The International Journal of Banking and Finance 2 (2004) 29-43

PURCHASING POWER PARITY IN DEVELOPING COUNTRIES: EVIDENCE FROM CONVENTIONAL AND FRACTIONAL COINTEGRATION TESTS

A. C. Arize *, John Malindretos, Elias C. Grivoyannis****

*College of Business and Technology, Texas A & M University U.S.A.
**Yeshiva University, New York, NY 10033 - 3201, U.S.A.

Abstract

This paper examines the long-run validity of purchasing power parity (PPP) for fourteen developing countries. The period examined is 1973:4 through 2002:8. The methods of Elliot, Rothemberg and Stock (1996), Kwiattkowski et al. (1992) and Geweke and Porter-Hudak (1983) are employed to detect the time series properties of exchange rates and consumer price indices of these countries. We find that these variables are nonstationary. We then utilize these data to test the PPP using both conventional and fractional approaches. Estimates of the cointegrating relations are obtained using estimators suggested by Stock and Watson (1993) and Phillips and Hanson (1990), respectively. The results are consistent with the argument that, during the recent floating exchange-rate period, PPP holds well, at least in a weak form , in developing countries where the general price level movements overshadow the factors causing deviations from the PPP.

1. Introduction

One area in international finance that still attracts the attention of many researchers is the empirical validity of the Purchasing Power Parity theory (henceforth, PPP) as an explanation for the long-run behavior of exchange rates. As new techniques for studying time series data have rapidly developed, it has become natural to use such statistical tools in studying whether old theories such as PPP are empirically valid. Testing for the validity of PPP continues to be important because it is the foundation of many long-run exchange rate models, and it generally serves as a guide to financial authorities when they intervene in the foreign exchange market to ensure that the level of exchange rate is consistent with PPP or when they compare real purchasing power between countries for policy polices.

*The author would like to thank Ed Manton, Keith McFarland and Lee Schmidt for helpful comments on an earlier draft. Special thanks to Kathleen Smith for excellent research assistance. This research is funded by a GSRF-TAMU-C grant.

The PPP theory postulates a stable long-run relationship between the exchange rate of two countries and the price levels of the involved countries. The underlying intuition—that international arbitrage equalizes prices across countries (that is, the law of one price)—makes it an attractive theoretical tool. In view of PPP's theoretical appeal, it is not surprising that a great deal of empirical testing has been conducted using data for industrial countries and, recently, the testing approach has utilized the conventional cointegration methods of Engle and Granger (1987) and the fractional cointegration method of Granger and Joyeux (1980). Interestingly, most of the empirical research has concluded that PPP fails to hold in the short run. Whether PPP holds as a long-run equilibrium relationship remains a controversial issue.

Researchers have expended great effort in attempting to investigate empirically the long- run PPP, but with mixed results. On the one hand, Herwartz and Reimers (2002), Cheung and Lai (1998), Pappel (1997), Pippenger (1993), Cheung and Lai (1993), Johansen and Juselius (1992) and Frankel (1978) are examples of studies suggesting that PPP is valid in the long run. A fail-to-reject decision concerning PPP (i.e., PPP holds) implies that the nominal exchange rate, when adjusted for inflation differentials must be stationary or constant. Most studies attribute the efficacy of PPP to the economic conditions in industrial countries, which are generally similar in structure and macroeconomic policies, and to the high degree of interdependence, as well as mobility in capital, goods and services.[1]

On the other hand, Coe and Serletis (2002), Crowder (1996), Flynn and Boucher (1993) and Chowdhury and Sdogati (1993) are examples of studies that find no support for long-run PPP. Rejection of the PPP hypothesis (i.e., PPP fails) implies that exchange rate follows a random-walk process and that it does not depend on relative prices. A possible explanation for the failure of PPP in industrial countries may be associated with the fact that, although inflation has been relatively stable across countries, real shocks have dominated nominal shocks in shaping the behavior of exchange rates, resulting in a divergence between the paths followed by exchange rates and those followed by prices. Some have suggested that the failure to find a cointegrating relation between nominal exchange rate and prices may be due to the econometric methods employed in previous studies. Clearly, such remarkably mixed evidence for developed countries is unfortunate in the light of the profound implications that PPP may have on economic policy. But little is known about the extent to which this conclusion may be confirmed for less developed countries.

Developing countries have significant dissimilarity with industrial countries. This is so because they are apt to experience factors that could affect the structural stability

[1]Macdonald (1995), in a survey of the PPP evidence, has concluded that there is considerable evidence supportive of a long-run relationship between relative prices and exchange rates, in the sense that these variables are cointegrated. However, the long-run relationship does not conform to traditional PPP since the hypotheses of proportionality and symmetry are generally rejected. He argues that these restrictions may not hold in the recent floating era because (a) measurement errors may exist in the price indices employed in these tests, (b) econometric methods employed may produce coefficient biases in small samples and (c) real disturbances and capital movements may have upset these restrictions.

of a long-run equilibrium relationship between nominal exchange rates and domestic and foreign price indices. Some of these factors are high inflation rates, volatile exchange rates, rapid monetary growth, trade controls, inefficient markets, large structural (real) shocks and fiscal as well as budget imbalances. Studies of less developed countries (or LDCs for short) experience have been few due to the unavailability of sufficient time series data, and no clear pattern of results or conclusions has emerged from such studies.

