Real Exchange Rate Stationarity in Managed Floats

Evidence from India

The paper tests for mean-reversion in real exchange rates for India during the recent float period. Using unit root tests with improved power, we test for stationarity of the real exchange rate, using several definitions of the real exchange rate. We also conduct cointegration and variance ratio tests to complement the evidence from unit root tests. We find evidence of mean-reversion in the real exchange rate series constructed with the consumer price index as deflator, as well as for a series constructed using the ratio of wholesale and consumer price indices to proxy for the shares of tradable and non-tradable goods.

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I
Introduction

Purchasing power parity, or the law of one price, is one of the most popularly tested theories in economics. In fact, tests for PPP have evolved along with time series analysis as much advancement in time series techniques have been applied in attempts to uncover parity reversion in real exchange rates. Testing for mean-reversion in the real exchange rate is important for many reasons. It is an important constituent of most models of exchange rate determination, being regarded as a long-run equilibrium or an arbitrage condition in goods and assets markets. The real exchange rate dynamics implied in models of inter-temporal smoothing of traded goods consumption [Rogoff 1992] and cross-country wealth redistribution/ transfers [Obstfeld and Rogoff 1995] makes the PPP hypothesis a meaningful one to examine. The concept is also of interest to policy-makers as it serves as a benchmark for computing an equilibirum exchange rate and assessing whether shocks to the real exchange rate dampen over time. A failure to reject the random walk hypothesis, for example, implies that permanent, real shocks explain real exchange rate movements. On the other hand, finding mean-reversion properties implies that nominal disturbances have no permanent impact upon the real exchange rate.

The behaviour of the real exchange rate and its responses to nominal and real disturbances as part of the macro-adjustment process assume significance for India, which has recently shifted to a market-determined (though managed) exchange rate regime. As Froot and Rogoff (1995) note, changes in exchange rate regime imply that deviations from parity might be eliminated through different processes altogether. Adjustments to parity are made through domestic price level movements in a fixed exchange rate regime, but when the regime is a float; parity reversion takes place via nominal exchange rate movements. Apart from changes in exchange rate regime, trade liberalisation and loosening of foreign exchange restrictions in India during the past decade has reduced many distortions, factors that suggest convergence in theory, at least in tradable goods. Finally, the exchange rate plays a central role in maintaining external and domestic equilibrium and understanding its response to shocks is important to policy-makers. This paper therefore aims to investigate how much evidence for PPP can be found for India during the float period.

Empirical evidence for developing countries on this issue is still fairly thin and concentrated around relatively homogeneous groups of countries, e g, Latin America or east Asia, leaving a gap of individual country studies with time series data [Edwards 1999]. This study attempts to fill this gap as it focuses exclusively on India. Existing empirical evidence for India finds some support for the PPP hypothesis for the pre-float period [Berg and Jayanetti 1995 for 1957-87; Baghestani 1997 for 1973:1-1991:2]. These studies however, restrict themselves to cointegration tests, which are weak tests for the PPP hypothesis. Our paper differs from these studies in two respects. One, we employ comprehensive tests, viz, unit root, variance ratio tests and cointegration, testing for both strong and weak mean-reversion in the real exchange rate. We check the robustness of results using different definitions of the real exchange rate. Two, we test for mean-reversion restricting ourselves to the post-float period, using monthly data for 1993:01-2001:03. The period under the float may indeed be considered too short to reveal mean-reversion in the real exchange rate. But we do not extend the sample backwards as gains from long-sample evidence may be offset by the fact that the float period is a very small proportion of the long sample. Moreover, some studies [e g, Baxter and Stockman 1989; Taylor 2000] have identified the regime dependency of real exchange rates. Combining fixed and floating exchange rate regimes in such circumstances is likely to inhibit uncovering parity reversion under the float. Instead we employ efficient unit root tests with improved power to overcome this gap.

The paper is organised into four sections. Section II briefly reviews existing empirical evidence on the issue and identifies problems encountered by researchers in uncovering parity reversion, Section III tests for mean-reversion with four different series of real exchange rate using unit root, variance ratio and cointegration tests. Section IV concludes.

