Tests of Purchasing Power Parity

by

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(ABSTRACT)

(This paper examines the long-run relationship between exchange rates and prices in ten countries in Southwest Asia, Africa, and the Pacific Rim for the post-Bretton Woods period. It uses cointegration tests to investigate the thesis that relative purchasing power parity exists as a long-run equilibrium condition between country-pairs. It expands upon tests for relative purchasing power parity suggested by previous authors by pretesting price index time series for structural breaks, in addition to pretesting the price indices and exchange rates for compatible stochastic properties. It compares the results of conventional cointegration tests for parity with a weaker form of the relationship suggested by Pippenger (1993) and Patel (1990), and finally, examines purchasing power parity by testing real bilateral exchange rates for stationarity.)
DEDICATION

This paper is dedicated to my children Lesley, Pete, Stephea, and Michael for their patience and forebearance during a cold, snowy winter.

Dad
ACKNOWLEDGEMENTS

I need to acknowledge Dr. David Meiselman's advice and counsel in suggesting this topic as a fertile field in economic research, and for timely navigation on the path. I also want to extend a special thanks to Dr. Russ Porter for sharing his intuition on the econometrics and quantitative tools.
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I. INTRODUCTION.

The purchasing power parity principle has been a cornerstone of many exchange rate models. It asserts there is a relationship between the market exchange rate of two countries and their price levels. While the theory has intuitive appeal, empirical tests over the past 15 years--particularly those covering the post-Bretton Woods period-- have been largely inconclusive.

The fundamental idea underlying the purchasing power parity theory is, at least in the long run, the law of one price holds for traded goods. When Cassel (1916, 1921) first proposed the concept, he produced evidence of a correlation between the exchange rates and the price indices of France, Germany, Sweden, Russia, and the United Kingdom. In the modern period of floating exchange rates, many countries have experienced persistent inflation which has been reflected in non-stationary price indices. The data generating processes for these time series appear to contain unit roots, and the presence of unit roots presents significant problems for conventional statistical tests.

The need to provide more conclusive empirical evidence on the purchasing power parity theory, and the need to more fully explore the appropriate application of unit root testing to this aspect of economic theory, provides the motivation for this research.
II. THE THEORY OF PURCHASING POWER PARITY.

There are two general interpretations of the parity relationship which appear in economic literature--absolute purchasing power parity and relative purchasing power parity. In absolute purchasing power parity, the law of one price is established as an arbitrage condition between countries. It assumes transactions costs are negligible. However, the assumption of small transaction costs in this interpretation is very strong and there is little empirical research supporting the commodity arbitrage view in the current period of floating exchange rates.

There are numerous possible explanations for the observed violations. Uppal (1992) shows absolute purchasing power parity may not hold if there are transportation costs, tariffs, or other barriers to trade. He also suggests that even when the law of one price holds, absolute purchasing power parity may not hold in the presence of non-traded goods, or because of differences in consumption tastes. Kravis and Lipsey (1977) suggest that divergences may also exist because many firms in international trade are in the position of being a discriminating monopolist--operating in markets with different demand elasticities. They also suggest that without perfect information between markets, the adjustment to changes in comparative advantage may be gradual rather than instantaneous.

By contrast, the interpretation as relative purchasing power parity is essentially a monetarist view. It establishes a relationship between exchange rates and price indices.
While this research is not directly concerned with exogeneity issues, the theory of purchasing power parity is consistent with two common exchange rate models. The first posits a causal chain running from monetary disturbances to prices, and from prices to exchange rates. The second is an asset approach, where currency demand stems from investors diversifying international portfolios in response to perceived changes in real domestic economic conditions. In the asset model, international capital flows determine exchange rates and prices simultaneously as endogenous variables.

Due to the somewhat formidable problems in testing absolute purchasing power parity, the empirical work has focused almost exclusively on relative purchasing power parity and price indices. Given the important distinctions between these forms of parity, and its relevance for both the intuition and the models, a more formal review of the development of the model is appropriate. Pippenger (1993) defines absolute purchasing power parity as the condition where:

$$E_t = \frac{DPL_t}{FPL_t}$$

(1)

where $E_t$ is the domestic price of a foreign currency, $DPL_t$ is the domestic price level (a weighted average of domestic prices) at time $t$, and $FPL_t$ is the foreign price level (a weighted average of foreign prices) at time $t$.

The addition of a simple constant scalar, $K$, to the equation establishes the condition for relative purchasing power parity, and the equation takes the form:

$$E_t = K \left( \frac{DPL_t}{FPL_t} \right)$$

(2)
Since the empirical work requires the use of price indices, it is essential to have an equation for relative purchasing power parity expressed in terms of the change in exchange rates and price levels. This can be derived by dividing relative purchasing power parity, in terms of price levels in period $t$, by the same variables in a base period $0$. This gives us the expression:

$$
\frac{E_t}{E_0} = \frac{\frac{DPL_t}{DPL_0}}{\frac{FPL_t}{FPL_0}} = \frac{DP_t}{FP_t}
$$

where $DP_t$ is the nominal value of a domestic price index and $FP_t$ is the nominal value of a foreign price index. These indices have a common base period $0$, and in an important aspect for later discussion, we can eliminate the constant $K$ in this expression of relative purchasing power parity.