Bahmani-Oskooee (1993) uses quarterly data from 1973-1988 and obtains mixed results about the existence of PPP in twenty-five developing countries. There was no support for PPP when effective exchange rates were used in testing for PPP, but he found support for PPP in seven out of the twenty-five countries when bilateral exchange rates were used. Nagayasu (1998) study uses panel cointegration tests on annual data for black market exchange rates and the consumer price index in sixteen countries from 1981-1994 and found support for the long-run PPP. However, no empirical support for the presence of a long-run PPP relation was obtained when cointegration tests were applied to the annual data for individual countries. Holmes (2000) found support for PPP when he applied panel unit root tests to quarterly data for selected African countries from 1974-1997. However, he found no support for PPP when unit root tests were applied to individual country data during the same period.

One problem with existing studies of developing countries is that they have presumed (either explicitly or implicitly) that the PPP statistical relationship (that is, the relationship between nominal exchange rates and price indices) is stable. It is possible that this may not be the case. There is no reason to believe a priori that, over the past twenty-nine years, the relative importance of factors influencing the PPP function in developing countries has remained unchanged. Nevertheless, as Hansen (1992) points out, it is a standard practice in econometrics to perform tests for parameter constancy on regression models, because the estimated parameters of a time series may be different over time and tests are important, since parameter non-constancy is indicative of model misspecification. In a similar vein, Ghysels and Hall (2000) point out that "structural instability is another form of nonstationarity ... and neglected structural instability can bias inferences about other aspects of the model."

A second problem with existing studies of developing countries is that they have employed tests that have difficulty distinguishing unit-root processes from stationary processes with substantial persistence. In other words, tests for purchasing power parity have relatively low power under both the null hypothesis that it holds and the null that it fails; the mixed evidence concerning PPP may have resulted from the nature of these tests.

The aim of this paper is to empirically investigate the hypothesis of long-run PPP over the generalized floating exchange rate period for fourteen developing countries — Cote d'Ivoire, India, Kenya, Korea, Madagascar, Malaysia, Mauritius, Morocco, Niger, the Philippines, Rwanda, Singapore, Sri Lanka and Thailand — in the context of conventional and fractional cointegration approaches. Our sample consists of seven Asian countries and seven African countries, and hence, may allow comparison of results. Monthly data are used for the period 1973:4 through 2002:8 (i.e., 357 observations).

This paper extends and improves on previous studies in several ways. First, it examines the time series properties of the variables by allowing for the possibility that the order of integration of nominal exchange rate and price indices may be fractional I(d) instead of integer. Conventional unit root tests typically have considered only integer values of zero or zero or some greater integer. Theoretically, the value of the order of differencing (d) is not limited to an integer. The tests used are: (a) the DF-GLS of Elliot, Rothenberg and Stock (1996), in which the null hypothesis is a unit root (nonstationarity) against the alternative hypothesis of I(0); (b) the KPSS due to Kwiatkowski et al. (1992), in which the null hypothesis of level (trend) stationarity is tested against the alternative of nonstationarity; and (c) the GPH test of Geweke and Porter-Hudak (1983), where the null hypothesis of d=1 is evaluated against the alternative of d<1 and the null hypothesis of d=0 is tested against the alternative hypothesis of d>0.

Second, the analysis uses the conventional cointegration methodology of Johansen (1995) to determine whether the relevant variables share a common stochastic trend. As a cross-check of the Johansen test results, which use the null hypothesis of no cointegration, we also test the null hypothesis of cointegration by means of cointegration tests suggested by Hansen (1992) and Harris and Inder (1994). Following Haug and Lucas (1996), estimates of the cointegrating relations are obtained using the dynamic ordinary least Squares (DOLS) estimator of Stock and Watson (1993)[2] and the fully modified ordinary least Squares (FMOLS) estimator of Phillips and Hansen (1990).

Third, the PPP plays a pivotal role in many theoretical and empirical models of exchange-rate determination because it identifies national prices as primary determinants of the equilibrium exchange rates, therefore, the issue of parameter stability is important if the long-run equilibrium relation is to be useful. in long-term planning, forecasting and policy formulation. For example, knowledge of the degree to which the PPP relation is structurally stable is important for the design of both exchange rate and monetary policies. If the PPP relation is structurally unstable, trade adjustment programs in developing countries that have strongly emphasized a relationship between inflation and exchange rate could be unsuccessful if the PPP relation has broken down. In addition, the intended effect of an monetary policy may be doomed by a breakdown of the PPP relation because market participants can no longer presume that any inflationary effects will be offset by exchange rate effects; this could precipitate a balance-of-payment crisis. In this study, the structural stability of the cointegrating relation is evaluated using the three tests suggested by Hansen (1992) -- Lc, M_F and the SupF test statistics.

