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PRINCIPAL COMPONENTS, STATIONARITY, AND NEW EVIDENCE OF PURCHASING POWER PARITY IN DEVELOPING COUNTRIES

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I. INTRODUCTION

PURCHASING power parity (PPP) is one of the most widely tested economic hypotheses. The argument that prices in different countries move towards equality in common currency terms is of potential interest to policymakers in developing countries (DCs) for at least two reasons. First, PPP becomes a prediction model for exchange rates and a criterion for judging over- and undervaluation of a currency. This may be particularly relevant for small open DCs and those experiencing large inflation differentials between domestic and foreign inflation rates. Second, many exchange rate theories employ some notion of PPP in their construction. Thus the quality of policy advice, insofar as it is based on these theories, may depend on the validity of PPP [Liu and Burkett (1995)]. Evidence on PPP for DCs has led to mixed conclusions regarding its validity [see, inter alia, McNown and Wallace (1989), Liu (1992), Bahmani-Oskooee (1993), Mahdavi and Zhou (1994)]. However, a general view emerges that evidence in favor of PPP is stronger among the high inflation DCs.[1] This study tests for relative PPP in thirty DCs using quarterly data for the period 1973Q2–1997Q3. For this purpose, a new methodology is employed, based on Snell (1996), that tests whether or not the first largest principal component (LPC) based on the growth in their real exchange rates with respect to the U.S. dollar is stationary or not. Using this methodology, PPP is confirmed if the first LPC is stationary.

The recent studies of PPP in DCs have utilized tests for unit roots in real exchange rates and cointegration between various measures of domestic prices and exchange rate–adjusted foreign prices. McNown and Wallace (1989) test for unit roots in U.S. dollar real exchange rates and they employ the Engle-Granger (1987) OLS test for cointegration. Using data on consumer and wholesale prices for the

[1] Studies on PPP for industrial countries have generally provided ambiguous results without a conclusive answer, for example, Balassa (1964) and Hakkio (1984) find in favor of PPP while Dornbusch (1980) and Frenkel (1981) find no evidence in favor of PPP. However, Frenkel (1978) suggests that PPP holds during periods of high inflation.

1970s and 1980s, evidence to support PPP is found in the cases of Argentina, Brazil, Chile, and Israel. Bahmani-Oskooee (1993) uses quarterly data on prices and effective exchange rates for twenty-five DCs for the period 1973–88. Using the same Engle-Granger technique, evidence in favor of PPP among major trading partners is confirmed in only a minority of cases with little evidence to suggest that PPP is more likely in high inflation countries. This finding is supported by Bahmani-Oskooee (1995) who generally rejects the null of stationarity for the real effective exchange rate across a sample of twenty-two DCs. Liu (1992) tests for PPP in a sample of ten Latin American economies using quarterly data from the 1940s and 1950s to 1989. Applying the Johansen (1988) maximum likelihood technique for estimating cointegrating vectors, Liu finds general evidence in favor of PPP with respect to the United States. The advantage of employing the Johansen methodology over the Engle-Granger technique is that the multivariate Johansen procedure is better suited to handling any simultaneity bias that might affect OLS regressions involving the domestic price level, foreign price level, and the exchange rate. Also, the Johansen procedure is able to identify the presence of multiple cointegrating vectors that might exist between these variables. Finally, Mahdavi and Zhou (1994) apply the Johansen technique to investigate PPP in a sample of DCs using quarterly data for 1973Q2 onwards. They conclude that incidences of PPP are more frequently observed among high inflation countries.[2]

Bearing in mind the existing evidence on PPP in DCs, the key reason of interest attached to this particular study is that a new test is applied in the search for PPP. The new technique is an extension of the principal components methodology, based on testing for the stationarity of the first LPC using data on DC growth rates in their real exchange rates with respect to the U.S. dollar. The advantage of this econometric methodology is that, unlike the Johansen maximum likelihood procedure and the Stock and Watson (1988) common trend framework, it does not require the estimation of a complete vector autoregression system (VAR). The size and power of this test is not affected by the VAR being constrained to an unreasonably low order on account of data limitations. This method also avoids the need for an entire sequence of tests for the stationarity of a multivariate system. As indicated by Snell, even if each test in the sequence had a reasonable chance of rejecting the false null, the procedure as a whole is likely to have low power.