I
Purchasing Power Parity: Methods, Problems and the Evidence

Testing for purchasing power parity has evolved considerably with the introduction of more powerful testing methods. Early tests for PPP focused on estimates of the coefficient on relative price levels, testing for absolute PPP, where the nominal exchange rate equals the foreign-domestic price levels, or relative PPP, where changes in relative price levels are offset by changes in the exchange rate. A useful survey of these early studies is to be found in Officer (1976). Simultaneity and the possibility of non-stationary exchange rate and prices shifted the focus upon testing time series properties of the residuals. Termed as stage two tests by Froot and Rogoff (1995), these consist of testing the hypothesis that the log of real exchange rate follows a random walk, or is non-stationary. Another strand of tests (Stage III tests) utilise cointegration techniques to test for a long-run equilibrium relationship between the nominal exchange rate and price levels. The serious small sample bias of cointegration tests and difficulties in interpretation of the long-run coefficients have however made unit root tests more preferable since they directly test mean-reversion in the real exchange rate [Froot and Rogoff 1995].

Empirical evidence for the industrialised countries conforms to the opinion that the random walk hypothesis is difficult to reject for floating currencies, implying that nominal disturbances to the real exchange rate have permanent, or infinite, effects. Most studies also report estimates of half-lives of parity deviations between 3-5 years, indicating a slow convergence to equilibirum. Evidence regarding permanent deviations from parity, i e, the Belassa-Samuelson effect, is not supported by data from industrialised countries except for Japan. The use of panel data and more powerful unit root tests has helped uncover more support for the hypothesis. But mean reversion in the real exchange rate in developing countries' data has been difficult to find, partly because the shift to floating exchange rates amongst this group of countries has been recent, making it meaningless to apply PPP-based models on such data. An up to date summary for developing countries can be found in Edwards (1999), who notes that empirical evidence on mean-reversion for developing countries is sparse, with far more evidence for Latin American countries. For currencies that have been formally stabilised, like the intra-European exchange rates, the evidence is more mixed.

Since extensive summaries of these empirical approaches/studies already exist [Froot and Rogoff 1995; Rogoff 1996, to name two), the focus here is only to shortlist the problems associated with testing for mean-reversion. Empirical studies investigating mean-reversion have found the results to be sensitive to the choice of price index, countries and time period. These are explained mainly by heterogeneity in the construction of price indices across countries, the presence of trade restrictions and the fact that many goods are not traded. Moreover, the post-Bretton Woods period, to which most of the early PPP studies pertain, does not provide a sufficiently long period for reasonable time series analysis. These factors have rendered empirical tests of PPP difficult and unstable.

Researchers have tried to overcome these problems in a number of ways. Apart from the use of more refined or disaggregated price indices, the most popular response has been to increase the sample period, thereby combining fixed and floating exchange rate regimes. This however has been questioned on grounds of regime changes. As Frankel and Rose (1995) argue, if different processes govern the real exchange rate during the float period, then test results are clearly biased. In fact, a growing body of literature refers to the variation in the statistical properties of real exchange rates across different nominal exchange rate regimes. For example, Taylor (2000) has investigated PPP, using a century of data for a group of 20 countries and finds that changes in the size of shocks to the real exchange rate depend on the political economy of monetary and exchange rate regime choice. Thus there are strong reasons for confining testing of time-series properties within a single regime.

A relatively recent response to overcome insufficient time series variation has been the introduction of cross-sectional variation through the use of panel data [Abuaf and Jorion 1990; Frankel and Rose 1995; Jorion and Sweeney 1996; Papell 1997; O'Connell 1998 amongst others). Studies with panel estimation have uncovered mean-reversion in several instances, indicating the utility of greater variation within a single regime. In the same vein are studies using tests with higher power [Cheung and Lai 1998; Culver and Papell 1999] and non-linear methods [Taylor and Sarno 1998]. This generation of research not only rejects the random walk hypothesis more frequently, but also demonstrates that increased power of unit root tests does uncover mean-reversion in small samples of the float period.

This paper seeks to combine these developments with data from a developing country, i e, India, and test for mean-reversion during a single regime. It responds to the problems identified above in three different ways. First, we confine ourselves to a single exchange rate regime, viz, the float period. In order to overcome the low power problems associated with conventional unit root tests, we employ unit root tests with improved power, using the modifications proposed by Elliott, Rothenberg and Stock (1996) and Park and Fuller (1995). We also compute variance ratio statistics to uncover parity. Finally, we employ a weaker test of mean-reversion, viz, cointegration, using the popular procedure due to Johansen and Juselius (1990). The following section deals with these in sequence.