Relative purchasing power parity states the percentage change in the exchange rate over a given period equals the ratio of the percentage change in the domestic country's price level and the percentage change in the foreign country's price level for the same period. Essentially, the difference between absolute purchasing power parity and relative purchasing power parity is that while absolute purchasing power parity
concerns price levels and ultimately price indices, relative purchasing power parity concerns only price indices.

Taking the logarithmic version of the expression above yields:

\[
(4) \quad e_t = e_o + dp_t - fp_t
\]

where \(e_t, e_o, dp_t,\) and \(fp_t\) are the logarithms of \(E_t, E_o, DP_t,\) and \(FP_t.\) Since we cannot be assured that relative purchasing power parity holds without error in the base period, we can let \(\alpha\) equal the sum of the error and the log of the exchange rate in the base period. Adding \(u_t\) lets us allow for temporary deviations in relative purchasing power parity, and gives us an equation we can estimate with OLS:

\[
(5) \quad e_t = \alpha + dp_t - fp_t + u_t
\]

This is the basic regression for the relative purchasing power parity model which appears (with minor algebraic variations) throughout the literature. In empirical practice, most researchers either, a) construct a “real exchange rate” from bilateral exchange rates and either wholesale or consumer price indices and interpret the purchasing power parity test as a test of stationarity for the real exchange rate series, or b) interpret the purchasing power parity test as a cointegration test for a long run equilibrium condition between exchange rates and price indices.

The concepts of stationarity, non-stationarity, and cointegration are central to the theory and tests and merit further discussion. Enders (1995) identifies a stationary time series as one which, 1) exhibits mean reversion in that it fluctuates around a constant
long run mean, 2) has a finite variance that is time-invariant, and 3) has a theoretical correlogram that diminishes as lag length increases.

By contrast, a nonstationary time series has permanent components and the mean and/or the variance is time dependent. In a nonstationary series, there is 1) no return to a long-run mean, 2) the variance is time dependent and goes to infinity as time approaches infinity, and 3) theoretical autocorrelations do not decay, but the sample correlogram dies out slowly in finite samples. A time series which becomes stationary after differencing x times is described a being integrated of order x, or I(x).

The concept of cointegration concerns long-run equilibrium relationships between two or more time series variables. Integrated variables are said to be cointegrated if there is a stationary linear combination of them. This linear combination, or cointegrating vector, is simply the coefficients of the cointegration regression. The stationarity of the linear combination is measured with tests on the residuals. The economic interpretation of cointegration is that of a stable relationship over time, i.e., the variables cannot meander without a tendency to return to a long-run equilibrium relationship. Cointegration testing for nominal exchange rates and price indices thus represents a somewhat natural test of relative purchasing power parity.

However, As Greene (1993) notes in his discussion of unit roots, the issue of stationarity in unit root and near-unit root processes rests on a “razor’s edge.” If the estimate of ρ, in the regression of variable on a constant and its lagged value,
(6) \[ y_t = \alpha + \rho y_{t-1} + \epsilon \]

is equal to one, the data generating process follows a random walk. The mean of \( y \) is not stationary over time. By contrast, if \( \rho \) is less than one, the time series is stationary and reverts to its mean. For the immediate issues of this research, a non-stationary real exchange rate would refute the hypothesis of parity as a long-run equilibrium condition. Changes in the real exchange rate would be expected to be permanent and deviations from purchasing power parity would be expected to become unbounded as the forecast horizon becomes longer. A finding that \( \rho \) is less than one, even if very close to one such as .98 or .99, would support the hypothesis of a long-run equilibrium condition.

A final consideration in the development of the model is the issue of a priori restrictions on the cointegrating vector. Most researchers including Enders (1988), Corbae and Ouliaris (1988) impose a restriction on the cointegrating vector, specifically \( \alpha = (1,1,1) \) or \( \alpha = (1,a,a) \) on the vector \((e_t, dp_t, fp_t)\). In other words, there is a requirement in their models that the coefficients for \( dp_t \) or \( fp_t \) to be equal to one, or to each other, to maintain relative purchasing power parity.

Pippenger (1993) asserts the restriction is inappropriate given the nature of the price indices, and that imposing such restrictions on the cointegrating vector may lead to false rejection of cointegration. While he accepts the propriety of the restriction on theoretical grounds, he notes that price indices do not have identical weights across
countries, and following an argument developed by Patel (1990), suggests the coefficients of the price indices could not be expected to equal to one, or to each other.

While the argument on weighting in the indices is persuasive, the shift from the univariate model (real exchange rates) or bivariate model (cointegration of exchange rates and a ratio of price indices) to a trivariate model, where prices indices are treated as fully independent variables, opens additional questions on the character of the test and what precisely is being measured. The cointegration test clearly still establishes a long run equilibrium relationship between the three variables, and it follows that the behavior of at least one variable must be restricted by the values of the others. However, whether or not the trivariate model describes a "parity" condition is far from certain. At best, we are left with a significantly weaker form of the parity relationship.

The absence of the restriction on the cointegrating vector re imposes a constant $K$ in equation (3) to yield,

\[
E_t / F_o = K \left( \frac{DPL_t}{DPL_o} / \frac{FPL_t}{FPL_o} \right) = K \left( \frac{Dp_t}{Fp_t} \right)
\]

and we cannot be assured that $K=1$, or is solely a statistical correction for the price indices. In this formulation of the parity condition, as $K \rightarrow 0$, the price indices become increasingly irrelevant measures of price levels.
III. REVIEW OF THE LITERATURE.

