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Testing purchasing power parity in a multivariate cointegrating framework

FRANCIS IN and IMAN SUGEMA

Department of Economics, Deakin University, Geelong, Victoria, 3217 Australia

This paper examines a bilateral PPP (purchasing power parity) relationship between Australia and the 11 major trading countries by means of two alternative econometric techniques – a multivariate cointegrating framework and a band-spectral regression. It is acknowledged that there is no strong evidence that classical PPP holds in all cases. However, the generalized version of PPP holds in all cases, and provides a better explanation of the long-run relations between exchange rates and relative prices. The use of different price indices, i.e. CPI and WPI, lead to different estimates and hence different policy implications.

I. INTRODUCTION

The law of one price plays a central role in the theory of trade and in models of exchange rate determination. From the 1970s until the early 1980s, there was a major consensus that PPP is a long-run and not a short-run phenomenon (Artus, 1978; Kravis and Lipsey, 1978; Dornbusch, 1980; Frenkel, 1981). But others, like Adler and Lehmann (1983), found that PPP holds in neither the short run nor the long run. Interestingly, they define any long-run adjustment toward the parity as a lengthy yearly process, whereby the short-run dynamics is a monthly or quarterly phenomenon.

Regardless of whether or not the proposition holds, past empirical work on which the arguments are based are biased. The main problem of any earlier analysis is that the estimates are contaminated by mistreatment of non-stationary series. Earlier studies fail to incorporate an appropriate robust-consistent estimation procedure. Hence standard Wald statistics used in the analysis may not be valid (Fisher and Park, 1989; Corbae and Ouliaris, 1990; Patel, 1990). The other problem is the failure to incorporate the long-run and short-run estimation within one framework. Moreover, they misdefine the long-run and the short-run properties as related to the periodicity of the data. Engle and Granger (1987) suggest that a long-run relationship between two or more non-stationary variables is present if the linear combination of the variables is stationary. So the long-run relationship is not necessarily associated with a particular data periodicity.

Considering the above problems, Baillie and Selover (1987), Corbae and Ouliaris (1990), Kim (1990), Patel (1990) and Phylaktis and Kassimatis (1994) test PPP using a cointegration framework. Among them only Kim (1990), using yearly data from 1920 to 1972, has successfully provided strong evidence supporting PPP. However, Baillie and Selover (1987) using monthly data, and Corbae and Ouliaris (1990) using yearly data, fail to reject the absence of such relations. In view of the difference in the construction of price indices, Patel (1990) used a generalized version of PPP. Using quarterly data, he found no strong evidence of such a relationship. Based on the above contradictory results we have no confidence that the relationship really exists.

The main purpose of this study is to investigate a bilateral PPP relationship between Australia and the 11 major trading countries, Canada, France Germany, Italy, Japan, Korea, the Netherlands, Singapore, Switzerland, the United Kingdom and the United States. The analyses will be based upon two distinct definitions of the relationship (i.e. classical and generalized PPP) and two different price indices (i.e. consumer price index and wholesale price index). This will allow us to determine which definition and/or price index provides a better explanation of the relationship. The paper is organised as follows. Section II provides an analytical framework that implies a long run cointegrating relationship. Data specification tests (i.e. unit root and seasonal unit root tests), a brief description of the multivariate cointegration method (Johansen method) and the band-spectral regression method, along with the relevant results are presented in Section III. Section IV examines whether the classical PPP or the generalised version of PPP holds between Australia and the 11 major trading countries over the relevant periods. Section V concludes the paper.

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II. ANALYTICAL FRAMEWORK

The classical PPP relation implies that exchange rates are equivalent to the ratio of domestic prices and foreign prices in the sense that prices all over the world are dictated by the law of one price. Deviations of price movements across countries are equibalanced by exchange rates. The condition implicitly assumes an open economy free trade system in which there are no significant tariff or non-tariff barriers. Moreover, the condition can hold if, and only if, the transport costs of traded goods are not significant. Assuming the condition holds and exchange rates are fixed, appreciation in prices in one country can be transmitted to its trading partners. For the same reason, floating the exchange rate insulates the domestic economy from outside shocks.

The classical PPP doctrine can be stated as follows:

$$EX = P^*/P \tag{1}$$

where EX is the index of bilateral exchange rates, P is the index of domestic prices and P^* is the index of foreign prices. Taking the log form, Equation 1 can be stated in the cointegration framework as:

$$\varepsilon_t = \gamma_0 ex_t + \gamma_1 p_t - \gamma_2 p_t^* \tag{2}$$

Assuming that exchange rates and prices are each non-stationary $I(1)$ processes the validating condition for estimation is that ε_t should be a stationary process. In other words, the variables should be cointegrated (see Engle and Granger (1987) for a formal proof). Under the heroic classical PPP doctrine, the long run γ parameters are all supposed to be unity, so that trade barriers and other distortions should be sufficiently captured by the error term ε_t and they do not significantly affect the long-run parameters. Otherwise the classical condition does not hold.