II
Testing for Mean-Reversion

This section probes the real exchange rate of the rupee for mean-reversion, using four different series. The real exchange rate is calculated as

$$q_t = e_t + p_t^* - p_t \qquad ...(1)$$

where e_t is the nominal (dollar) exchange rate expressed in domestic currency, p_t is the domestic price index and p_t^* is the analogous US price index. All variables are expressed in logarithms. The first series is constructed using the consumer price indices of India and the US respectively to obtain a CPI-deflated real exchange rate series. The second series is constructed using the ratio of wholesale/producer and consumer price indices (India and the US) on the assumption that these two price series proxy for the share of tradable and non-tradable goods respectively. Since the manufacturing index covers a greater proportion of tradable goods' prices, one would expect, a priori, PPP to hold more strongly with these indices. Finally, we use the two trade-weighted real exchange rate indices (36 countries and 5 countries)

Figure 1: CPI Deflation Exchange Rate

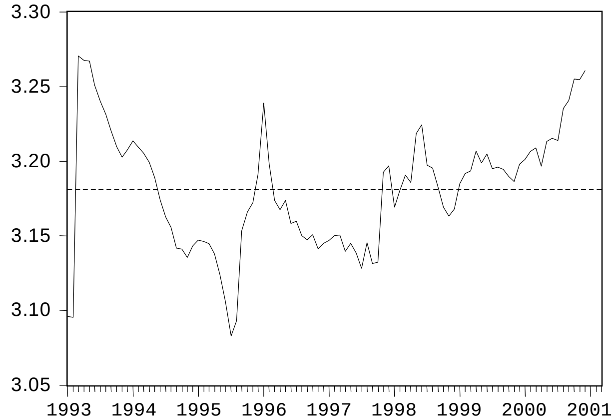


Figure 2: 5-Country Weighted Real Effective Exchange Rate

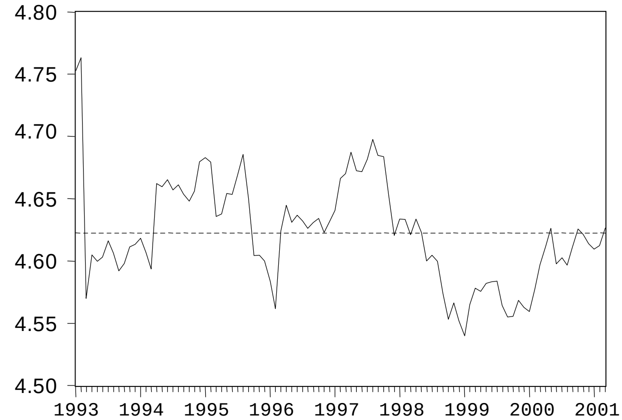


Figure 3: 36-Country Weighted Real Effective Exchange Rate

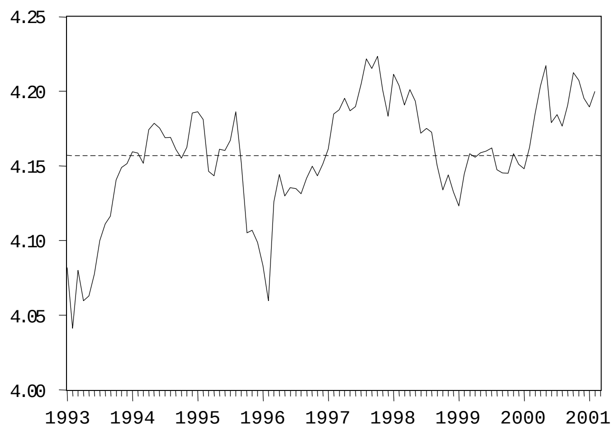
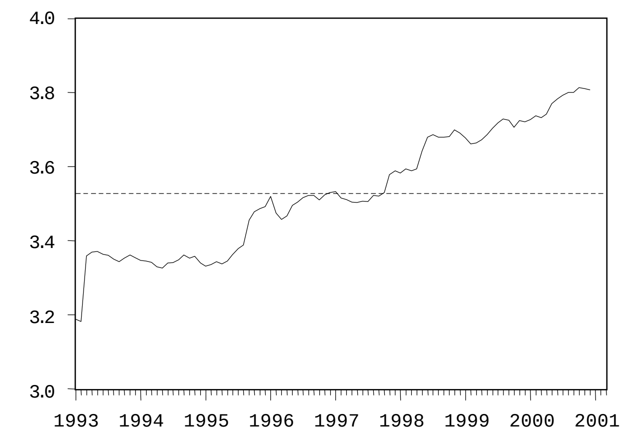