This paper is set out as follows. Section II formally describes the empirical methodology, Section III discusses the data set and results, and Section IV concludes.

[2] Further evidence on PPP in DCs based on tests for unit roots and cointegration can be found in Conejo and Shields (1993) and Hoque (1995). While the latter study rejects PPP, Conejo and Shields find evidence in favor of PPP with respect to the United States in the cases of Brazil and Mexico.

II. METHODOLOGY

The methodology involves a test of growth rates in real exchange rates for a sample of thirty DCs. Let P_t^i be the price level in country i where $i = 1, 2, \ldots, n$, P_t^* be the base country price level, and e_t^i be the country i nominal spot price of foreign (base country) currency. Under absolute PPP, we should have $e_t^i = P_t^i/P_t^*$ which implies that the prices of a standard market basket of goods expressed in a common currency are the same. If PPP holds then deviations from absolute PPP should be stationary, in other words the real exchange rate, defined as $e_t^i P_t^*/P_t^i$, should not contain a unit root. Absolute PPP is rather restrictive. The actual exchange rate may deviate from its PPP value on account of imperfections in published price levels (for example, in reality the price indices of different countries do not reflect the same basket of goods). Furthermore, deviations from PPP may occur on account of transport costs, tariffs, and differential speeds of adjustment in the goods and foreign exchange markets. PPP can be redefined in relative terms to allow for any constant of proportion based on these factors that drives a wedge between P_t^i and $e_t^i P_t^*$. Relative PPP states that the percentage change in the nominal spot exchange rate should equal the inflation differential between country i and the base country, i.e., $\dot{e}_t^i = \dot{P}_t^i - \dot{P}_t^*$. Define

$$u_t^i = \dot{e}_t^i - (\dot{P}_t^i - \dot{P}_t^*), \tag{1}$$

where u_t^i denotes the growth in the real exchange rate of country i. The stationarity of u_t^i would suggest that deviations from relative PPP are self-correcting. Thus, relative PPP is confirmed if the u_t^i's across the sample of DCs are stationary.

With a multi-country study there are n deviations, corresponding to the $(n + 1)$ countries in the sample. We construct principal components using each u^i. Let X_t be an $(n \times 1)$ vector of random variables, namely the u_t^i's for each of the n countries, which may be integrated up to order one. The principal components technique addresses the question of how much interdependence there is in the n variables contained in X_t. We can construct n linearly independent principal components which collectively explain all of the variation in X_t where each component is itself a linear combination of the u_t^i's.[3] Since I(1) variables have infinite variances, whereas stationary, I(0), variables have constant variances, it follows that the first LPC, which explains the largest share of the variation in X_t, is the most likely to be I(1) and so corresponds to the notion of a common trend [Stock and Watson (1988)]. However, if the first LPC is I(0) then all the remaining principal components will also be stationary, and there are no common trends which suggests that the u_t^i's contained in X_t are themselves stationary. This will confirm PPP across the sample of n deviations.

[3] See, for example, Child (1970).

More formally, following Stock and Watson (1988) we can argue that each element of X_t may be written as a linear combination of $k \leq n$ independent common trends which are I(1), and $(n - k)$ stationary components which correspond to the set of $(n - k)$ cointegrating vectors among the u_t^i's. The k vector of common trends and $(n - k) \times 1$ vector of stationary components may respectively be written as

$$\tau_t = \alpha' X_t, \tag{2}$$
$$\xi_t = \beta' X_t, \tag{3}$$

where α is an $(n - k)$ matrix of full column rank, β is an $n \times (n - k)$ matrix that forms the $(n - k)$ cointegrating vectors, $\alpha' \alpha = I$ and $\alpha' \beta = 0$. If there are k common trends, it can be shown that the k LPC's of X_t may be written as

$$\tau_t^* = X_t^{*\prime} \alpha^*, \tag{4}$$

where X_t^* is a vector of observations on the u_t^i's in mean deviation form, α^* represents the k eigenvectors corresponding to the largest eigenvalues of X_t and is defined as αR where R is an arbitrary, orthogonal $(k \times k)$ matrix of full rank. This relationship guarantees that under the null hypothesis of k common trends, each of the k LPC's will be I(1). Similarly, for the $(n - k)$ remaining principal components, it can be shown that

$$\xi_t^* = X_t^{*\prime} \beta^*, \tag{5}$$

where β^* corresponds to the $(n - k)$ eigenvectors that provide the $(n - k)$ smallest principal components and is defined as βS where S is an arbitrary orthogonal $(n - k) \times (n - k)$ matrix.