[2]As Hoffman et al. (1995) point out, a potential problem of multicollinearity may arise in the use of the Johansen estimator, since domestic and foreign price indices may be collinear.

Finally, this paper relaxes the assumption in the conventional cointegration analyses that the cointegrating residual must be integrated of order zero (i.e., I(0)) before the null hypothesis of non-cointegration can be rejected. The strict nature of the I(0) assumption may lead to invalid conclusions since it may not be adequate in assessing each country's data for the existence of a long-run equilibrium relationship. In this study, we employ tests for fractional cointegration, and the two estimators used are the GPH test of Geweke and Porter-Hudak (1983) and the exact maximum likelihood (EML) of Sowell (1992) and others. This generalized form of cointegration enables us to capture slow or subtle mean-reverting dynamics because the integration order of the cointegrating residual is allowed to take any value on the real line. Therefore, it enables the capture of a long-run equilibrium relationship, that is mean reverting though stationary but not exactly I(0).

2. Model specification

From an empirical standpoint, there is a consensus in the literature that the long-run equilibrium PPP relation may be written as:

$$e_t = \beta_0 + \beta_1 p_t + \beta_2 p_t^* + \mu_t \tag{1}$$

where p_t and p_t* are the logarithm of domestic and foreign price indices. e_t is the logarithm of nominal exchange rate, defined as the bilateral exchange rate of domestic currency against the U.S. dollar measured relative to the base year of the price indices in order to eliminate unit dependency. The error term is $:_t$,and the t subscript denotes time.

The PPP hypothesis is based upon the presumption that, in an integrated competitive market, the law of one price would prevail and, as such, the price of a given good should be the same when quoted in different currencies. However, in the presence of trade and transactions restrictions, the sum of estimated coefficients on domestic and foreign price indices should be positive but not equal to unity. Aggregate price indices are usually used when testing for long-run PPP, but these indices have significant measurement errors and would not confirm their theoretical counterparts.[3] For this reason, the symmetry ($\$_1=-\$_2$) and proportionality ($\$_1=\$_2=1$) conditions are not imposed in equation (1). Nevertheless, the concept of cointegration developed in Engle and Granger (1987) is attractive because we can relax these restrictions and still test the equilibrium relation among the variables in equation (1); for more on this, see Johansen and Juselius (1992) and Cheung and Lai (1993).

[1]Empirical studies concerning PPP theory do not use domestic and foreign price levels because these series are not easily available; they use price indices as proxies. The use of an aggregative price index may understate the true elasticity since goods with relatively low elasticities may exhibit the largest variation in price and the effect of such variation will be more than its relative importance in the construction of the aggregative price index (Leamer

The basic idea of cointegration is that two or more nonstationary time series may be regarded as defining a long-run equilibrium relationship if a linear combination of the variables in the model is stationary (converges to an equilibrium over time). Thus, if the PPP relation describes a stationary long-run relationship among the variables in equation (1), this can be interpreted to mean that the stochastic trend in nominal exchange rate is related to the stochastic trends in domestic and foreign price indices. In other words, even though deviations from equilibrium should occur, they are mean reverting (Arize, 1996).

3. Empirical results

A. The Data

The data used in this study consist of monthly data on nominal exchange rate of seven Asian countries (India, Korea, Malaysia, the Philippines, Singapore, Sri Lanka and Thailand) and seven African countries (Cote d'Ivoire, Kenya, Madagascar, Mauritius, Morocco, Niger, and Rwanda). The exchange rates used are bilateral exchange rates of these countries against the US dollar. The price index used is the consumer price index (CPI). All the data are taken from the CD-ROM version of the International Financial Statistics of the International Monetary Fund (December 2003).

B. Unit Root Test Results

Because cointegration tests require a certain stochastic structure of the time series involved, the first step in the estimation procedure is to determine whether the variables are stationary or nonstationary in levels. The tests employed for such empirical investigation fall into three key classes. Tests of the null hypothesis of a unit root, H_0: d=1, tests for the null hypothesis of stationarity, H_0: d=0, and tests for long memory, H_0: 0<d<1.

The four tests employed here are the DF-GLS of Elliot et al. (1996), the KPSS (Kwiatkowski et al. (1992), the LR (Lobato and Robinson, 1998) and the GPH (Geweke and Porter-Hudak, 1983). The first test has a null hypothesis of nonstationarity (i.e., differencing order of one, d=1), whereas the null hypothesis of stationarity (i.e., differencing order of zero, d=0) is assumed for LR and KPSS tests. Unlike standard unit root tests, the GPH test avoids the knife-edged I(1) and I(0) distinction and considers the interesting alternative that the effects of shocks have no permanent effect, that is, the process is mean reverting. Hence, we test the null hypothesis of d=1 against the alternative of d<1, and the null hypothesis of d=0 is tested against the alternative of d>0.