Figure 4: Ratio of WPI to CPI Deflated Exchange Rate



published by the Reserve Bank of India. The data is monthly, spanning the 1993-2001:03 time horizon.

The four series are plotted in Figures 1-4, centred by their respective sample means. The two trade-weighted REER indices and the CPI-based real exchange rate series exhibit some patterns of reversal. The fourth series, constructed as the ratio of wholesale to consumer prices for US/India respectively, appears to be stationary around a strong deterministic trend. One explanation of a trend in the relative prices of traded and non-traded goods is widening differentials in productivity growth in the two sectors, as has been shown for Japan by Obstfeld (1993). Other reasons could be change in composition of tradable goods over time, changes in the cost of goods arbitrage (reflecting trade liberalisation, reductions in transportation costs, etc) or simply, measurement errors that may affect the proportionality of non-traded and traded goods prices.[1] The presence of these effects implies that PPP fails to hold or that shocks to the real exchange rate are infinitely lived.

Unit Root Tests

A strict version of PPP requires that the real exchange rate be constant. Termed as stage two tests by Froot and Rogoff (1995), the test is based upon whether the real exchange rate contains a unit root. If this hypothesis is rejected then there is evidence of mean-reversion, i e, the real exchange rate is not governed by permanent shocks. Thus in the equation

$$r_t = \mu + \rho r_{t-1} + \varepsilon_t \qquad ...(2)$$

the stationarity of ε_t is necessary if the real exchange rate is mean-reverting. If $\rho \geq 1$ then r_t is a non-stationary process and the nominal exchange rate and the price differential deviate from one another, suggesting that some shocks to the real exchange rate are permanent. There can be valid economic reasons for the real exchange rate to be non-stationary. The permanent components in real exchange rate movements can be explained by increased productivity induced real exchange rate appreciation [Balassa 1964]; permanent changes in relative productivity of traded and non-traded sectors [Baumol-Bowen 1966]; permanent changes in government spending [Froot and Rogoff 1991; Alesina and Perotti 1995], or simply a bias in the measurement of the consumer price index.

A conventional test for unit roots in a series is the Augmented Dickey-Fuller (ADF) test, which involves regressing the first difference of a series on a constant, its lagged level and p lagged first differences, i e,

$$\Delta q_t = \mu + \rho q_{t-1} + \Sigma_{t=1}^{p} \alpha_i \Delta q_{t-1} + \varepsilon_t \qquad ...(3)$$

The p lagged first-differences are included to control for autocorrelation errors. The value of p was selected using the recursive t-statistic procedure, starting from a maximum value of p=12, since the data is monthly in frequency. A significance level of 5-10 per cent (1.645) was used to assess the significance of the last lag.

The results of the ADF tests, demeaned and detrended, are presented in Table 1.[2] The table also shows the results of the Phillips-Perron unit root tests, which allows for conditional heteroskedasticity of

the residuals. We leave the detailed comparison of the results for later. Suffice it to say for now that the data supports the stationarity hypothesis when we use a relaxed criterion for the CPI and WPI/CPI deflated real exchange rate series. Stationarity is unequivocally accepted only in the case of the 5-country trade weighted REER series. Using stricter levels of significance, i e, 1 per cent level of significance, we are unable to reject the random walk hypothesis with the rest of the series.