For several reasons we have to assume that the true price indices for which the strict PPP relation holds are unknown or unobservable. First, the theory does not specify which price index should be used and under what condition the relation can be expected to hold. Second, there is no strong convention among researchers about which price index is appropriate. Current research by Fisher and Park (1989), Corbae and Ouliaris (1990), Kim (1990), and Patel (1990) use different indices. Some examples of indices available for research are the consumer price index (CPI), the wholesale or producers price index (WPI), the export price index (EPI), the import price index (IPI) and the traded goods price index.

So far only Kim (1990) strongly suggests that WPIs are more appropriate than CPIs to capture the relations. He argues that CPIs are more contaminated by non-traded goods than WPIs which put more weight on traded goods. However, we may still argue that the definition of traded and non-traded goods varies widely across countries. For example, for countries like Japan and Germany, manufactured goods are the major traded goods, but for countries like Australia it is mining and agricultural products that are the major traded goods. Hence it would be difficult to find close or similar indices for different countries unless they have similar economic structures.

Moreover, it is evident that there is strong intercountry variation in the construction of price indices (Gernberg, 1977). Each country has a different definition and puts different weights on a bundle of goods. Hence it would be difficult to expect convergent price movements if the source of the economic shock is real in origin. However, if the shock is of monetary origin, we would expect conformity of intercountry price movements (see Frenkel (1981) for further discussion).

If the true indices are not obtainable, can we test PPP using a proxy variable? If so, are such estimates valid? Suppose we have proxies for domestic and foreign price indices like WPIs and CPIs, and consider that there is a relationship between the true and observable proxy variables so that:

$$p_t = \eta_1 px_t + v_t \tag{3}$$

$$p_t^* = \eta_2 px_t^* + v_t^* \tag{4}$$

where p_t and p_t^* are the true domestic and foreign price indices which px_t and px_t^* are observable proxies of the price indices where PPP holds. The condition for the proxies to be valid indices is that they should be cointegrated with the true indices. Now if we substitute Equations 3 and 4 into Equation 2, we get:

$$z_t = \beta_0 ex_t + \beta_1 px_t - \beta_2 px_t^* \tag{5}$$

where $z_t = \varepsilon_t - \gamma_1 v_t + \gamma_2 v_t^*$

$$\beta_0 = \gamma_0$$

$$\beta_i = \gamma_i \eta_i \quad \text{for } i = 1 \text{ and } 2$$

Since the proxies are cointegrated with the true variables, we can expect Equation 5 to be valid and cointegrated. The sum of the stationary errors v_t, v_t^* and ε_t should also be stationary. The two error terms v_t and v_t^* represent factors affecting the construction of the price indices or measurement errors, while ε_t is the disturbance that captures the trade barriers, the transport costs, the change in productivity differences, tastes, etc. If those distortions are quite large, then the classical PPP may no longer hold (see Gernberg, 1977; Baille and Selover, 1987; and Patel, 1990 for discussions as to why classical PPP does not hold).

As an alternative to the classical version, following the work of Patel (1990) we may use a weaker definition of the PPP relation that does not restrict the magnitude of the coefficients to unity but instead simply restricts the sign. We call this relation a generalized version of PPP. For the purpose of our analysis we will compare which version has better explanatory power regarding the long run relations between the exchange rate and the price indices. We also want to evaluate the belief of Kim (1990) that WPIs are more appropriate than CPIs for capturing PPP relations.

III. EMPIRICAL ANALYSIS

Data specification

For our analyses, we use quarterly data of bilateral exchange rates, wholesale price indices, and consumer price indices. The data are seasonally unadjusted, and are obtained from the Reserve Bank of Australia (RBA) and International Financial Statistics published by the IMF. The period of investigation is from 1972:1 to 1986:2, except for Singapore and France. Because of the limitation in data availability, the periods for the two countries are 1974:1–1986:2 and 1972:1–1985:4 respectively. The exchange rates are defined as the unit of foreign currency per Australian dollar. All of the series are measured in indices with 1980 as the base year, to remove the effect of the differences in the currency unit on the long run coefficient. First, we employ the Phillips–Perron unit root test (PP test) using the Shazam computer package version 6.2. The results of the unit root tests are reported in Table 1.

