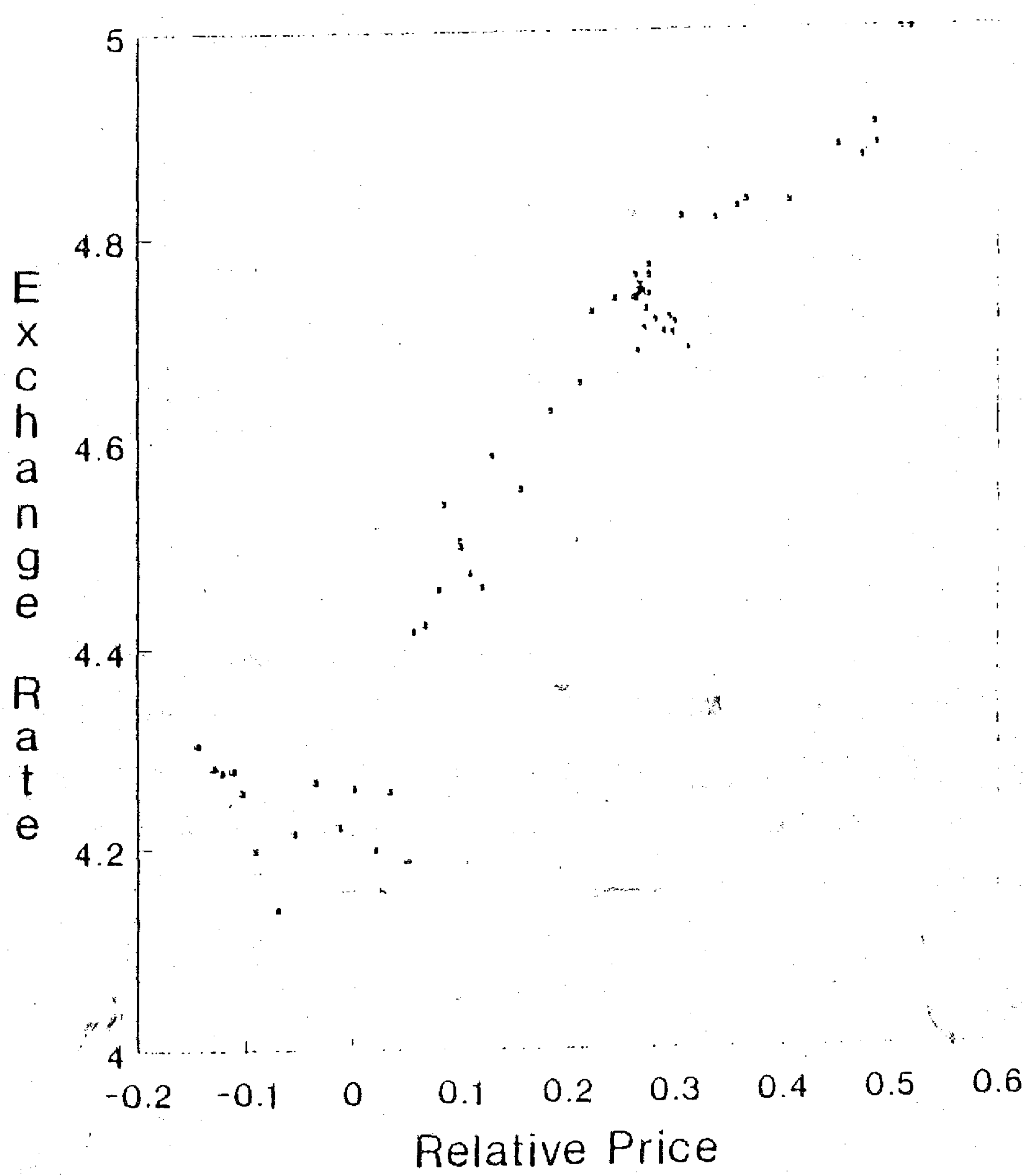


Figure 2: Deviations From PPP
Ratio of Exchange Rate to Long-Run Level

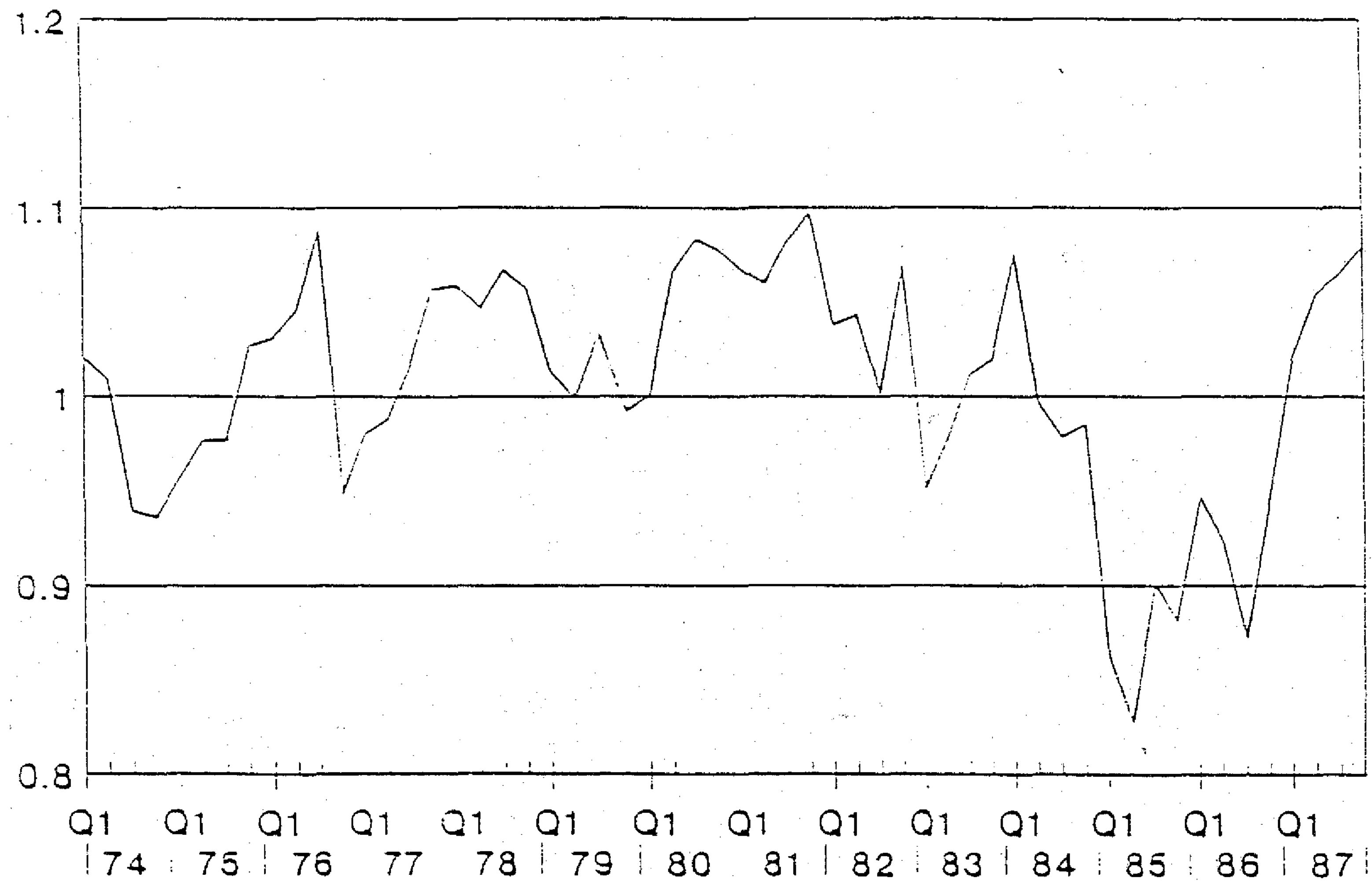
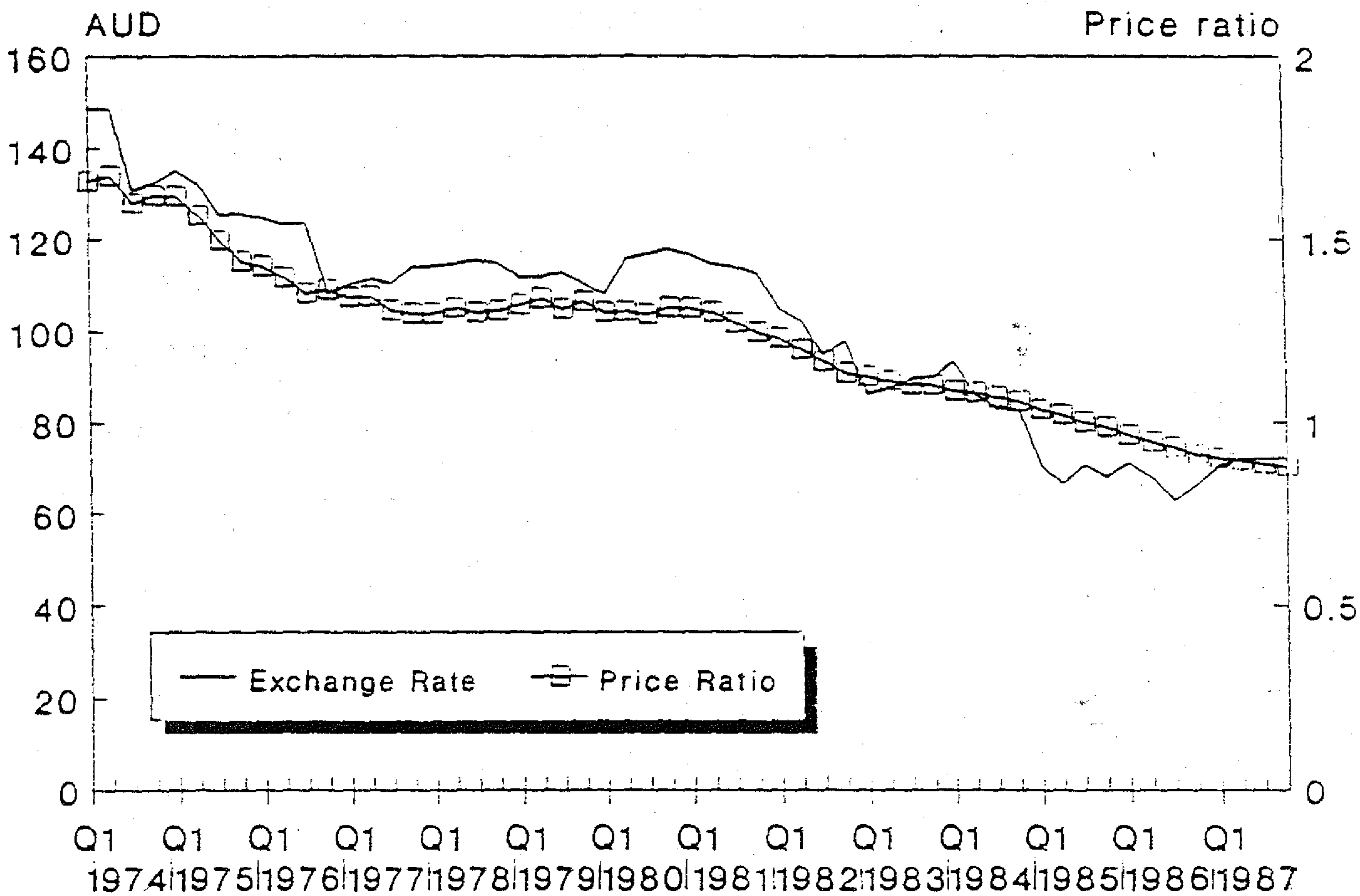


Figure 1: Purchasing Power Parity
Price Ratio and Exchange Rate



Empirical Findings

Before any tests for the existence of a long-run PPP relationship can be performed, suitable price indices must be defined. Using a small country argument and in the absence of reliable traded goods prices, we have taken the position that the Australian currency denominated in US dollars [AUD] might react to Australian domestic prices [Aus^P] and US export prices [US^e]. The exchange rate and price ratio are presented in Figure 1.