Unit root tests are well known to have low power, especially in small samples. One response to this problem has been to search for improved efficiency unit root tests. Cheung and Lai (1998) employ two efficient univariate tests proposed by Elliott, Rothenberg and Stock (1996) and Park and Fuller (1995) to uncover parity reversion. These tests require much shorter sample sizes than conventional unit root tests to attain the same statistical power. The ERS (1996) modification to the augmented Dickey-Fuller test uses the generalised least squares estimation (DF – GLS test). This test is based upon an analysis of the sequence of Neyman-Pearson tests of the null hypothesis H_0:ρ =1 against the local alternative H_a:ρ =1 + eIT, where e<0. The locally detrended data process is obtained by regressing the real exchange rate series q_t on z_t

$$q_t = q_t - z_t\beta \qquad ...(4)$$

where β is the least squares coefficient of $\tilde{q}_t$ on $\tilde{z}_t$. $\tilde{q}_t = (q_1, (1-\rho L)q_2,..., (1-\rho L)q_T)'$ and $\tilde{z}_t = (z_1,(1-\rho L)Z_2,...,(1-\rho L)Z_T)'$. The demeaned process is similarly obtained by replacing q_t above with q_t^μ and z_t = 1. The $DF-GLS^t$ (detrended) and the $DF-GLS^\mu$ (demeaned) tests are then based on the following regression:

$$\Delta q_t = \phi_0 q_{t1} + \Sigma_{j=t}^{p} \phi_j \Delta q_{tj} + \nu_t \qquad ...(5)$$

and the test statistic is the t-statistic on q_{t1}, testing H_0:ϕ_0 = 0 against H_0:ϕ_0 < 0. ERS (1996) show that this modification yields substantial power improvement with small samples, which is equivalent to the power attainment of conventional ADF tests in large samples. The parameter defining the local alternative e for calculating the series with time trend is set equal to -13.5 and without trend is set to -7, as recommended by ERS (1996). The results of the ADF-GLS tests are presented in Table 1, but before we analyse these, we discuss the modification of the augmented Dickey-Fuller test proposed by Park and Fuller (1995).

The Park and Fuller modification makes use of weighted symmetric least squares estimation after demeaning and detrending the real exchange rate series q_t. The DF – WS test requires minimising a weighted sum of errors with respect to ρ and α in (1) above, where α = (α_1, α_2,...α_p) and 'p' is the number of lags selected on the basis of significance of the 't' statistic on α. The weights w_t in the estimation are specified as w_t(t=1,2,...,T).[3] The test statistic is $\tau_{ws} = \{V(\hat{\rho})\}^{-1/2} (\hat{\rho}-1)$ where $V(\hat{\rho})$ is the estimated variance from the WSLS regression and the hypothesis H_0:ρ =1 is tested

Table 1: Unit Root Tests

(1993-2000)

Exchange Rate	Augmented Dickey Fuller		Phillips-Perron	
	Detrended	Demeaned	Detrended	Demeaned
$e + pcpi^* - pcpi$	3.37***	3.34**	3.33***	3.30**
$e + \frac{pppi^*}{pcpi^*} - \frac{pwpi}{pcpi}$	3.97**	1.39	3.66**	1.23
REER-5 country weighted	4.29*	4.39*	4.18*	4.16*
REER-36 country weighted	2.56	2.52	2.73	2.52

Critical values at 1 per cent, 5 per cent and 10 per cent for the ADF and Phillips-Perron tests are 3.5, 2.89 and 2.58 respectively for the intercept. Critical values at the 1 per cent, 5 per cent and 10 per cent for the ADF and Phillips-Perron are 4.05, 3.45 and 3.15 respectively for trend and intercept. *(**(***)) indicate significance at 1, 5 and 10 per cent level respectively.

Exchange Rate	DF-GLS				DF-WS	
	MAIC		Ng-Perron		Ng-Perron	
	Detrended	Demeaned	Detrended	Demeaned	Detrended	Demeaned
$e + pcpi^* - pcpi$	2.66	0.74	2.48	1.30	3.08***	2.80**
	(2.73)	(1.77)	(2.74)	(1.81)	(2.96)	(2.26)
$e + \frac{pppi^*}{pcpi^*} - \frac{pwpi}{pcpi}$	2.61	1.00	2.61	1.00	3.16***	1.76
	(2.73)	(1.80)	(2.73)	(1.80)	(3.12)	(2.35)
REER-5 country weighted	2.45	1.22	2.62	1.35	2.12	2.15
	(2.72)	(1.80)	(2.74)	(1.81)	(2.99)	(2.26)
REER-36 country weighted	2.11	0.80	2.40	1.07	2.82	2.08
	(2.71)	(1.79)	(2.74)	(1.81)	(2.99)	(2.26)