Unit root tests indicate that each data series is $I(1)$ non-stationary in log levels. Having established that all the series are $I(1)$ in Table 1, we then tested for the possible presence of seasonal unit roots. There is a need to pay attention to the seasonal characteristics of the series to complete data specification tests. Since the data are seasonally unadjusted, the HEGY test (Hylleberg *et al.* 1990, pp. 215–38) is recommended for use. If data are seasonally adjusted, then we know from Ghysels and Perron (1993, pp. 57–98) that the size of the Phillips and Perron procedure is distorted and the test has low power. The HEGY test for quarterly data uses the following representation:

$$\Delta_4 y_t = \sum_{i=1}^{4} \alpha_i Q_{it} + \delta \text{ trend} + \pi_1 Z_1 y_{t-1} + \pi_2 Z_2 y_{t-1} + \pi_3 Z_3 y_{t-1} + \pi_4 Z_3 y_{t-2} + \sum_{j=1}^{p=1} \Psi_j \Delta_4 y_{t-j} + \varepsilon_t \quad (6)$$

where p is the number of lagged terms included, and the $Q_{i,t}$ are seasonal dummy variables. The presence of the seasonal dummies and the linear trend term are due to the arguments of Perron (1988). Let the polynomials in the lag operation be defined as follows:

$$Z_1 = (1 + L + L^2 + L^3), Z_2 = -(1 - L + L^2 - L^3) \text{ and } Z_3 = -(1 - L^2) \quad (7)$$

Hylleberg *et al.* (1990) propose a test that looks at the unit roots of all the seasonal frequencies as well as the zero frequency. The interpretation of the HEGY test can be summarized as: (1) the series is stationary and it has no unit roots, if each of the πs is different from zero; (2) there will be no seasonal unit roots if π_2 and either π_3 or π_4 are different from zero; (3) if $\pi_1 = 0$, then we have a unit root at the zero frequency; (4) if $\pi_2 = 0$, then we have a seasonal unit root at the π frequency; (5) if π_3 and $\pi_4 = 0$, then we have a pair of complex roots at the $\pi/2$ frequency. The decision rule for the HEGY tests is therefore straightforward. We reject the various null hypotheses for large values of the test statistic (in absolute terms) and this implies no unit root at that frequency.

Table 1. *Phillips–Perron unit root test results*

	CPI				WPI				EX			
	level	nl	diff.	nl	level	nl	diff.	nl	level	nl	diff.	nl
Australia	− 2.27	2	− 5.64	2	− 2.40	2	− 4.06	1				
Canada	− 1.84	3	− 3.54	2	− 2.59	5	− 4.02	4	− 0.71	0	− 6.42	7
France	− 1.60	2	− 3.14	1	− 0.98	4	− 3.14	4	− 1.48	1	− 4.49	2
Germany	− 0.35	4	− 4.26	1	− 2.62	1	− 4.06	1	− 0.34	0	− 5.75	2
Italy	− 1.30	1	− 4.60	1	− 2.55	1	− 3.63	0	− 1.85	1	− 4.78	1
Japan	− 1.24	4	− 4.46	7	− 1.07	2	− 3.41	0	− 0.84	1	− 4.78	7
Korea	− 1.76	3	− 3.69	2	− 2.19	3	− 3.71	2	− 2.27	0	− 6.65	2
Netherlands	− 0.19	4	− 4.91	4	− 2.25	4	− 5.05	4	− 0.41	0	− 5.87	2
Singapore	− 2.19	5	− 3.79	5	− 1.39	2	− 3.05	0	− 0.16	0	− 7.03	7
Switzerland	− 2.36	4	− 4.13	3	− 2.37	4	− 3.11	2	− 0.57	1	− 5.54	2
UK	− 2.54	4	− 4.22	4	− 3.15	2	− 3.99	3	− 2.30	1	− 5.36	7
USA	− 1.69	3	− 2.98	2	− 3.07	4	− 3.51	3	− 2.15	0	− 5.76	7

Notes: Critical values 90% (95%) = − 2.60 (− 2.93) – see Fuller (1976, Table 8.5.2)
CPI = consumer price index, WPI = wholesale price index, EX = exchange rates, diff. = difference and nl = indicates optimal truncation lag.