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and Stern, 1970). Also, price indices are expressed in relation to a base period, and it may not be possible to know whether the PPP was valid in the base period. Further, domestic and foreign price indices are never based on the same basket of goods. This may introduce a nonstationarity that could lead to a false rejection of PPP as a long-run equilibrium condition when cointegration tests are used. Finally, most indices include trade and non-tradable good and services; however, the latter are determined mostly by domestic factors.

Table 1
Time series properties

	DF - GLS		KPSS				Lobato-Robinson Test		Geweke and Porter-Hudak [GPH (1983)]					
			W / O Trend		Trend				Price			E.Rate		
	Price	E.Rate	Price	E.Rate	Price	E.Rate	Price	E.Rate	d	H_0:d=1	H_0:d=0	d	H_0:d=1	H_0:d=0
Asian Countries														
India	-2.972(10)	-2.166(1)	7.27	7.08	0.52	1.19	-7.518	-7.506	0.945	-1.776*	30.731*	0.990	-0.203	19.905*
Korea	-2.321(10)	-2.577(8)	6.62	1.36	5.56	0.58	-7.293	-7.430	0.996	-0.194	44.382*	1.082	0.517	6.833*
Malaysia	-2.891(6)	-2.497(1)	7.02	0.90	4.52	0.84	-7.493	-6.804	0.926	-4.489*	56.178*	0.916	-0.555	6.017*
Philippines	-1.292(9)	-2.609(8)	7.14	1.40	6.86	0.52	-7.179	-7.189	0.974	-1.165	43.112*	1.040	0.608	15.664*
Singapore	-2.214(6)	-1.501(2)	6.75	1.16	5.89	0.60	-7.492	-7.304	0.820	-6.159*	28.076*	0.990	-0.052	4.982*
Sri Lanka	-2.700(6)	-2.408(7)	7.25	6.84	0.45	1.13	-7.412	-7.543	0.975	-1.022	39.774*	1.010	0.148	14.903*
Thailand	-2.604(6)	-2.354(7)	6.90	5.33	1.13	0.72	-7.427	-7.456	0.935	-2.873*	41.294*	1.097	0.541	6.101*
African Countries														
Cote d Ivoire	-3.856(0)	-2.218(1)	6.74	5.13	1.13	0.55	7.286	-7.460	1.036	0.794	22.730*	1.132	0.512	4.381*
Kenya	-1.531(3)	-2.743(7)	7.20	7.10	0.49	0.66	-7.284	-7.364	1.013	0.522	41.194*	1.060	0.834	14.667*
Madagascar	-2.166(3)	-2.714(1)	7.23	7.17	0.31	0.62	-7.351	-7.398	0.996	-0.122	32.913*	0.959	-0.672	15.593*
Mauritius	-3.908(2)	-2.295(3)	6.95	6.76	1.31	0.62	-7.126	-7.023	0.957	-1.775*	39.402*	0.963	-0.635	16.542*
Morocco	-1.603(10)	-1.472(1)	6.96	5.86	1.68	0.93	-7.451	-7.474	0.979	-1.282	60.488*	1.043	0.484	11.723*
Niger	-2.433(1)	-2.218(1)	5.12	5.13	0.96	0.55	-7.404	-7.286	1.024	0.358	15.373*	1.132	0.512	4.381*
Rwanda	-2.327(4)	-1.556(10)	6.82	0.72	5.47	1.45	-7.367	-7.221	0.993	-0.284	37.848*	0.916	-0.525	5.729*
United States	-1.694(10)		6.82		1.61		-7.436		0.982	-1.361	85.913*			

Notes: Price is consumer price index. The DF-GLS critical value is —2.01 at the 5 percent level, and KPSS statistics are evaluated at lag four All Lobato-Robinson test values have probability values below 0.05. The critical value for the GPH is 1.4 at the 5 percent level.

Table 1 reports the results for the four tests. Because it is possible for different tests to yield conflicting conclusions, we designate a series as a nonstationary variable if the null hypothesis of nonstationarity is not rejected by three of the four tests or if the nonstationarity threshold of d$0.5 is satisfied in case of the GPH test. The DF-GLS statistic indicates that the null hypothesis of nonstationarity cannot be rejected in all cases, except for the consumer price index of Cote d'Ivoire. The KPSS statistic rejects the null hypothesis of I(0) in all cases. Likewise, the Lobato and Robinson (1998) test statistic rejects the null hypothesis of I(0) and finds evidence consistent with long memory. The GPH test rejects the null hypothesis of short memory in all cases; however, there is evidence of long memory for the consumer price indices of India, Malaysia, Singapore, Thailand and Mauritius. The results in Table 1 suggest that, while it is reasonable to carry an empirical investigation of the PPP relation predicated on nonstationary processes, some caution is neccessary.