There is an extensive, inconclusive literature reporting on the results of empirical tests for real exchange rates and purchasing power parity. Roll (1979), Frenkel (1981), Adler and Lehmann (1983), and Hakkio (1984, 1986) find support for the random walk hypothesis for real exchange rates. And in papers closely related to this research, Mark (1986), Mecagni and Pauly (1987), and Corbae and Ouliaris (1988) find that nominal exchange rates are not cointegrated with price indices. Enders (1988) reports mixed results in similar research on real exchange rates, but concludes it is not possible to reject the random walk hypothesis. Patel (1990) removes a priori restrictions on the cointegrating vector and also reports mixed results in cointegration tests.


These tests for relative purchasing power parity vary considerably in their approach to the data and testing methods, and in general, the literature is characterized by progressive refinements in both. The later research naturally uses larger sample sizes, and demonstrates increasing concern on the stochastic properties of the time series. We can draw only two general conclusions on the issues from the review of the literature:
a. Wholesale price indices provide better support for the purchasing power parity hypothesis than consumer price indices—an unsurprising observation given the theoretical foundations, and,

b. Pretesting the time series for compatible stochastic properties is an extremely important aspect of the methodology for cointegration tests, and the stochastic properties of exchange rates and price indices may not be consistent over time.

As an illustration of the latter, Pippenger (1993) finds the wholesale price index for the United States and the industrial production index in the United Kingdom to be I(2) series for the period between 1973 and 1988. Given that comparable indices from other countries are generally I(1) series, he appropriately questions the validity of prior cointegration tests which use the United States as a base country and rely upon compatible stochastic properties in the time series. However, any conclusion that the US and UK price indices are I(2) as a long-run condition would imply consistently accelerating inflation in these countries over time, and preclude any possibility of finding relative purchasing power parity as a generalized condition.
IV. THE DATA. The data for this research consists of monthly observations from the International Monetary Fund's International Financial Statistics series on compact disk for the period January 1973 through June 1994. The countries include Japan, Korea, Thailand, Australia, Philippines, India, Pakistan, Indonesia, South Africa, and Singapore. The exchange rates are the monthly end of period rates. The price indices are the reported wholesale prices indices for all countries with the exception of Australia, where the comparable Industrial Goods Index is used due to the unavailability of a wholesale price index. Since a wholesale price index for Singapore is not available prior to January 1974, all statistical tests using the variable rely upon the slightly shorter period.

Japan is chosen initially as the base (domestic) country due to the scope of its economy and influence in exchange markets, industrial exports, and extensive bilateral trade relationships. Given that the indices for most countries in the sample are comparable, we consider the most important consideration is reducing the chance of rejecting the hypothesis of cointegration due to the influence of transactions costs. India is used as the base country where the stochastic properties of the variables makes the use of Japan inappropriate.

The statistical tests are conducted using Econometrics Toolkit, Version 3.0.
V. THE TESTING METHODS. If the exchange rate and price index series for a
country-pair are cointegrated, we can conclude there is a long-run equilibrium condition
among the variables. Testing for cointegration, or more precisely the lack of
cointegration, among a group of variables thought to be integrated of of order one
requires two steps.

a) Proving the time series $e_t$, $dp_t$, and $fp_t$ are integrated of order 1, $I(1)$. This uses an Augmented Dickey-Fuller (ADF) test for unit roots on both the exchange rate series (using a selected base country) and the wholesale price index for each country.

b) Proving the residuals from the cointegration regressions of the
country-pairs are stationary, $I(0)$. This also uses ADF tests for unit roots.

Our test for unit roots follows a sequence suggested by Dickey and Pantula
(1987) which relies upon the fact that first differencing a series with a single unit root
results in a stationary series. Because our lengthened sample increases the probability of
structural breaks in the series, and the ADF test has low power in detecting trend
stationary processes in the presence of structural breaks, we supplement conventional

Letting $x$ represent each series, the Dickey-Pantula sequence calls for a null
hypothesis of a unit root and estimation of the following regression with OLS to test:
\[
(8) \quad \Delta x_t = a_0 + a_2 t \text{ (time)} + \gamma x_{t-1} + \sum_{i=2}^{p} \beta_i \Delta x_{t-i-1} + \epsilon_t
\]

(In the formula above, the more familiar procedure of testing for \( a_1 = 1 \) in the model \( x_t = a_1 x_{t-1} + \epsilon_t \), has been modified by subtracting \( x_{t-1} \) from both sides of the equation to write the equivalent form \( \Delta x_t = \gamma x_{t-1} + \epsilon_t \), where \( \gamma = a_1 - 1 \). Testing the hypothesis \( a_1 = 1 \) is equivalent to testing the hypothesis \( \gamma = 0 \).)