Finite sample critical values for the DF-GLS and DF-WS are given in parentheses. *(**(***)) indicate significance at 1, 5 and 10 per cent level respectively. The lag lengths for the DF-GLS test were selected using both minimised AIC and Ng-Perron's 't' statistic criteria. Use of the MAIC criterion provides substantial size improvements in the DF-GLS context (Ng-Perron 2000).

Table 2: Unit Root Tests

(1993-99)

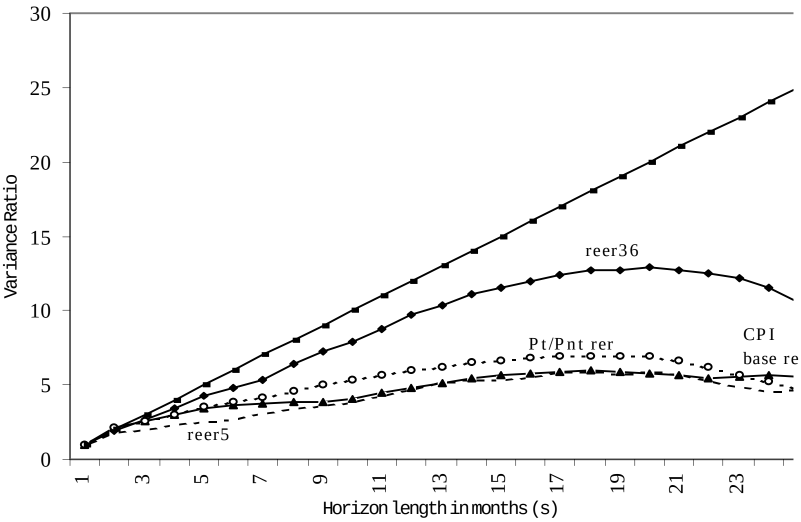
Exchange Rate	Augmented Dickey Fuller		Phillips-Perron	
	Detrended	Demeaned	Detrended	Demeaned
$e + pcpi^* - pcpi$	3.41***	3.39*	3.42***	3.39**
$e + \frac{pppi^*}{pcpi^*} - \frac{pwpi}{pcpi}$	3.78**	1.64	3.31**	1.46
REER-5 country weighted	3.78**	3.65*	3.78**	3.65*
REER-36 country weighted	2.35	2.48	2.35	2.48

Critical values at 1 per cent, 5 per cent and 10 per cent for the ADF and Phillips-Perron tests are 3.5, 2.89 and 2.58 respectively for the intercept. Critical values at the 1 per cent, 5 per cent and 10 per cent for the ADF and Phillips-Perron are 4.05, 3.45 and 3.15 respectively for trend and intercept. *(**(***)) indicate significance at 1, 5 and 10 per cent level respectively.

Exchange Rate	DF-GLS				DF-WS	
	MAIC		Ng-Perron		Ng-Perron	
	Detrended	Demeaned	Detrended	Demeaned	Detrended	Demeaned
$e + pcpi^* - pcpi$	2.76***	1.51	2.56	1.81	2.36	2.47**
	(2.75)	(1.79)	(2.77)	(1.84)	(2.96)	(2.26)
$e + \frac{pppi^*}{pcpi^*} - \frac{pwpi}{pcpi}$	2.86***	0.55	2.86***	1.69	3.26***	0.95
	(2.75)	(1.83)	(2.75)	(1.79)	(2.99)	(2.26)
REER-5 country weighted	2.48	1.02	2.69	1.19	3.70**	0.80
	(2.75)	(1.83)	(2.77)	(1.84)	(3.58)	(2.23)
REER-36 country weighted	1.74	1.05	2.29	1.34	1.93	2.21
	(2.76)	(1.79)	(2.72)	(1.84)	(3.04)	(2.23)

Finite sample critical values for the DF-GLS and DF-WS are given in parentheses. *(**(***)) indicate significance at 1, 5 and 10 per cent level respectively. The lag lengths for the DF-GLS test were selected using both minimised AIC and Ng-Perron's 't' statistic criteria. Use of the MAIC criterion provides substantial size improvements in the DF-GLS context (Ng-Perron 2000).