The first LPC will be I(1) provided there is at least one common trend among the u_t^i's contained in X_t. We therefore test the null hypothesis that the first LPC is nonstationary against the alternative hypothesis that the first LPC is I(0). Rejection of the null means that all principal components are stationary and so there are no common trends among the u_t^i's contained in X_t. This confirms relative PPP across the sample. To test the stationarity of the first LPC, we use the familiar Augmented Dickey-Fuller (ADF) test based on

$$\Delta z_t = \rho z_{t-1} + \sum_{i=1}^{p} \gamma_i \Delta z_{t-i} + e_t, \tag{6}$$

where $z = \alpha_1^* X_t^*$ using α_1^* as the first column of α^*, and e_t is a white-noise error term.

III. THE DATA AND RESULTS

The thirty DCs included in the sample are Argentina, Barbados, Brazil, Chile, Columbia, Costa Rica, Ecuador, El Salvador, Ghana, Guatemala, Honduras, India, Indonesia, Israel, Jamaica, Kenya, Mauritius, Mexico, Morocco, Netherlands Antilles, Nigeria, Pakistan, Philippines, Singapore, South Africa, Sri Lanka, Suriname, Thailand, Uruguay, and Venezuela. All price and exchange rate data are taken from the *International Financial Statistics* database. Inflation rates are based on the consumer price index (line 64), and exchange rates are end-of-period spot rates with respect to the U.S. dollar. Quarterly data for the period 1973Q2–1997Q3 provides a sample of size of ninety-eight observations on each series for each country where the use of quarterly data is dictated by data availability across this large sample. The start of 1973 is consistent with Bahmani-Oskooee (1993) and Mahdavi and Zhou (1994) in their investigations of PPP in DCs and can be regarded as the start of the modern "floating rate" period with respect to the U.S. dollar. Deviations from relative PPP for each country are calculated according to equation (1) above.

This sample may be organized in a number of ways. First, we can organize the sample of countries according to inflationary experience and test the hypothesis that PPP is more likely to hold in high-inflation countries. It can be argued that for all countries, real shocks affect PPP, but if these are stationary then they will cancel out.[4] A regime of high inflation, say as a consequence of monetary disturbances, means that the sheer size of price changes dominates the impact of these relative effects, thus the nominal exchange rate follows its PPP path more closely. Furthermore, Copeland (1989) argues that high inflation penalizes agents for maintaining sticky prices, and so attempts to fix the nominal exchange rate may be undermined. If we define a "high inflationary country" as one which experienced an average annual inflation rate in excess of 30 per cent over the sample period 1973Q2–1997Q3, then the "high inflation" countries include Argentina, Brazil, Chile, Ecuador, Ghana, Israel, Mexico, Suriname, and Uruguay while the group of "low inflation" countries comprise Barbados, Columbia, Costa Rica, El Salvador, Guatemala, Honduras, India, Indonesia, Jamaica, Kenya, Mauritius, Morocco, Netherlands Antilles, Nigeria, Pakistan, Philippines, Singapore, South Africa, Sri Lanka, Thailand, and Venezuela.

The second method of organizing the sample is according to region. This provides five groupings: Africa—Ghana, Kenya, Mauritius, Morocco, Nigeria, and South Africa; Asia—India, Indonesia, Israel, Pakistan, Philippines, Singapore, Sri Lanka, and Thailand; Central America—Costa Rica, El Salvador, Guatemala, Honduras, and Mexico; South America—Argentina, Brazil, Chile, Columbia, Ecuador, Suriname, Uruguay, and Venezuela; and Other—Barbados, Jamaica, and Netherlands Antilles.

The principal components results for the full sample of countries along with the inflationary and regional groupings are reported in Table I. The first LPC for X_t which is an $(n \times 1)$ vector of u_t^i's offers the largest explanation of the variation in the u_t^i's. The greater is the explanatory power of the first LPC, then the more closely do deviations from relative PPP move together over time. The explanatory power of

[4] An exception may be productivity shocks as described by Balassa (1964).