And to test for a unit root in the series \( \Delta x \), we estimate the following regression with OLS:

\[
(9) \quad \Delta^2 x_t = a^*_0 + \gamma^* \Delta x_{t-1} + \sum_{i=2}^{p} \beta^*_i \Delta^2 x_{t-i-1} + \epsilon^*_t
\]

Thus our null hypothesis for equation (8) is \( (a_0, a_2, \gamma) = (a_0, 0, 0) \), and the appropriate test statistic is \( \Phi_3 \) from Dickey and Fuller (1981). If we cannot reject the null hypothesis from equation (8), we conclude the time series does not include a time trend, and omit this deterministic regressor to improve power in testing the first differences for stationarity in equation (9). Our null hypothesis in equation (9) is \( \gamma^* = 0 \), and the appropriate test statistic is \( \tau_4 \) also from Dickey and Fuller. If we cannot reject the null hypothesis of unit root in equation (8), but can reject the null hypothesis of a unit root in equation (9), we conclude the series is \( I(1) \).

The autoregressive lag length \( p \), the inclusion of a time trend, and the appropriate test statistics all merit further discussion. In Augmented Dickey-Fuller tests,
the lag length (p) should be chosen to induce the residuals $\epsilon_i$ and $\epsilon^*_i$ into a white noise series. Most authors suggest the problem should be approached with data dependent methods, such as using the Box-Ljung Q statistic or minimizing the Akaike Information Criterion. Including too many lags reduces the power of the test to reject the null hypothesis of a unit root since the increased number of lags results in the loss of degrees of freedom. On the other hand, too few lags will not fully capture the error process, and result in poor estimates of $\gamma$ and its standard error. For this research, we start with relatively long lag lengths and pare them down by examining the significance of the $t$ statistic on the last lag. Given an extensive sample, we chose the minimum lag length consistent with a 95% significance level for the Box-Ljung Q statistic.

The null hypothesis tested in equation (8) examines whether or not the data generating process is a random walk or a random walk with drift; the possibility of a deterministic trend in the time series is included in the alternative hypothesis. The cointegration tests used here are conducted properly only on variables with compatible stochastic properties. For example, a trend stationary process cannot be cointegrated with an $I(1)$ process, and an $I(1)$ process with drift and linear trend cannot be cointegrated with an $I(1)$ process with drift.

The adequacy of pretesting methods in making such distinctions is an important consideration. When a time trend is not included in the equation, the ADF tests (as well as the widely used Phillips-Perron tests) have little power to distinguish between trend
stationary and drifting processes. In finite samples, Enders (1995) has shown any trend stationary process can be well approximated by a unit root process, and a unit root process can be well approximated by a trend stationary process.

The Dickey-Fuller $\tau$ and $\Phi$ statistics are computed as conventional $t$ and $F$ statistics, but under the null hypothesis of a unit root there is no mean reversion in our variable, and its variance is infinite in the limit. The distribution of $\tau$ and $\Phi$ thus differ from the $t$ and $F$ statistics. The critical values are significantly higher, and moreover, the critical values increase with the number of deterministic regressors included in the model. The critical values for the ADF test were computed by Fuller (1976) and appear in a number of texts on applied econometrics. For the samples of 250 observations used here, the 95% significance levels for $\tau$, is $-2.88$ and for $\Phi$, is $6.34$.

We next consider the issue of structural breaks. Our use of sample spanning more than 20 years significantly increases the chances of encountering structural breaks, and Perron (1989) has shown that standard unit tests against trend stationary alternatives cannot reject the unit root hypothesis (even asymptotically) if the true data generating process is stationary around a trend function with a one-time break. While Perron's findings have broad significance in issues pertaining to the characterization of macroeconomic time series, our immediate concern is simply determining whether or not our test series have compatible stochastic properties.
There are few obvious events, other than the time span, which would suggest structural breaks in the 1973-1994 period. Thus it is important to remain sensitive to concerns on "data mining" the choices of possible dates. Our method simply consists of visual comparisons of the time series plots, and introduces formal tests for structural change as a pretesting step only when suggested by concurrent shifts in the direction of the data.

For the formal test, Perron provides a series of tests which permit us to measure shifts in drift and trend, as well as one-time shocks in the data generating process. In the most general model,

\[ x_t = \alpha_0 + (\alpha_1 - \alpha_0)D_{U_t} + \beta_1 t + (\beta_2 - \beta_1)D_{T_t} + D_{TB_t} + \chi x_{t-1} + \sum_{i=1}^{k} \theta_i \Delta x_{t-1+i} + \epsilon_t \]

three dummy variables (including a pulse variable \(D_{TB_t}\)) have been inserted in the ADF test and with values set as,

\[ D_{TB_t} = 1 \text{ if } t = \text{time break} + 1, \text{ and } 0 \text{ otherwise} \]

\[ D_{U_t} = 1 \text{ if } t > \text{time break}, \text{ and } 0 \text{ otherwise, and} \]

\[ D_{T_t} = t \text{ if } t > \text{time break}, \text{ and } 0 \text{ otherwise.} \]

The parameters of interest in this equation are the \(\tau\) statistics, and particularly, the statistic for the test \(\chi = 1\). The position of the time break in the series also affects the critical values of the statistic. If the proportion of observations before the time break equals 0 or 1, the test statistics are the same as those for the ADF tests. However, the
critical values of the Perron t statistic are somewhat larger than the ADF statistics in absolute value over the range and reach maximum difference when the time break occurs at the mid-point.