Using quarterly data from 1974.1 to 1988.1, ADF tests were performed on the level of each price and exchange rate, and the first-differences; all variables are expressed in logarithmic form. From Table 1 it can be noted that each ADF statistic on the price variables (1 - 3) with two lags are all less than the 95% critical value of -2.93 implying that nonstationarity cannot be rejected. Using similar tests on the changes in those prices, the presence of a second unit root cannot be rejected for the US index.

Table 1: ADF Tests on Price Variables

Variable	y	Δy
1. Exchange Rate	-0.79	-5.11
2. Aus. GDP Deflator	-2.49	-5.02
3. US Export Deflator	-1.85	-2.52
4. Ratio (3) to (2)	-1.41	-3.99
5. Ratio (1) to (4)	-1.86	-5.75

From the tests on the log of the ratio of the price deflators, also presented Table 1, it can be concluded that there is only one unit root. If it is accepted that the US index is I(2) and the Australian index is I(1) using Engle and Granger's terminology, the conclusion that the difference is I(1) is logically inconsistent. Given the relative magnitudes of the ADF statistics on the Δy series, it is reasonable to conclude that the US export price index is I(1) and that the anomaly is due to a power problem.

The final row in Table 1 shows that the real exchange rate, defined as the ratio of the nominal exchange rate to the relative price index involving US exports and Australian GDP is I(1). Thus, a simple test for absolute PPP would be rejected on this basis.

In order to further motivate the econometric approach taken here, let the y_t vector be defined by the three variables in Table 1 and given in rows 1 to 5. A second order VAR was estimated, that is equation (1) with p = 2, and the canonical decomposition was performed. The rows of M', Λ and ADF tests on the canona are presented in Table 2.

Table 2: Canonical Decomposition of 3-Equation VAR

Canona	Weights				
	AUD	US^e	Aus^p	λ	ADF
1	.4340	-.6632	.6097	.501	-2.53
2	.0493	.8886	-.4561	.975	-0.07
3	.0017	.7012	.7129	.999	-1.15

It follows from Table 2 that, on the basis of ADF tests alone, there are no stationary canona or cointegrating relationships. However, the root associated with the first canona is only 0.501 and the test of exclusion of canona 1 from the error correction model, proposed in Bewley and Elliott (1989b), is 3.16 which can be compared to an approximate F-distribution critical value of 2.79 at the 5% level with 3 and 52 degrees of freedom. On that basis we conclude that there is one cointegrating relationship with coefficients given by the first row in Table 2. Since the coefficients on the prices are numerically close to being equal and opposite, support is given to the notion that homogeneity of degree zero in prices is reasonable.

In the next stage of the analysis, the homogeneity restriction is imposed and the canonical decomposition of the 2-equation VAR is presented in Table 3. Again there is a root near unity and one near 0.5. The ADF test does not support the notion of cointegration but the exclusion test of the error correction term does at the 5% level with an observed F-ratio of 3.87 compared to a critical value of 3.20.

Table 3: Canonical Decomposition of 3-Equation VAR

Canona	Weights			
	AUD	US^e/Aus^p	λ	ADF
1	-.6153	.7883	.517	-2.55
2	.0535	.9986	.994	-1.47

Interestingly, the C_i matrix in (4) had no significant elements in the 2-equation model suggesting an over-parameterization. When the lag length was reduced by one, the exclusion test statistic increased to 5.05 which is close to the 1% critical value. A test of p = 2 against p = 1 is cle[illegible] rejected in the 3-equation model using an adjusted LR statistic [Sims, 1980] of 54.57 compared with an approximate X^2 critical value of 16.92. The main contribution to this rejection arises because of the delayed effect of US export prices on Australian domestic prices, over and above any relationship through the cointegrating canona.