TABLE I

PRINCIPAL COMPONENTS BASED ON GROWTH IN REAL EXCHANGE RATES

	Eigenvalue	Cumulative R^2	Sample Size (n)
All countries	5.609$^{\#}$	0.189	30
High-inflation countries	1.882$^{\#}$	0.209	9
Low-inflation countries	5.039$^{\#}$	0.240	21
Africa	2.624	0.437	6
Asia	2.533$^{\#}$	0.317	8
Central America	1.726$^{\#}$	0.345	5
South America	1.716	0.214	8
Other	1.847$^{\#}$	0.616	3

Notes: 1. Estimation is for the period 1973Q2–1997Q3. Growth in real exchange rates, or deviations from relative PPP, are with respect to the United States. The full sample of countries comprises Argentina, Barbados, Brazil, Chile, Columbia, Costa Rica, Ecuador, El Salvador, Ghana, Guatemala, Honduras, India, Indonesia, Israel, Jamaica, Kenya, Mauritius, Mexico, Morocco, Netherlands Antilles, Nigeria, Pakistan, Philippines, Singapore, South Africa, Sri Lanka, Suriname, Thailand, Uruguay, and Venezuela.

2. $^{\#}$ indicates stationarity at the 5 per cent significance level or better of the first largest principal component (LPC) as reported in Table II.

the first LPC can be measured by its eigenvalue or the cumulative R^2 (measured as the eigenvalue divided by the number of countries in that particular group).[5] If we refer to Table I, it can be seen that the variation in the u_t^i's explained by the first LPC is modest, for example only 18.9 per cent for the full sample of countries. However, this figure varies across the regional groupings with 43.7 and 61.6 per cent in the cases of the African and Other groups respectively as opposed to 21.4 per cent in the case of the South American group. These results suggest that there may be regional factors that influence the extent to which deviations from relative PPP move together. In the case of the high- and low-inflation groups, the cumulative R^2's are comparable. Of course, the extent to which deviations from relative PPP are synchronized across countries does not necessarily imply that PPP holds, thus we now need to test the stationarity of the first LPC.

Table II reports the ADF unit root tests on the first LPC's. At the 5 per cent significance level, the first LPC is confirmed as stationary in the majority of cases, thereby suggesting that relative PPP holds across the sample. In particular, this conclusions holds when the full sample is estimated together irrespective of whether high- and low-inflation DCs are considered. Indeed, in both these cases the null of nonstationary is strongly rejected at the 1 per cent significance level. We therefore support the earlier viewpoint held by McNown and Wallace (1989), Liu (1992),

[5] Since the n components that explain the variation in X_t are orthogonal to each other, it must be the case that the sum of their respective contributions equal unity.

TABLE II

ADF UNIT ROOT TESTS ON THE FIRST LPC

	ADF Statistic	Lag
All countries	-3.139^{**}	5
High-inflation countries	-3.774^{***}	8
Low-inflation countries	-3.603^{***}	5
Africa	-2.582^{*}	8
Asia	-3.249^{**}	4
Central America	-4.006^{***}	8
South America	-2.230	4
Other	-3.156^{***}	8

Notes: 1. The lag lengths are chosen to ensure white noise residuals. Following the application of the Schwarz Information Criteria, all regressions exclude a time trend. Further tests based on Dickey and Fuller (1981, Tables I–IV) revealed the time trend to be insignificant.

2. ***, **, and * indicate rejection of the null of nonstationarity at the 1 per cent, 5 per cent, and 10 per cent significance levels with critical values taken from Fuller (1976).

and Mahdavi and Zhou (1994) that PPP is likely to hold in the case of high-inflation countries. However, we also have strong evidence that relative PPP also holds for low-inflation countries. One source of explanation for this difference in conclusions lies in the methodology. The earlier tests for cointegration are subject to low test power which makes rejection of the null of non-cointegration unlikely. A further finding from Table II is that PPP might be a regional phenomenon. The first LPC is clearly nonstationary in the case of South America, and the null of nonstationarity is only rejected at the 10 per cent significance level in the case of Africa.