Figure 5: Variance Ratio Tests



against H_0:ρ < 1.

The ADF-GLS and ADF-WS unit root test results are presented in Table 1. Both demeaned and detrended cases are considered for four real exchange rate series. The DF-GLS test rejects stationarity of the real exchange rate, however defined, for the entire 1993-2000 sample. But the DF-WS test supports stationarity for the demeaned and detrended CPI-deflated real exchange rate series as well as the detrended WPI/CPI deflated series. Stationarity of all other series is however, rejected when the DF-GLS and Df-WS tests are applied to the data.

Are these results stable over the sample period? Trimming the sample down to 1993-99 shows that much of the results are undisturbed (Table 2) except for the DF-GLS test for the CPI-deflated and the WPI/CPI real exchange rate series, both of which are now stationary. One must recall though that, these are borderline cases, since the non-stationarity null is rejected at a 10 per cent level of significance. Apart from these changes, the full sample results are basically preserved, indicating the robustness of the unit root tests

Finally, both ADF-GLS and the ADF-WS tests reject mean-reversion in the 5 and 36 country trade-weighted REER series, except for the detrended 5-country REER where parity reversion is uncovered over the 1993-99 time horizon. The 36-country REER does not exhibit mean-reversion at all.

What conclusions can be drawn from these tests, given the mixed evidence on parity reversion? Both ADF and Phillips-Perron test show the CPI-base, the P_t/P_{nt}-base and the 5-country index of real exchange rates to be a stationary series. For the CPI based real exchange rate series, it is not possible to find mean-reversion, using stricter confidence levels. However, when we subject the series to the DF – GLS test for unit roots, the series exhibits mean-reversion at a 5 per cent significance level. The same is the case with the series constructed to reflect the concept that parity holds for traded goods. The CPI-based and the P_t/P_{nt} based real exchange rate series are the only two series where parity can be detected with the aid of a more efficient unit root tests, viz, the DF – WS tests. This conforms to the findings of other researchers too, notably Cheung and Lai (1998).

The evidence regarding trend stationarity of the WPI/CPI real exchange rate series is significant. Noting that PPP holds only with the inclusion of the trend term, it hints at the presence of the Balassa-Samuelson effect discussed earlier in the paper. This issue is worthy of further exploration in research. The results reveal that the real exchange rate computed with the US dollar as the base currency exhibits a greater tendency to display mean-reversion than vis-a-vis other currencies as the base. This is in conformity with stylised evidence on the issue for other countries as well. Moreover, this result is also in line with evidence where it is more common to detect parity for countries which stabilise their currencies. It is widely believed that the rupee is stabilised with respect to the US dollar. Therefore it is not surprising to find more compelling evidence for mean-reversion in the rupee's real exchange rate vis-a-vis the dollar. This difference in the different exchange rate series is confirmed when we computed variance ratio statistics to test for unit roots, which are presented next.

Variance Ratio Tests

In the next stage of investigating real exchange rate stationarity we computed the variance ratio test statistic proposed by Cochrane (1988). Under the null of a random walk, the variance of series q_t is hypothesised to grow linearly over time. For a stationary series, the variance ratio statistic, Var $(q_{t+s} – q_t)/(q_{t1}–q_t)$, converges to zero as S increases. If the true process is I (1), this statistic gives a quantitative measure of the effects of permanent shocks upon the real exchange rate.

Figure 5 plots the variance ratio statistic for the four measures of the real exchange rate. From the figure we can observe that the variance ratio statistic rises more sharply over the horizon (s = 1 to s = 25) for the 36-country real effective exchange rate (reer36) index than for the 5-country index (reer5), the CPI-based or the wholesale/consumer prices-based real exchange rate series. As the horizon length increases, the variance ratio statistic for all series, except the 36-country REER, converges. While by itself, this does not constitute conclusive evidence of stationarity, it does lend further support to the evidence from the unit root tests where we were unable to reject the random walk hypothesis for the 36-country, trade-weighted REER series.