C. Conventional Cointegration Analysis

Having accepted the combined univariate test results as evidence that the variables can be treated as nonstationary for empirical analysis, the next step is to apply conventional cointegration tests to examine whether nominal exchange rate, domestic and foreign price indices are cointegrated, as given by equation (1). The results of applying these tests are summarized in Table 2.

Table 2

Cointegration analysis test results

Countries		Johansen						Harris-Inder		
		Maximum Eigenvalue			Trace Statistics			KPSS Tests		
	Lags H_0:	$r=0$	$r\leq 1$	$r\leq 2$	$r=0$	$r\leq 1$	$r\leq 2$			
	H_a:	$r=1$	$r=2$	$r=3$	$r\geq 1$	$r\geq 2$	$r\geq 3$	0 lags	3 lags	6 lags
Asian Countries										
India	4	28.84	11.98	3.87	44.69	15.85	3.87	0.068	0.044	0.041
Korea	7	26.26	12.57	4.66	43.49	17.22	4.66	0.063	0.041	0.044
Malaysia	615.36	6.82	1.69	23.87	8.51	1.69	0.091	0.066	0.062	
Philippines	7	14.59	10.66	6.01	31.26	16.67	6.01	0.088	0.069	0.063
Singapore	9	33.66	13.29	3.21	50.16	16.49	3.21	0.187	0.141	0.150
Sri Lanka	8	29.39	11.42	2.12	42.94	13.55	2.12	0.046	0.024	0.022
Thailand	7	13.84	11.51	4.17	29.52	15.68	4.17	0.073	0.055	0.053
African Countries										
Cote d' Ivoire	4	40.47	7.24	3.53	51.24	10.77	3.53	0.054	0.050	0.051
Kenya	7	25.58	10.65	1.95	38.17	12.59	1.95	0.069	0.054	0.052
Madagascar	5	18.44	7.34	2.93	28.72	10.27	2.93	0.040	0.033	0.040
Mauritius	7	23.49	9.45	2.65	35.60	12.11	2.65	0.085	0.061	0.066
Morocco	12	24.71	9.32	8.21	42.24	17.53	8.21	0.162	0.112	0.108
Niger	5	20.99	8.08	3.19	32.27	11.27	3.19	0.033	0.035	0.037
Rwanda	7	27.78	13.75	0.36	41.88	14.11	0.36	0.033	0.022	0.022
Critical Values 5%		21.12	14.88	8.07	31.54	17.86	8.07			

Note: r denotes the number of cointegrating vectors and Johansen's critical values are for the 5% level of significance. For the Harris and Inder test, the critical values at the 5 percent level is 0.3202 and 0.2335 at the 10 percent level

The first panel of Table 2 presents the results of applying the Johansen (1995) system- based cointegration method to test for the presence or absence of long-run equilibria among exchange rate, domestic and foreign price indices. The test utilizes two likelihood-ratio (LR) test statistics for the number of cointegrating vectors: namely, the maximum eigenvalue (8-max) and the trace statistics. A constant was allowed to enter the VAR unrestrictedly and this permits intercept in the cointegrating relations ans a drict in the nonstationary part of the process. The lag order of the Vector Autoregressive (VAR) was determined by Sim's likelihood ratio (SLR) test and the F-version of the Breusch and Godfrey (1981) statistic for residual autocorrelation.

Focusing on the calculated 8-max test results, it is observed that the null hypothesis of no cointegration (H_0:r=0) is rejected in nine of the fourteen cases. The exceptions are Madagascar, Malaysia, Niger, the Philippines and Thailand. For example, for India, the calculated 8-max test statistic of 28.84 is larger than its critical value of 21.12 from Pesaran and Pesaran (1997) at the 5 percent level. For the trace test results, the null hypothesis of no cointegration is rejected in ten of the fourteen countries, but it is not rejected in Madagascar, Malaysia, the Philippines and Thailand. Both 8-max and trace statistics reject the null hypothesis of no cointegration in Madagascar, Malaysia, the Philippines and Thailand. It is worth pointing out that there is a single cointegration vector in those cases where the null hypothesis of r=0 can be rejected since the null hypothesis of at most one cointegrating (H_0:r#1) is in no case rejected.

Taken together, the test results indicate that, in most of the countries studied, the nominal exchange rate is cointegrated with domestic and foreign price indices. The results of implementing the Harris and Inder (1994) test are reported in the second section of Table 2. As can be seen from the results, the null hypothesis of cointegration cannot be rejected in each of the fourteen countries at the conventional levels. The critical value is 0.3202 at the 5 percent level and 0.2335 at the 10 percent level.