Table III reports the factor loadings applying to the first LPC for all countries, high inflation countries, low inflation countries and the regional groups. These factor loadings are the squared coefficients of correlation between the u_t^i's and the first LPC. Perfect synchronization of deviations from relative PPP would require factor loadings of unity attached to the first LPC across all countries. Factor loadings that are insignificant or low on the part of individual countries imply some degree of independence from the rest of the sample. The factor loadings on LPC^a, which is the first LPC for all countries, are noticeably lower for many Latin American countries as compared to the rest of the sample. This suggests that Latin American deviations from relative PPP are generally less synchronized with the rest of the world. In the case of the first LPC applicable to high-inflation countries (LPC^b), the different signs on some factor loadings suggest that Brazil, Ghana, and Israel have had quite different experiences with regard to deviations from relative PPP as compared to Mexico, Suriname, and Uruguay. This can be contrasted with the factor loadings

TABLE III

FACTOR LOADINGS ATTACHED TO THE FIRST LPC

	LPC^a	LPC^b	LPC^c	LPC^d	LPC^e	LPC^f	LPC^g	LPC^h
Argentina	0.102	0.050					0.278***	
Barbados	0.485***		0.567***					0.400***
Brazil	0.128	0.240**					0.248**	
Chile	–0.312***	–0.031					–0.008	
Columbia	0.396***		0.380***				0.768***	
Costa Rica	–0.650***		–0.690***			–0.035		
Ecuador	0.065	–0.130					0.505***	
El Salvador	–0.180		–0.128			0.262***		
Ghana	0.406***	0.698***		0.358***				
Guatemala	0.000		0.024			0.968***		
Honduras	–0.037		–0.006			–0.040		
India	0.447***		0.474***		0.937***			
Indonesia	0.657***		0.679***		0.003			
Israel	0.483***	0.788***			–0.004			
Jamaica	0.447***		0.375***					0.114
Kenya	0.475***		0.492***	0.619***				
Mauritius	0.772***		0.756***	0.846***				
Mexico	–0.079	–0.701***				0.093		
Morocco	0.604***		0.573***	0.830***				
Netherlands Antilles	0.514***		0.582***					0.886***
Nigeria	–0.010		0.020	–0.321***				
Pakistan	0.734***		0.759***		0.206**			
Philippines	0.522***		0.502***		0.212**			
Singapore	0.291***		0.294***		–0.003			
South Africa	0.559***		0.532***	0.778***				
Sri Lanka	–0.002		0.034		–0.055			
Suriname	–0.241***	–0.385***					–0.633***	
Thailand	0.707***		0.731***		0.178			
Uruguay	0.351***	–0.239**					0.356***	
Venezuela	0.237**		0.195**				0.453***	

Notes: 1. Factor loadings are for the first LPC reported in Tables I and II. LPC^a is the first LPC for the full sample of countries, LPC^b applies to the high-inflation countries, LPC^c applies to the low-inflation countries, LPC^d applies to the group of African countries, LPC^e applies to the Asian countries, LPC^f applies to Central American countries, LPC^g applies so South American countries, and LPC^h applies to Other countries.

2. *** and ** indicate significance of the factor loadings at the 1 per cent and 5 per cent levels based on Pearson correlation coefficients [see Child (1970)].

on the first LPC for the low-inflation countries (LPC^c) which are all positive with the exception of Costa Rica. The first LPCs for the regional groups also presents a picture of diverse experience among DCs. In particular, where PPP is confirmed in the cases of Asia (LPC^e), Central America (LPC^f), and Other (LPC^h), the reported factor loadings are all non-negative. However, where non-stationarity of the first

LPC is accepted in the cases Africa (LPC^d) and South America (LPC^g), experiences are more diverse with negative factor loadings in the cases of Nigeria and Suriname.

IV. SUMMARY AND CONCLUSION

This study has tested for relative PPP among a sample of thirty DCs using a new econometric technique that investigates the stationarity of the largest principal component based on deviations from relative PPP against the United States. This technique has advantages over existing studies that employed Engle-Granger and Johansen techniques that can suffer from low test power as a result of demands on limited data which makes rejection of the null of a nonstationary real exchange rate or the null of no cointegration between domestic prices, foreign prices, and nominal exchange rates unlikely. Using quarterly data for 1973–97, PPP is generally confirmed and, unlike earlier studies, there is no evidence that PPP is confined to high-inflation developing countries.

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