From this point our testing method splits in two directions. If the ADF tests verify the variables are I(1) series and there is no evidence of structural breaks invalidating the method, we proceed with cointegration tests. If Perron tests indicate the price indices are trend stationary with structural breaks, we must rely on the test for real exchange rate stationarity for the relevant country-pairs.

The next step in the cointegration test is to estimate the long-run equilibrium relationship. We repeat the cointegrating regression for convenience.

\[ (11) \quad e_t = \alpha + \beta_1 d_p_t - \beta_2 f_p_t + u_t \]

If the variables are cointegrated, an OLS regression yields a super-consistent estimator of the parameters. Stock (1987) proves the OLS estimates converge faster than in OLS models using stationary variables. The reason is the effect of integrating trend dominates the effect of the stationary components.

The formal test for cointegration examines the residuals from estimating this equation. If the deviations from the long-run equilibrium are found to be stationary, the variables are cointegrated. Our test again is the ADF test, and since we use regression residuals, there is no need to include an intercept term.

\[ (12) \quad \Delta u_t = \gamma_0 u_{t-1} + \sum_{i=1}^{p} \gamma_i \Delta u_{t-i} + \epsilon_t \]
In most studies, as Enders (1995) explains, it is not possible to use the ADF test statistics. The problem is the cointegration regression provides us only an estimate of the error $u_t$, and since the residual variance is made as small as possible, the procedure is prejudiced toward finding a stationary error process in equation (12). The test statistic $\gamma_o$ must reflect this fact, and the appropriate tables are found in Engle and Yoo (1987). For a 200 observation sample, the critical value of the $\tau$ statistic ($t$ statistic on $\gamma$) at 95 per cent is 3.37 in the 2 variable case, and 3.78 for the 3 variable case.

To provide a measure of relative purchasing power parity in those cases where the stochastic properties of the variables precludes the appropriate use of cointegration, and to compare results of tests for cointegration, we conclude by examining the stationarity of the real bilateral exchange rates $r$, where:

\begin{equation}
(13) \quad r = e_t + f_p_t - d_p_t
\end{equation}

Our test of relative purchasing power again uses the ADF, and in this application, we include a constant but omit the deterministic time trend. Our regression takes the form:

\begin{equation}
(14) \quad \Delta x_t = a_0 + \gamma x_{t-1} + \sum_{i=2}^{n} \beta_i \Delta x_{t-i} + \epsilon_t
\end{equation}

and the null hypothesis ($\gamma=0$) is that the data generating process for the real exchange rate is a random walk possibly with drift. It is important to note the theory of purchasing power parity does not allow for a deterministic time trend, or multiple unit roots. Any
such findings would refute the theory as posited. The critical values for the $\tau_\alpha$ statistic for the 250 observation sample is 2.88.
VI. RESULTS OF THE TESTS.

In pretesting the wholesale price indices for unit roots, the Dickey-Pantula test sequence confirms the wholesale prices are I(1) series for all countries tested with the exception of Australia and Singapore. The results are reported in Tables 1 and 2. Although not summarized here, the Australian series shows a deterministic time trend. Singapore was not included in the initial group selected for testing.

Because of the lengthened sample, we incorporate a Perron test for structural change. We add Singapore to the test group (with a slightly shorter available wholesale price index series, 1974-1994) and, based on simple visual examination, hypothesize a structural break at March, 1981. The test for structural change rejects the hypothesis of a unit root in the wholesale price indices for Japan, Korea, Singapore, and Indonesia at the 99 per cent significance level. The results support a conclusion these series are trend stationary, possibly with drift, with a concurrent structural break in March, 1981.

With incompatible stochastic properties, we limit further unit root testing for the exchange rates to the countries with I(1) wholesale price series--Philippines, India, Thailand, South Africa, and Pakistan. We select India as our base country for cointegration, and examine the stationarity of the four exchange rate series and report the results in Tables 4 and 5. As shown in Table 5, we cannot reject the hypothesis of a unit root in the first differenced form of the South African-Indian exchange rate.
Having completed a somewhat rigorous pretesting process, we are left with two country groups. Group I consists of India, Thailand, Pakistan, and the Philippines. Each of these countries has price and exchange rate series integrated of order one and appropriate for cointegration tests. Group II consists of Japan, Korea, Indonesia, Singapore, Australia, and South Africa. These countries have trend stationary price indices which are inappropriate for cointegration tests, or in the case of South Africa, a trend stationary exchange rate with the selected base country.

We proceed with cointegration tests for Group I and compare the results of tests both with and without a priori restrictions on the cointegrating vector. These results are presented in Tables 6, 7, and 8. We find some evidence of stationary residuals, and thus a long-run equilibrium relationship, in only the unrestricted form of the cointegration test for the Philippines at slightly less than the 90 per cent significance level.

These weak results for cointegration between India and the Philippines, however, do not strongly support the further hypothesis of relative purchasing power parity. The coefficients for the cointegrating vector are (1, 1.67, 1.08) and indicate strong divergence of the price indices in the equilibrium condition, i.e. providing a scalar significantly different from 1 for the relationship described in equation 7. Moreover, while the coefficients for the cointegrating regression with Pakistan are very small and indicate strong covariance in the price levels in Pakistan and India, the coefficient reported for fp, has an unexpected negative sign.
Because of the limited scope and poor results from our cointegration tests, we expand our scope with a more generalized test of the stationary of real bilateral exchange rates using Japan as the base country. There are no requirements for pretesting variables in this test, and the results are summarized in Table 9. We find we cannot reject the null hypothesis that the data generating process follows a random walk, possibly with drift, in any of the bilateral relationships examined.
VII. CONCLUSIONS.