As can be seen from Table 2, we first, accept that $\pi_1 = 0$ for each series which supports a unit root at the zero frequency. In other words, all the variables show evidence of a unit root at the non-seasonal frequency in terms of the t-statistic on π_1. Second, given the critical values we accept that $\pi_2 = 0$ for most variables, which supports a seasonal unit root at the π frequency. For example, we cannot reject that $\pi_2 = 0$ for 32 out of the 35 series at the 1% level of testing. Third, we cannot reject that π_3 and/or π_4 are zero, which supports the unit root at the $\pi/2$ frequency in many

Table 2. *HEGY test results*

Null I (0, 1)	$\pi_1 = 0$	$\pi_2 = 0$	$\pi_3 = 0$	$\pi_4 = 0$	$\pi_3 \cap \pi_4 = 0$
Australia					
CPI	− 1.28	− 2.59	− 2.11	− 3.61	9.24
WPI	1.07	− 2.48	− 4.66	− 1.67	12.75
Canada					
Exchange rate	− 0.96	− 2.87	− 1.68	− 2.18	4.56
CPI	− 0.21	− 4.15	− 4.15	− 0.56	8.96
WPI	− 0.62	− 2.44	− 1.15	− 2.63	4.68
France					
Exchange rate	− 1.99	− 2.48	− 3.71	− 1.16	7.90
CPI	− 0.81	− 1.24	− 3.93	− 2.16	10.17
WPI	− 2.10	− 3.82	− 3.21	− 0.55	5.35
Germany					
Exchange rate	− 1.92	− 3.37	− 3.52	− 2.06	9.59
CPI	− 0.26	− 2.37	− 1.46	− 0.59	1.67
WPI	− 2.07	− 2.91	− 1.86	− 2.01	4.53
Italy					
Exchange rate	− 1.46	− 2.44	− 2.55	− 1.55	4.26
CPI	0.76	− 2.88	− 3.52	− 0.85	8.81
WPI	− 0.66	− 2.14	− 1.46	− 2.88	4.79
Japan					
Exchange rate	− 1.24	− 2.05	− 3.04	− 1.78	7.16
CPI	− 3.62	− 1.07	− 1.90	− 2.44	4.95
WPI	− 0.87	− 2.50	− 2.94	− 0.81	4.87
Korea					
Exchange rate	− 2.24	− 3.17	− 2.08	− 2.84	7.13
CPI	− 0.60	− 2.85	− 2.96	− 1.89	6.49
WPI	− 0.81	− 2.83	− 1.70	− 1.99	4.32
Netherlands					
Exchange rate	− 2.10	− 4.09	− 3.61	− 2.37	10.82
CPI	− 0.87	− 1.18	− 3.54	− 1.39	7.35
WPI	− 1.02	− 1.68	− 2.16	− 1.78	5.38
Singapore					
Exchange rate	− 0.74	− 2.61	− 3.36	− 2.61	10.85
CPI	− 1.45	− 1.76	− 2.54	− 1.60	4.64
WPI	− 0.08	− 2.20	− 2.90	− 1.02	5.25
Switzerland					
Exchange rate	− 2.49	− 3.38	− 3.14	− 2.19	8.36
CPI	− 3.11	− 2.90	− 2.47	− 2.91	8.40
WPI	− 2.79	− 2.47	− 2.96	− 1.36	6.39
U.K.					
Exchange rate	− 1.83	− 1.96	− 1.99	− 1.37	3.44
CPI	− 0.57	− 0.84	− 2.04	− 1.39	2.47
WPI	− 2.55	− 2.96	− 2.29	− 0.08	2.65
USA					
Exchange rate	− 0.98	− 3.52	− 1.60	− 1.75	3.19
CPI	− 0.11	− 2.68	− 2.11	− 1.16	2.91
WPI	− 0.36	− 1.74	− 3.60	− 0.08	8.85
Critical value (1%)	− 4.46	− 3.80	− 2.75	− 4.46	9.27
Critical value (5%)	− 3.71	− 3.08	− 1.91	− 3.66	6.55

Note: HEGY (Hylleberg *et al.*, 1990, see Tables 1a and 1b) test gives the critical values for the t-tests on π_1, π_2, π_3, and π_4 and the F-test on $\pi_3 \cap \pi_4 = 0$.

series, depending on the use of 5% or 1% level of testing. Overall, the results of the HEGY tests are quite conclusive and there are seasonal unit roots.[1]