Armed with evidence that a single cointegrating vector prevails among the three-variable menu of nominal exchange rate, domestic and foreign price indices, we follow the specification suggestions by Huag and Lucas (1996), Hoffman, Rasche and Tieslau (1995) and Stock and Watson (1993) and obtain estimates of the cointegrating vector by applying the dynamic ordinary least squares (DOLS) and the fully modified ordinary least squares (FMOLS). These estimates are reported in Table 3.

Without discussing each estimator in detail, we note that the t-ratios are all significantly different from zero at the 5 percent level. Also, observe that, because of measurement errors, the symmetry condition that $\$_1$ is equal to -$\$_2$ is rejected in the FMOLS (M) results for all countries except Korea, the Philippines, Cote d'Ivoire, Mauritius and Rwanda. Nevertheless, following the lead of Cheung and Lai (1993) and Macdonald (1995), a finding of cointegration among exchange rate and price indices without the constraints imposed is consistent with long-run PPP with the presence of measurement errors in prices indexes and unequal weights attached to the same good in different indexes.

Table 3
Long-run elasticities

Countries	Stock and Watson			Phillips and Hansen			Hansen's Stability Tests		
	β_1	β_2	ϕ	β_1	β_2	ϕ	L_C	M_F	SupF
Asian Countries									
India	1.40	-0.92	108.32	1.38	-0.87	131.39	0.20	2.43	6.86
	[32.23]	[11.08]	[0.000]	[26.04]	[9.36]	[0.000]	[0.200]	[0.200]	[0.200]
Korea	0.86	-0.83	0.032	0.79	-0.61	0.77	0.29	3.22	5.79
	[3.52]	[2.05]	[0.858]	[2.54]	[1.20]	[0.380]	[0.200]	[0.200]	[0.200]
Malaysia	1.88	-1.15	72.68	1.86	-1.13	123.47	0.22	4.45	11.06
	[7.32]	[6.29]	[0.000]	[7.73]	[6.11]	[0.000]	[0.200]	[0.163]	[0.200]
Philippines	1.29	-1.45	11.79	1.28	-1.40	0.95	0.69	34.22	108.36
	[24.58]	[15.05]	[0.001]	[14.77]	[6.91]	[0.329]	[0.036]	[0.010]	[0.010]
Singapore	-1.53	0.39	23.27	-1.032	0.17	21.26	1.00	18.95	49.33
	[-3.21]	[1.59]	[0.000]	[2.63]	[0.80]	[0.000]	[0.010]	[0.010]	[0.010]
Sri Lanka	0.23	1.47	616.56	0.25	1.37	417.16	0.18	3.02	11.21
	[4.72]	[13.33]	[0.000]	[4.23]	[10.11]	[0.000]	[0.200]	[0.200]	[0.200]
Thailand	1.96	-1.82	10.22	1.92	-1.72	15.48	0.11	2.21	4.45
	[7.46]	[6.56]	[0.001]	[6.57]	[5.23]	[0.000]	[0.200]	[0.200]	[0.200]
African Countries									
Cote d' Ivoire	1.78	-2.20	6.210	1.48	-1.53	0.09	0.56	8.64	14.71
	[6.57]	[5.17]	[0.013]	[5.01]	[3.41]	[0.768]	[0.138]	[0.027]	[0.123]
Kenya	1.00	-0.87	3.173	1.00	-0.77	8.72	0.24	4.87	18.23
	[26.92]	[8.05]	[0.075]	[24.07]	[6.59]	[0.003]	[0.200]	[0.124]	[0.012]
Madagascar	1.26	-1.01	4.216	1.24	-9.40	5.30	0.47	7.65	14.63
	[21.69]	[5.79]	[0.040]	[21.61]	[5.10]	[0.021]	[0.113]	[0.020]	[0.062]
Mauritius	1.17	-1.16	0.002	1.23	-1.16	0.18	0.27	5.01	11.91
	[7.18]	[3.92]	[0.962]	[6.46]	[3.22]	[0.673]	[0.200]	[0.061]	[0.166]
Morocco	0.59	-0.18	16.750	0.92	-0.48	14.52	0.25	7.07	22.66
	[1.94]	[0.47]	[0.000]	[3.00]	[1.17]	[0.000]	[0.200]	[0.029]	[0.010]
Niger	0.64	0.01	40.470	0.63	0.15	123.18	1.61	12.82	35.25
	[3.38]	[0.07]	[0.000]	[3.74]	[0.87]	[0.000]	[0.110]	[0.010]	[0.010]
Rwanda	1.36	-1.52	3.005	1.35	-1.48	2.13	0.36	7.02	39.15
	[18.98]	[10.26]	[0.083]	[17.27]	[9.37]	[0.145]	[0.200]	[0.030]	[0.010]

Notes: The values in square brackets are the t-statistic. For Hansen's Stability test, the values in parentheses are the p-values. ϕ is the Walk test of $\beta_1 = - \beta_2$, and each p-value is beneath the test statistic.