These results provide scant evidence of relative purchasing power parity as a long-run equilibrium condition for the countries tested in the post-Bretton Woods period. As always, rejection of a null hypothesis does not provide a proof of the alternative. However, we have demonstrated the absence of the parity condition in the presence of a strong price level covariance (Pakistan-India) and a concurrent structural break in the price level trends for four Asian nations. This is not sufficient to reject the purchasing power parity theory, but we conclude anecdotally the results are inconsistent with the notion of a causal chain leading from monetary disturbances to price levels, and from price levels to exchange rates.

We have also illustrated some of the statistical problems associated with testing for unit roots, and provided proof of trend stationarity with structural breaks in the data generating processes for several Asian wholesale price indices. This is somewhat comforting from a theoretical perspective since finding unit roots in macroeconomic time series has implications which are potentially profound. If there are time series which are truly integrated of order one, one would have to conclude temporary shocks (such as monetary shocks) have permanent effects. This would be a finding inconsistent with much of current economic theory.
Table 1--Unit Root Tests for Wholesale Price Indices

<table>
<thead>
<tr>
<th>Country</th>
<th>Lag</th>
<th>Sig of Q</th>
<th>a₂t</th>
<th>τₑ (a₂t)</th>
<th>γ</th>
<th>τₑ (γ)</th>
<th>Φ₃</th>
</tr>
</thead>
<tbody>
<tr>
<td>Japan</td>
<td>17</td>
<td>.9992</td>
<td>.00000</td>
<td>-1.47</td>
<td>-.0056</td>
<td>-1.65</td>
<td>3.327</td>
</tr>
<tr>
<td>Korea</td>
<td>22</td>
<td>1.000</td>
<td>.00000</td>
<td>0.282</td>
<td>-.0067</td>
<td>-1.84</td>
<td>3.566</td>
</tr>
<tr>
<td>Thailand</td>
<td>18</td>
<td>.9884</td>
<td>.00002</td>
<td>0.759</td>
<td>-.0088</td>
<td>-1.20</td>
<td>1.307</td>
</tr>
<tr>
<td>Philippines</td>
<td>24</td>
<td>.9962</td>
<td>.00150</td>
<td>1.511</td>
<td>-.1361</td>
<td>-1.07</td>
<td>2.111</td>
</tr>
<tr>
<td>Australia</td>
<td>24</td>
<td>1.000</td>
<td>-.0001</td>
<td>-1.65</td>
<td>.0039</td>
<td>0.83</td>
<td>3.646</td>
</tr>
<tr>
<td>Indonesia</td>
<td>5</td>
<td>.9998</td>
<td>.00014</td>
<td>1.245</td>
<td>-.0197</td>
<td>-1.85</td>
<td>4.306</td>
</tr>
<tr>
<td>India</td>
<td>24</td>
<td>.9720</td>
<td>.00028</td>
<td>2.694</td>
<td>.0429</td>
<td>-2.59</td>
<td>4.205</td>
</tr>
<tr>
<td>Pakistan</td>
<td>6</td>
<td>.9971</td>
<td>-.0008</td>
<td>-.657</td>
<td>.1615</td>
<td>0.88</td>
<td>2.273</td>
</tr>
<tr>
<td>S. Africa</td>
<td>24</td>
<td>.9517</td>
<td>.00019</td>
<td>0.932</td>
<td>-.0189</td>
<td>-0.97</td>
<td>1.351</td>
</tr>
</tbody>
</table>

This is equation (8) testing the joint hypothesis γ = a₂ = 0. The appropriate test statistic is Φ₃. The critical values of Φ₃ for a sample of 250 observations are 6.34 at the 95 per cent level and 8.43 at the 99 per cent significance levels.
Table 2--Unit Root (First Difference) Tests for Wholesale Price Indices

<table>
<thead>
<tr>
<th>Country</th>
<th>Lag</th>
<th>Sig of Q</th>
<th>$\gamma$</th>
<th>$\tau_\gamma$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Japan</td>
<td>17</td>
<td>.9992</td>
<td>-0.2658</td>
<td>-3.888</td>
</tr>
<tr>
<td>Korea</td>
<td>22</td>
<td>1.000</td>
<td>-0.2475</td>
<td>-2.560</td>
</tr>
<tr>
<td>Thailand</td>
<td>18</td>
<td>.9794</td>
<td>-0.5091</td>
<td>-3.551</td>
</tr>
<tr>
<td>Philippines</td>
<td>13</td>
<td>.9645</td>
<td>-0.4781</td>
<td>-4.347</td>
</tr>
<tr>
<td>Australia</td>
<td>24</td>
<td>1.000</td>
<td>-0.1425</td>
<td>-1.221</td>
</tr>
<tr>
<td>Indonesia</td>
<td>5</td>
<td>.9996</td>
<td>-0.7747</td>
<td>-5.984</td>
</tr>
<tr>
<td>India</td>
<td>24</td>
<td>.9893</td>
<td>-0.6199</td>
<td>-3.116</td>
</tr>
<tr>
<td>Pakistan</td>
<td>24</td>
<td>.9762</td>
<td>-0.9151</td>
<td>-3.966</td>
</tr>
<tr>
<td>S. Africa</td>
<td>24</td>
<td>.9573</td>
<td>-0.6573</td>
<td>-3.353</td>
</tr>
</tbody>
</table>