Multivariate cointegrating framework: Johansen method

Johansen (1988) and Johansen and Juselius (1990) set out a maximum likelihood procedure for the estimation and testing of the cointegrating vectors in a VAR system. Suppose the vector of p-variables, $Z_t = (Z_{1t}, \dots, Z_{pt})'$, is generated by the k-order vector autoregressive process with Gaussian errors:

$$Z_t = A_1 Z_{t-1} + \cdots A_k Z_{t-k} + \mu + \Psi D_t + \varepsilon_t, t = 1, \dots, T \quad (8)$$

where Z_t is a $p \times 1$ vector of stochastic variables, $\varepsilon_1, \dots, \varepsilon_T$ are iid $N_p(0, \Sigma)$ and D_t are centred seasonal dummies, and μ is a vector of constants. Since we want to distinguish between stationarity by linear combinations and by differencing this process may be rewritten in error correction form as:

$$\Delta Z_t = \Gamma_1 \Delta Z_{t-1} + \cdots + \Gamma_{k-1} \Delta Z_{t-k+1} + \Pi Z_{t-k} + \mu + \Psi D_t + \varepsilon_t, t = 1, \dots, T \quad (9)$$

The matrix Π contains the long-run information in the system and is analogous to the error correction representation of Engle and Granger (1987). Information about the number of cointegrating vectors is found in the rank of Π. Denote the rank of Π as $r < p$. There exists a representation of Π such that $\Pi = \alpha\beta'$ where α and β are both $p \times r$ matrices. The matrix β is called the cointegrating matrix and has the property that $\beta' Z_t \sim I(0)$, where $I(d)$ indicates 'integrated of order d'. Thus we can interpret the relations of $\beta' Z_t$ as the stationary relations among potentially non-stationary variables, i.e. as cointegrating relations. Johansen (1988) and Johansen and Juselius (1990) develop a maximum likelihood estimation procedure for μ, Γ_i, α, β and Σ and also provide tests for the number of cointegrating vectors. Since Johansen's test procedure is well known, a description of the specific details is omitted here for the sake of brevity. A constant and quarterly seasonal dummies are included in the regression, although seasonal effects seem to be very small. Table 3 shows the λ-max, trace statistics and their corresponding critical values.

It is necessary to select about the number of lags in the autoregression specification. The lag lengths included in each estimation are the best lag lengths, in the sense they should be enough to whiten the error term and to leave considerable degrees of freedom. For the WPI-based analysis, there is little problem for Canada and France, since they seem to have a full rank matrix at the 10% significance level. However, since the data show that all series are stationary, there should be no full rank matrix. It is concluded that there should be at least one cointegrating vector for Canada, France, Germany, Korea, the Netherlands, Singapore, Switzerland and the USA. For Italy, Japan and the UK there should be at least two cointegrating vectors.

For the CPI-based analysis, Korea and the UK seem to have a very weak cointegrating relationship. In Korea's case, only the trace test supports the presence of one cointegrating vector at the 5% significance level. For the UK, only the λ-max test shows the presence of a cointegrating vector at the 10% significance level. However, by referring to Johansen and Juselius (1990) it is confirmed that cointegration is present for both cases. The presence of cointegration for the other countries is also quite significant. In summary, we conclude that there are two cointegrating vectors for Canada, Germany, and the Netherlands, and one cointegrating vector for the rest of the countries.

Band-spectral regression framework

In this section, the band-spectral regression, which is based in the frequency domain, has been applied and the relevant results are presented. It would have been useful to compare the time-domain and frequency-domain methods. The technique of running spectral regressions on subsets of frequencies is known as band-spectral regression. This methodology is well known and those interested in the technique should refer to Engle (1980) for its theory to Nachane and Chrissanthaki (1991) for its application. Following Nachane and Chrissanthaki (1991), the PPP relationship was tested in its relative version:

$$\Delta e_t = \alpha + \beta \Delta p_t \quad (10)$$

where e_t is the logarithm of a country's nominal dollar-dominated spot exchange rate, p_t is the logarithm of the ratio of that country's wholesale price index to the Australian wholesale price index and Δ is the first difference operator. The band-spectral results for the Australian dollar-based PPP (for the 11 countries) are reported in Table 4.

The distinct features from Table 4 are summarized as follows. (1) Full-spectrum regression (column 3) does not support PPP for any of the countries except France and Singapore and the overall R^2 values are very low. (2) For every truncation point except the truncation point of 90

[1] The existence of seasonal unit roots on the cointegration tests has some implications. Clearly, modifications to the procedure would be required in order to deal with cointegration at all seasonal frequencies. For example, Engle *et al.* (1990) develop and test a full model of seasonal cointegration at the single, two period cycle and all seasonal cycles. The cointegrating regressions are performed using filtered data which adjust the data for all unit roots except that at the cycle of interest.