On reworking the analysis with only one lag, the new cointegrating relationship can be expressed as

$$-0.6077\ ln[AUD] + 0.7942\ ln[US^e/Aus^P] = -2.65,$$

or

$$ln[AUD] = 4.36 + 1.22\ ln[US^e/Aus^P]$$

Since the real exchange rate analysed in Table 1 is not stationary, we reject the hypothesis that the coefficient on relative prices is unity. The deviations from PPP are displayed in Figure 2. and it can be noted that the MX missile crisis effect in early 1985 which was reinforced by the commodity price collapse in May 1986 caused the exchange rate to overshoot. More recently, the boom in commodity prices, together with an optimistic view of Australia's current account deficit problem caused an overshooting, in the exchange rate in the opposite direction. If Australia's exchange rate is to overact in the post-float period, some acknowledgement of the accompanying heteroscedasticity must be made. However, with the relatively short post-float period in Australia, such recognition is difficult. The logarithms of the exchange rate and relative prices are presented in a scatter plot in Figure 3. The 13 points in the south west corner are all post-1984 and, in some sense, define a second regime.

Although the practice of altering sample periods to remove data apparently not consistent with one's theories is fraught with problems, the fact that the second regime commenced at a well documented point in time makes it of particular interest to repeat the analysis for the shorter time period. In each of the following cases only one lag was necessary to capture all dynamic effects.

Tables 4 to 6 are replications of Tables 1 to 3 using only data from 1974.I to 1984.IV.

Table 4: ADF Tests on Price Variables Pre 1985

Variable	y	Δy
1. Exchange Rate	-0.36	-6.99
2. Aus. GDP Deflator	-1.28	-8.53
3. US Export Deflator	-2.01	-2.29
4. Ratio (3) to (2)	-0.39	-5.96
5. Ratio (1) to (4)	-2.77	-8.30

It is now apparent from Table 4 that the real exchange rate has an ADF statistic of -2.77 which is quite close to the 5% critical value and above the 10% critical value of 2.60. More importantly, the autoregressive parameter in that regressions is 0.60 and power becomes the central issue.

The smallest root in Table 5 has reduced from 0.501 to 0.215 and the accompanying ADF statistic easily rejects the nonstationarity hypothesis and a similar improvement can be noted when comparing Tables 3 and 6. The error correction exclusion statistics in these two models are 5.70 and 5.03 both of which are well above 5% critical values.

Table 5: Canonical Decomposition of 3-Equation VAR Pre 1985

Canona	Weights				
	AUD	US^e	Aus^p	λ	ADF
1	-.4554	.6561	-.6018	.215	-3.94
2	-.0763	.5080	-.8580	.878	-0.53
3	.0319	.8799	.4742	.999	-2.06

The implied coefficient in the cointegrating equation from Table 6 is 1.149 and, because of the earlier comment on the stationarity of the real exchange rate, it can be concluded that absolute PPP should not be ruled out over an extended period of floating exchange rates that do not contain major currency crises. The second analysis highlights the influential nature of the recent data and should, therefore, be treated as part of a full diagnostic analysis rather than a direct result on PPP.

Table 6: Canonical Decomposition of 3-Equation VAR Pre 1985

Canona	Weights			
	AUD	US^e/Aus^p	λ	ADF
1	-.6564	.7544	.248	-3.56
2	.0617	.9981	.983	-0.54

Conclusions

Our analysis provides some support for a weak form of PPP. That is, there exists a long-run relationship between the nominal exchange rate between Australian and US currencies and the ratio of US export prices and the Australian GDP deflator. However, the absolute version of PPP requires that the coefficient on price is unity and this is rejected by the data over the full period but, if the more recent highly volatile data are omitted, the existence of absolute PPP is not an unreasonable conjecture.

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