In testing the first differenced form of the wholesale price indices under equation (9), the test is for $\gamma = 0$ and the appropriate test statistic is $\tau_\gamma$. The critical values of $\tau_\gamma$ for a 250 observation sample are -2.88 at 95 per cent and -3.46 at the 99 per cent significance levels.
<table>
<thead>
<tr>
<th></th>
<th>(α)</th>
<th>(DU)</th>
<th>(T)</th>
<th>(DT)</th>
<th>(TB)</th>
<th>γ</th>
<th>Adj R²</th>
</tr>
</thead>
<tbody>
<tr>
<td>Japan</td>
<td>.1228</td>
<td>.0160</td>
<td>.0001</td>
<td>.0002</td>
<td>.0016</td>
<td>.9715</td>
<td>.99756</td>
</tr>
<tr>
<td></td>
<td>2.995</td>
<td>2.586</td>
<td>2.505</td>
<td>-2.73</td>
<td>0.292</td>
<td></td>
<td>100.9</td>
</tr>
<tr>
<td>Korea</td>
<td>.2001</td>
<td>.0905</td>
<td>.0010</td>
<td>-.001</td>
<td>-.010</td>
<td>.9317</td>
<td>.99934</td>
</tr>
<tr>
<td></td>
<td>3.628</td>
<td>3.501</td>
<td>3.894</td>
<td>-3.88</td>
<td>-.787</td>
<td></td>
<td>49.39</td>
</tr>
<tr>
<td>Indonesia</td>
<td>.1945</td>
<td>.0719</td>
<td>.0012</td>
<td>-.001</td>
<td>-.014</td>
<td>.9180</td>
<td>.99868</td>
</tr>
<tr>
<td></td>
<td>3.779</td>
<td>2.875</td>
<td>3.258</td>
<td>-3.00</td>
<td>-.521</td>
<td></td>
<td>38.92</td>
</tr>
<tr>
<td>Singapore</td>
<td>.3103</td>
<td>.0621</td>
<td>.0006</td>
<td>-.001</td>
<td>.0081</td>
<td>.9867</td>
<td>.98677</td>
</tr>
<tr>
<td></td>
<td>3.119</td>
<td>2.901</td>
<td>3.572</td>
<td>-3.45</td>
<td>0.517</td>
<td></td>
<td>38.15</td>
</tr>
</tbody>
</table>

The statistics reported are the regression coefficients and related t statistics. The critical values for t(γ) in the additive outlier model are 4.91 at the 99% significance level and we can reject the null hypothesis of a unit root.
Table 4--Unit Root Tests for Foreign Exchange Rates -Base India

<table>
<thead>
<tr>
<th>Country</th>
<th>Lag</th>
<th>Sig of Q</th>
<th>$a_2t$</th>
<th>$\tau_2 (a_{2t})$</th>
<th>$\gamma$</th>
<th>$\tau_2 (\gamma)$</th>
<th>$\Phi_3$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Philippin</td>
<td>15</td>
<td>.9972</td>
<td>.00001</td>
<td>0.108</td>
<td>-.0163</td>
<td>-1.20</td>
<td>1.108</td>
</tr>
<tr>
<td>Pakistan</td>
<td>20</td>
<td>.9874</td>
<td>-.0000</td>
<td>-1.21</td>
<td>-.0266</td>
<td>-1.45</td>
<td>1.526</td>
</tr>
<tr>
<td>S. Africa</td>
<td>15</td>
<td>.9710</td>
<td>.00047</td>
<td>2.544</td>
<td>-.1176</td>
<td>-2.13</td>
<td>3.373</td>
</tr>
<tr>
<td>Thailand</td>
<td>15</td>
<td>.9831</td>
<td>-.0001</td>
<td>-1.74</td>
<td>-.0095</td>
<td>-0.82</td>
<td>1.848</td>
</tr>
</tbody>
</table>

The critical values of $\Phi_3$ for a sample of 250 observations are 6.34 at the 95 per cent level and 8.43 at the 99 per cent significance levels.
<table>
<thead>
<tr>
<th>Country</th>
<th>Lag</th>
<th>Sig of Q</th>
<th>$\gamma$</th>
<th>$\tau_c$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Philippines</td>
<td>14</td>
<td>.9974</td>
<td>-0.8869</td>
<td>-3.687</td>
</tr>
<tr>
<td>Pakistan</td>
<td>20</td>
<td>.9922</td>
<td>-0.9496</td>
<td>-2.923</td>
</tr>
<tr>
<td>S. Africa</td>
<td>15</td>
<td>.9957</td>
<td>-0.9781</td>
<td>-1.085</td>
</tr>
<tr>
<td>Thailand</td>
<td>16</td>
<td>.9529</td>
<td>-1.0973</td>
<td>-4.011</td>
</tr>
</tbody>
</table>