Table 3. *λ-max and trace statistics*

Country	*n*-lag	λ-max			Trace test		
		$r=0$	$r \leqslant 1$	$r \leqslant 2$	$r=0$	$r \leqslant 1$	$r \leqslant 2$
		WPI-base analysis					
Canada	5	36.243	5.112	3.568	44.923	8.681	0.568
France	8	32.495	9.625	2.830	44.950	12.455	2.830
Germany	4	22.020	9.682	1.881	33.583	11.563	1.881
Italy	1	36.131	13.068	2.664	51.862	15.731	2.664
Japan	1	28.499	19.932	0.223	48.654	20.155	0.223
Korea	8	19.035	6.308	1.721	27.063	8.029	1.721
Netherlands	9	35.613	6.158	0.286	42.058	6.444	0.286
Singapore	1	32.469	11.006	0.089	43.564	11.095	0.089
Switzerland	2	26.324	10.969	2.724	40.017	13.693	2.724
UK	4	21.693	16.750	1.967	40.360	18.718	1.967
USA	3	21.934	9.941	1.990	33.865	11.931	1.990
		CPI-based analysis					
Canada	2	23.955	16.338	2.530	42.823	18.868	2.530
France	8	24.657	9.714	2.906	37.277	12.620	2.906
Germany	1	37.153	14.243	1.753	53.148	15.996	1.753
Italy	4	21.478	9.586	2.749	33.812	12.335	2.749
Japan	1	38.468	9.076	1.058	48.602	10.134	1.058
Korea	3	16.202	10.388	2.927	29.517	13.315	2.927
Netherlands	9	21.407	15.269	0.854	37.531	16.123	0.854
Singapore	5	21.255	6.360	0.760	28.374	7.120	0.760
Switzerland	1	23.001	19.091	2.907	44.998	21.980	1.541
UK	4	19.858	4.854	0.143	24.855	4.996	0.143
USA	5	25.054	9.805	2.195	37.055	12.000	2.195

Notes: Critical value (Johansen and Juselius, 1990):

90%	18.697	12.099	2.816	26.791	13.338	2.816
95%	18.697	14.036	3.962	29.509	15.197	3.962

months, the long-run version fares better than the corresponding short-run version. It is worth noting that this empirical finding is consistent with that of Nachane and Chrissanthaki (1991). (3) The long-run version of the PPP holds for France, Japan, Italy, the UK, Germany, Canada, Korea and the USA for some of the truncation points. For example, the long-run PPP holds for France for all of the truncation points except 15 months; for Japan, the corresponding periods are 6, 15 and 24 months; for Italy and the UK the corresponding periods are 15 and 24 months; for Germany the corresponding periods are 6 and 24 months; whereas, in the case of Canada, Korea and the USA, PPP holds in 90, 6, and 24 months, respectively. (4) It is interesting to note that there are some countries such as the Netherlands, Singapore and Switzerland where long-run PPP does not hold for any of the truncation points considered. (5) The intercept term in all the regressions is very close to zero, which is its expected theoretical value under PPP.

IV. TESTING GENERALIZED PPP VERSUS CLASSICAL PPP

The main focus of this paper is to examine whether the generalized PPP or the classical PPP holds between Australia and the 11 major trading countries over the relevant periods. The generalized version of PPP relates the equilibrium changes in exchange rates to changes in the ratio of foreign and domestic prices. Basically Johansen's procedure defines the long-run equilibrium error as:

$$z_t = \beta_0 ex_t + \beta_1 px_t + \beta_2 px_t^* \quad (11)$$

where β_1 and β_2 are assumed to be positive and negative, respectively. In Table 5 we report β_1 and β_2 normalized by β_0 or we set β_0 to be unity. It is concluded that generalized PPP holds if at least one vector has the correct sign. Based on Table 5 it is evident that generalized PPP holds for all countries and for both CPI and WPI based analyses. The magnitudes of the coefficients vary considerably across