D. Constancy of the Cointegration Space

Having provided evidence concerning cointegration and some relevant hypotheses, it seems prudent to examine whether the estimated elasticities are stable over time. Hansen (1992) suggests three tests for parameter instability in cointegrated systems.

The three tests have parameter stability as their null hypothesis, but they differ in their alternative hypothesis. Also, each test assumes that the location of the potential break point in the cointegrating relation is unknown.Having provided evidence concerning cointegration and some relevant hypotheses, it seems prudent to examine whether the estimated elasticities are stable over time.

The first test statistic, Lc, permits testing of parameter stability as well as cointegration because it considers instability due to relatively constant parameter variation over the sample period. The second test statistic, SupF, is similar to the classical Chow F-tests. It assumes as the alternative hypothesis a sudden shift in regime at an unknown point in time, and it entails computing the classical Chow F-statistic for a subset of possible change point and then basing a test upon the supremum of these statistics. Specifically, the supremum of this sequence is:

$$SupF = \underset{t/T \in [0.15, 0.85]}{SupF_{t/T}} \tag{2}$$

where $F_{t/T}$ is the F-test statistic corresponding to the classical Chow test or the split sample test for a fixed t.

Finally, the Mean F-test is designed to detect, instead, a slow shift of parameters and tests for overall stability of the model. It is computed as the average value of the $F_{t/T}$ defined earlier. A low P value, below 0.05 (say), for a particular test statistic is interpreted as the instability of the parameters of the cointegrating vector. The test results are reported in Table 3. Our results obtained using the FM-OLS and the covariance parameters are estimated using Quadratic spectral (QS) kernel on residuals pre-whitened with a VAR(1).

The Lc statistic rejects the null hypothesis of parameter stability and cointegration in the Philippines and Singapore at the 5 percent level. The Mean F-test statistics are statistically significant in the seven countries: namely, Cote d'Ivore, Madagascar, Morocco, Niger, the Philippines, Rwanda and Singapore. This implies that parameter stability is rejected. Also, the SupF test statistic rejects parameter stability in six countries: namely, Kenya, Morocco, Niger, the Philippines, Rwanda and Singapore.

Three conclusions emerge from these results. First, the probability values for the estimated test statistics show that, for most of the Asian countries, the null hypothesis of parameter stability cannot be rejected. We interpret this conclusion as strong support for the proposition that PPP holds as a long-run equilibrium condition in Asia. Another important result is that, while the nominal exchange rate and price indices are cointegrated in each African country, the existence of a constant-parameter equilibrium long-run relationship is rejected by either the MeanF and SupF test statistics. The exception is Mauritius. Finally, there are two countries -- the Philippines and Singapore -- where all three tests (Lc, MeanF and SupF) reject the PPP relation.

E. Fractional Cointegration Tests

The empirical results reported above have been obtained using the conventional cointegration approaches. Such results are inadequate when the residuals from equation (1) are mean reverting but not I(0). In this section, we present the results from implementing the fractional cointegration technique which deals with cases where the residuals from equation (1) have a long memory, with $0<d<1$. The residuals will be fractionally integrated but mean-reverting if $0 < d < 1$. From amongst the different methods used for implementing cointegration tests based on the estimation of the Autoregressive Fractionally-Integrated Moving Average (ARFIMA) model, we have opted to use the GPH and the exact maximum likelihood (EML) of Sowell (1992).

The GPH method of testing for cointegration is a three-step approach. In the first step, the residuals from equation (1) are obtained by means of ordinary least square (OLS). In the second step, to ensure stationarity of the series, the residuals in first differences are generated and in the third step, the GPH test is used to estimate the differencing order $\hat{d}$. The value of estimated d can be used to test whether $:_t$ in equation (1) is I(1). The presence of a unit root in the residuals may be examined by testing the null $d = 1$ ($\hat{d} = 0$), against the alternative $d < 1$ ($\hat{d} < 0$). That is, the null hypothesis that nominal exchange rate, domestic and foreign price indices are not cointegrated is tested against the alternative hypothesis that nominal exchange rate, domestic and foreign price indices are cointegrated. The t-statistic is used for testing the null hypothesis, and it is a test of the estimated regression coefficients from

$$\ln\{I(w_j)\} = \beta_0 + \beta_1 \ln\{4Sin^2(w_j/2\} + \eta_t \tag{3}$$

where $\beta_1 = -d$, $I(w_j)$ is the periodogram of x_t at frequency w_j, $w_j = 2B_j/T$ ($j = 1,\ldots,T - 1$) and 0 are asymptotically i.i.d. The number of low-frequency ordinates (n) used in this test is equal to $n = T^v$, where T is the number of observations. Following the recommendation of Hurvich, Deo and Brosky (1998), we use a v equal to 0.60. Table 4 reports these test results.