The critical values of $\tau_c$ for a 250 observation sample are -2.88 at 95 per cent and -3.46 at the 99 per cent significance levels.
Table 6--Cointegration Test Results--Thailand and India

<table>
<thead>
<tr>
<th>Form</th>
<th>$\beta_{dp}$</th>
<th>t($\beta_1$)</th>
<th>$\beta_{fp}$</th>
<th>t($\beta_2$)</th>
<th>Lags</th>
<th>Sig of Q</th>
<th>$\tau$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Unrestricted</td>
<td>-1.411</td>
<td>-24.56</td>
<td>1.226</td>
<td>15.06</td>
<td>20</td>
<td>.9813</td>
<td>-1.269</td>
</tr>
<tr>
<td>Restricted</td>
<td>-1.623</td>
<td>-36.60</td>
<td>1.623</td>
<td>36.60</td>
<td>21</td>
<td>.9855</td>
<td>-1.509</td>
</tr>
</tbody>
</table>

Table 7--Cointegration Test Results--Pakistan and India

<table>
<thead>
<tr>
<th>Form</th>
<th>$\beta_{dp}$</th>
<th>t($\beta_1$)</th>
<th>$\beta_{fp}$</th>
<th>t($\beta_2$)</th>
<th>Lags</th>
<th>Sig of Q</th>
<th>$\tau$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Unrestricted</td>
<td>-.0177</td>
<td>-1.559</td>
<td>-.0463</td>
<td>-2.986</td>
<td>20</td>
<td>1.000</td>
<td>-1.566</td>
</tr>
<tr>
<td>Restricted</td>
<td>-.0099</td>
<td>-0.886</td>
<td>0.0099</td>
<td>0.886</td>
<td>20</td>
<td>.999</td>
<td>-1.431</td>
</tr>
</tbody>
</table>

Table 8--Cointegration Test Results--Philippines and India

<table>
<thead>
<tr>
<th>Form</th>
<th>$\beta_{dp}$</th>
<th>t($\beta_1$)</th>
<th>$\beta_{fp}$</th>
<th>t($\beta_2$)</th>
<th>Lags</th>
<th>Sig of Q</th>
<th>$\tau$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Unrestricted</td>
<td>-1.669</td>
<td>-22.34</td>
<td>1.0817</td>
<td>26.81</td>
<td>12</td>
<td>.9817</td>
<td>-3.223</td>
</tr>
<tr>
<td>Restricted</td>
<td>-.4759</td>
<td>-29.87</td>
<td>0.4759</td>
<td>29.87</td>
<td>12</td>
<td>.9799</td>
<td>-1.025</td>
</tr>
</tbody>
</table>

The critical values of $\tau$ in the test of stationary (no constant or time trend) on the residuals from the cointegration regressions are derived from Engle and Yoo. For a 200 observation bivariate sample (the restricted form), the critical values are 3.02 at the 90 per cent level, 3.37 at the 95 per cent level, and 4.00 at the 99 per cent level. For a 200 observation trivariate sample (the unrestricted form), the critical values are 3.47 at the 90 per cent level, 3.78 at the 95 per cent level, and 4.35 at the 99 per cent level.
Table 9--Tests for Stationarity on Real Exchange Rates--Base Japan

<table>
<thead>
<tr>
<th>Country</th>
<th>$\alpha$</th>
<th>$\tau_1(\alpha)$</th>
<th>$\gamma$</th>
<th>$\tau_1(\gamma)$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Korea</td>
<td>-.0235</td>
<td>-.907</td>
<td>-.0142</td>
<td>-.837</td>
</tr>
<tr>
<td>Thailand</td>
<td>.0161</td>
<td>.725</td>
<td>-.0096</td>
<td>-.813</td>
</tr>
<tr>
<td>Philippines</td>
<td>.0969</td>
<td>1.853</td>
<td>-.0538</td>
<td>-1.865</td>
</tr>
<tr>
<td>Australia</td>
<td>.1209</td>
<td>1.452</td>
<td>-.0260</td>
<td>-1.474</td>
</tr>
<tr>
<td>Indonesia</td>
<td>-.0236</td>
<td>-1.113</td>
<td>-.0096</td>
<td>-1.037</td>
</tr>
<tr>
<td>India</td>
<td>-.0218</td>
<td>-1.032</td>
<td>.0067</td>
<td>.779</td>
</tr>
<tr>
<td>Pakistan</td>
<td>.0034</td>
<td>.163</td>
<td>-.0023</td>
<td>-.239</td>
</tr>
<tr>
<td>S. Africa</td>
<td>.0921</td>
<td>1.609</td>
<td>-.0230</td>
<td>-1.641</td>
</tr>
<tr>
<td>Singapore</td>
<td>-.0727</td>
<td>-1.666</td>
<td>-.0165</td>
<td>-1.730</td>
</tr>
</tbody>
</table>

The critical values for the $\tau_1$ statistic for a sample of 250 observations are 2.57 at the 90 per cent level, 2.88 at the 95 per cent level, and 3.46 at the 99 per cent level.
LITERATURE CITED


VITA

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He began his professional career in 1973 as a program analyst with the Department of the Air Force, transferring to Headquarters United States Air Force in 1976. In 1981, he became the Director of Administration for the Defense Logistics Agency (DLA) and currently serves as Assistant to the Deputy Director (Corporate Administration), DLA.