Table 4. *Spectral regressions for PPP (Aus \$ base)*

Exchange rate and sample period	Regression statistics	Full-spectrum regression	Band-spectral regression: Trunc. pt freq. = 1.0476 per. = 6 mths High	Low	Trunc. pt. freq. = 0.4189 per. = 15 mths High	Low	Trunc. pt. freq. = 0.2619 per. = 24 mths High	Low	Trunc. pt. freq. = 0.0698 per. = 90 mths High	Low
Canada	Intercept	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
January 1972	Slope	− 0.62	0.56	2.61	0.42	2.76	0.49	1.87	0.87*	0.63
to August 1986	R^2	0.01	0.01	0.74	0.01	0.38	0.01	0.12	0.03	0.02
	DW	2.37	1.98	0.02	2.14	0.11	2.47	0.36	1.85	2.09
France	Intercept	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
January 1972	Slope	0.59	0.95*	1.11*	0.17	1.28	0.82*	1.18*	1.03*	0.96*
to December 1985	R^2	0.03	0.07	0.37	0.04	0.46	0.04	0.30	0.11	0.10
	DW	2.25	2.21	0.01	2.34	0.08	2.54	0.25	2.06	2.11
Germany	Intercept	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
January 1972	Slope	− 0.62	1.11*	2.93	0.72	3.16	0.97*	2.02	1.37	1.46
to August 1986	R^2	0.01	0.02	0.38	0.01	0.39	0.02	0.11	0.04	0.05
	DW	1.66	1.93	0.01	2.04	0.07	2.39	0.33	1.79	1.87
Italy	Intercept	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
January 1972	Slope	0.29	0.32	1.26	0.12	1.17*	0.18	0.85*	0.48	0.59
to June 1986	R^2	0.01	0.01	0.34	0.00	0.31	0.00	0.15	0.02	0.03
	DW	2.28	2.23	0.02	2.36	0.09	2.56	0.28	1.99	2.25
Japan	Intercept	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
January 1972	Slope	4.62	− 1.74	1.01*	− 2.65	0.99*	− 3.19	0.89*	− 0.70	− 1.00
to August 1986	R^2	0.02	0.00	0.33	0.01	0.12	0.01	0.03	0.00	0.00
	DW	1.86	2.97	0.02	2.98	0.15	3.02	0.41	2.92	3.07
Korea	Intercept	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
January 1972	Slope	0.19	0.37	0.88*	0.36	0.65	0.28	0.69	0.42	0.52
to August 1986	R^2	0.01	0.03	0.52	0.02	0.22	0.01	0.03	0.00	0.00
	DW	2.48	2.09	0.02	2.54	0.08	2.52	0.32	1.95	2.21
Netherlands	Intercept	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
January 1972	Slope	0.52	0.38	2.97	0.21	2.99	0.29	1.64	0.62	0.76
to July 1986	R^2	0.01	0.00	0.46	0.00	0.33	0.00	0.09	0.01	0.02
	DW	2.41	2.04	0.01	2.19	0.09	2.50	0.33	1.88	2.05
Singapore	Intercept	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
January 1972	Slope	0.20	0.12	0.54	0.09	0.60	0.15	0.22	0.13	0.25
to August 1986	R^2	0.03	0.00	0.48	0.00	0.15	0.01	0.01	0.00	0.01
	DW	2.44	2.14	0.03	2.28	0.14	2.61	0.37	2.03	2.27
Switzerland	Intercept	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
January 1972	Slope	− 0.25	0.42	1.54	0.15	1.81	0.22	1.40	0.69	0.71
to August 1986	R^2	0.00	0.01	0.30	0.00	0.26	0.00	0.11	0.02	0.02
	DW	1.97	1.97	0.01	2.12	0.08	2.38	0.29	1.85	1.89
UK	Intercept	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
January 1972	Slope	− 0.64	0.48	0.79	0.28	1.02*	0.26	0.90*	0.56	0.37
to August 1986	R^2	0.01	0.01	0.04	0.00	0.07	0.00	0.04	0.01	0.00
	DW	2.49	2.29	0.02	2.41	0.09	2.61	0.25	2.11	2.28
USA	Intercept	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
January 1972	Slope	0.05	0.45	2.05	0.42	1.89	0.51	1.16*	0.62	0.71
to August 1986	R^2	0.00	0.01	0.84	0.01	0.33	0.02	0.08	0.03	0.03
	DW	2.89	2.13	0.03	2.29	0.14	2.64	0.37	2.00	2.26

Notes: (*) indicates that the slope coefficients does not differ significantly from unity at the 5% level. Following Nachane and Chrissanthaki (1991), four frequency truncations are considered, 0.0698, 0.2619, 0.4189 and 1.0476 which correspond to periods of 90, 24, 15 and 6 months, respectively. The 'high' and 'low' columns really capture the short-run and long-run effects, respectively, of PPP.