Table 4
Fractional cointegration tests

Countries	GPH $d{:}H_0{:}d=1$	Maximum Likelihood ARFIMA *(p,d,q)*	d	T-Value
Asian Countries				
India	-0.007(5.33)	(0,d,0)	0.1528	3.25[0.001]
Korea	0.075(9.59)	(0,d,0)	0.3279	5.17[0.000]
Malaysia	-0.113(4.01)	(0,d,0)	0.1838	3.74[0.000]
Philippines	0.033(5.55)	(0,d,0)	0.0749	1.70[0.090]
Singapore	-0.171(6.53)	(0,d,0)	0.1573	3.06[0.002]
Sri Lanka	-0.045(8.96)	(0,d,0)	0.3322	6.00[0.000]
Thailand	0.103(6.80)	(0,d,0)	0.1601	3.15[0.002]
African Countries				
Cote d' Ivory	0.018(6.39)	(0,d,0)	0.0704	1.47[0.143]
Kenya	-0.049(3.55)	(0,d,0)	0.0441	0.912[0.362]
Madagascar	-0.337(3.37)	(1,d,0)	-0.3803	-3.05[0.002]
Mauritius	-0.227(3.87)	(0,d,0)	0.1758	3.30[0.001]
Morocco	0.005(5.23)	(0,d,0)	0.2716	4.63[0.000]
Niger	-0.217(4.72)	(0,d,0)	-0.0226	-0.49[0.625]
Rwanda	-0.523(4.79)	(0,d,0)	0.2112	3.45[0.001]

Notes: GPH test for v=0.6 and the t - values are in parentheses.

As can be seen, all residual series are fractionally integrated since the fractional difference parameter is significantly different from zero. These results indicate that the null hypothesis of no cointegration can be rejected by the GPH estimator. Since the GPH test has poor size properties, as a cross-check, we report the results obtained using the exact maximum likelihood procedure.

The second section of Table 4 reports the parameter estimates for various models from ARFIMA (0,d,0) to ARFIMA (1,d,0). These parameterizations were obtained using the AIC selection technique, and the results are qualitatively the same as those obtained from using the GPH estimator. Observe that the null hypothesis of no fractional cointegration is rejected in India, Korea, Madagascar, Malaysia, Mauritius, Morocco, the Philippines, Rwanda, Singapore, Sri Lanka and Thailand. It was not rejected in three countries, namely Cote d'Ivoire, Kenya and Niger. The null hypothesis of no fractional cointegration is rejected in all Asian countries, as can be seen from the p-values associated with the t-values. The obtained evidence for the Philippines is weak and only statistically significant at the 10 percent level.

4. Summary and conclusion

This paper has examined the validity of the purchasing power parity as a long-run equilibrium condition by applying time series tests for the cointegration of exchange rates and price indices. The focus was on the PPP between the United States and fourteen developing countries: Cote d'Ivoire, India, Kenya, Korea, Madagascar, Malaysia, Mauritius, Morocco, Niger, the Philippines, Rwanda, Singapore, Sri Lanka and Thailand. The empirical study was conducted using monthly data from 1973:4 through 2002:8 (i.e., 357 observations). It employs the Johansen estimator, which tests the null hypothesis of no cointegration, and the Harris and Inder (1994), which tests the null hypothesis of cointegration. To account for the possibility that the residuals may be mean reverting, but not integrated of order zero, we implemented the GPH and the exact ML tests for fractional cointegration. The application of both conventional and fractional cointegration test techniques, as well as the testing for structural stability of the PPP relation are distinct features of this paper compared to earlier studies on developing countries.

The empirical results provide ample evidence of cointegration between exchange rate and price indices. These results indicate that PPP holds as a long-run equilibrium condition. Another attractive feature of the results is that for cases such as Madagascar, Malaysia, the Philippines and Thailand, where the Johansen estimator did not find evidence of a long-run equilibrium relationalship, it was shown that strong evidence of cointegration is easily obtained using either the fractional cointegration techniques or the Harris and Inder (1994) test. Moreover, an appealing aspect of the results is that the long-run PPP relationship is generally stable in most of the countries. We found this stability of the parameter estimates to be strong evidence in favor of the long-run PPP in developing countries. We also found evidence that the PPP relations for Asian countries were generally more stable than those of the African countries. Our results provide a wider and more significant support for PPP than has been found in most previous studies of developing countries. We attribute the difference to the (more appropriate) statistical techniques we have implemented in this study—existing studies have utilized ADF-type cointegration tests, but it is now well recognized that these tests have low test power against stationary alternatives.

Finally, this paper illustrates the need to adequately test for cointegration in order to correctly interpret the results from finite samples. This is consistent with the evidence in Gregory (1994), which suggests that instances of conflicting test results are very likely to occur in cointegration analyses because of sharp power differences. In this light, Gregory (1994: 359) notes that it is "of considerable practical importance to calculate and report several tests for cointegration in applied studies."

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