Table 5. *Estimates of cointegrated vectors*

Country	Vector 1 β_1	Vector 1 β_2	Vector 2 β_1	Vector 2 β_2
	WPI-based analysis			
Canada	2.886	− 3.257		
France	1.703	− 1.895		
Germany	2.751	− 5.158		
Italy	2.153	− 4.546	1.950	− 0.078
Japan	2.161	− 2.832	0.922	− 1.530
Korea	2.262	− 1.461		
Netherlands	3.641	− 2.671		
Singapore	1.206	− 0.909		
Switzerland	1.467	− 2.699		
UK	0.490	− 0.391	− 0.041	− 1.073
USA	1.968	− 2.141		
	CPI-based analysis			
Canada	0.589	− 0.543	1.246	− 2.332
France	2.577	− 2.438		
Germany	5.569	− 11.075	4.149	− 8.876
Italy	2.497	− 1.815		
Japan	0.320	− 0.764		
Korea	1.307	− 1.048		
Netherlands	− 4.772	9.029	0.503	− 1.232
Singapore	2.039	− 3.186		
Switzerland	3.713	− 7.156	0.887	− 1.360
UK	0.401	− 0.058		
USA	7.236	− 7.315		

countries. Furthermore, there is no consistency in the magnitude obtained from the WPI and CPI-based regressions. This indicates that the measures the considerably different in reflecting the relations.

In using the CPI, it is interesting to find a very unreasonable magnitude of the coefficient for Germany, Switzerland and the USA. Similarly with the use of the WPI, the magnitudes of the b vectors for Germany and Italy seem to deviate too widely. However, as Juselius (1991) argues, the β vectors obtained from the decomposition of the Π matrix are not unique, but the cointegrating space is unique. Since the research focus is to find the cointegrating spaces, the magnitude has less significance. The classical PPP relation implies that β_1 and β_2 should be 1 and -1 respectively, therefore we have two restrictions in the cointegrating space. Holding the dimension of the cointegrating space fixed, consider the hypothesis matrix $H = (1 \;\; 1 \;\; -1)$ such that $\beta = H\Phi$ where Φ is an unknown weighting matrix. The corresponding test statistics is:

$$-2\ln\Phi = T\sum_{i=1}^{r}\ln[(1-\sigma_i)/(1-\tau_i)] \qquad (12)$$

This statistic has a chi-square distribution with $r(p-s)$ degrees of freedom. Table 6 shows the test statistics and their corresponding critical values.

Table 6. χ^2 *statistics for testing classical PPP*

Country	WPI χ	WPI df	CPI χ^2	CPI df
Canada	32.25	2	1.46	1
France	10.01	2	7.05	2
Germany	16.47	2	12.48	1
Italy	9.74	1	8.08	2
Japan	9.15	1	9.19	2
Korea	13.45	2	5.81	2
Netherlands	27.49	2	4.86	1
Singapore	23.97	2	7.97	2
Switzerland	13.29	2	3.11	1
UK	1.51	1	15.17	2
USA	15.87	2	20.03	2

Note: df = freedom and critical value (95%): $\chi^2(1) = 3.84$, $\chi^2(2) = 5.99$.

At the 5% significance level, in the WPI based analysis, classical PPP holds only for the UK. A better result is obtained in the CPI-based analysis, where four countries confirm the relations, (Canada, Korea, Netherlands, and Switzerland). However, empirical findings are still not strong enough to support the classical proposition.

V. CONCLUDING REMARKS

This study has provided an investigation of a bilateral PPP relationship between Australia and the 11 major trading countries by means of cointegration and alternative econometric techniques and band-spectral regression. The results seem to provide very weak evidence supporting classical PPP relations. Consequently, the relation should not be imposed as an assumption in building an economic model; nor is it a useful guide for policy purposes.

It seems clear that any course of action based on the classical PPP relation will be misleading. The departure from PPP can be due to deviations in productivity differentials, changes in tastes, and shifts in comparative advantage (Baillie and Selover, 1987), measurement error in the construction of price indices (Gernberg, 1977) and tariff and non tariff barriers (Frenkel, 1981). Those problems multiplicatively cause non-stationary errors while classical PPP requires stationarity. In the light of the problem in the construction of the price indices, the performance of CPIs and WPIs in reflecting the PPP relationship has been compared. The two measures produce different estimates, but these did not show which one is better. The results contradict the contention of Kim (1990), who suggested that WPIs are more appropriate than CPIs for capturing PPP relations. So the potential danger in applying PPP to price indices such as CPIs, expressed by Kim (1990), should be treated cautiously. Moreover, one should be aware that the two measures produce different policy implications.

Following the work of Patel (1990), the presence of a generalized version of PPP has been tested. There is strong evidence that the generalized version provides a better explanation of the equilibrium relationship between exchange rates and the price level, i.e. exchange rates are inversely related to domestic–foreign price ratios. Distortions in international markets are not enough to be represented as a stationary process in which classical PPP holds. Researchers should impose only a sign restriction in the model formulation, and not magnitude restrictions. Indeed, the co-movement between price ratios and exchange rates is present but has different speed. This indicates that each country has a different speed of adjustment to an external shock.

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