Cointegration Tests

Cointegration tests are weak tests of PPP since they require only that some linear combination of domestic price level (P_t), the foreign price level (P_t^*) and the nominal exchange rate (e_t) be stationary. The null hypothesis here is that of no cointegration between the three series. A stronger test for PPP can be implemented by imposing proportionality and symmetry restrictions upon the coefficients on p_t^* and to be 1, -1. Thus in the equation

$$\ln e_t = \beta_0 + \beta (\ln p_t – \ln p_t^*) + u_t \quad ...(6)$$

the restriction β = 1 is tested, which is a bivariate test. Alternately, the trivariate case

$$\ln e_t = \beta_0 + \beta_1 \ln p_t + \beta_2 \ln p_t^* + u_t \quad ...(7)$$

may be examined, where the symmetry restriction $(\beta_1, \beta_2) = (1, –1)$ is tested. The evidence using cointegration tests suggests that while the cointegration null, a weaker hypothesis, is often rejected, the proportionality restrictions are typically violated.

We use the cointegration proccedure due to Johansen and Juselius (1990) and the results are presented in Table 3. As the table shows, we experimented with several combinations of the four exchange rate series. Both trivariate and bivariate systems were tested for cointegrating relationships. Inferences regarding the existence of a cointegrating vector between the nominal exchange rate, foreign and domestic prices are primarily based upon finite sample critical values as finite sample analyses can bias the likelihood ratio tests towards finding cointegration too often.[4] Finite sample critical values were obtained by adjusting the asymptotic critical values, details of which are reported below the table.

Using finite sample critical values, we are unable to reject the null of no cointegration for any of the bivariate and trivariate cases. However, at a more relaxed significance (>5 <10 per cent) level, it is possible to conclude that a stationary combination of the nominal exchange rate and the foreign and domestic price levels exists with a broader price index as the deflator. Moreover, evidence of cointegration can be found when a bivariate relationship between the nominal exchange rate and divergence between the wholesale price index (India) and the consumer price index (US) is hypothesised (Row 6, Table 3). The ratio of wholesale and consumer price levels (Row 7) proxies for the hypothesis that only the prices of tradable goods should be equalised across the two countries and is a close candidate for rejection of the no-cointegration null.

Thus, the data provides weak support for the hypothesis that parity with foreign price level holds for a more aggregate class of goods and to a large extent, for tradable goods. This result is not surprising in view of the stylised evidence for PPP tests where CPI-based tests are less frequently rejected than WPI based tests. This is probably due to the fact that the non-traded goods component in the consumer price index is higher than in the wholesale price index. Since the WPI has heavier weights for manufactured goods, PPP may hold to a larger extent for WPI than the CPIs [McKinnon 1971]. Finally, the estimates of β_1 and β_2 range widely from 0.46 to 4.04 and are often wrongly signed. Thus a one-to-one association between prices and the exchange rate does not hold over this time horizon.

The fact that we find less evidence for mean-reversion through cointegration tests is not very surprising as much of existing evidence shows that that the cointegration null is rejected more often than the unit root hypothesis. The tests are subject to power problems, i e, serious small sample bias, with failure to reject the null of no cointegration too often.

IV
Conclusion

This paper has examined mean-reversion in the real exchange rate for India after change in exchange rate regime in 1993. Tests for stationarity used unit root tests as well as cointegration and variance ratio tests and were applied to four different series of the real exchange rate. The evidence on parity reversion uncovered in this paper for the period after the float is somewhat mixed with bulk of the findings indicting mean-reversion for the consumer prices and the WPI/CPI deflated real exchange rate series. The use of recent developments in unit root testing like the ADF-GLS and ADF-WS tests, which yield substantial power improvements in small samples, shows that the real exchange rate series defined vis-a-vis the US dollar exhibits mean-reverting tendencies. The data does not however, support stationarity for the REER series, which is computed with a broader base of currencies. This may reflect the potential effects of currency stabilisation on unit root tests. It may also imply that models emphasising real determinants of the exchange rate may offer a more accurate description of real exchange rate behaviour as opposed to nominal disturbances.

Several caveats accompany the results obtained in this paper. As with most time-series studies suffering from data inadequacy handicaps, this study too is no exception. Thus the results of this study are to be interpreted with a great deal of caution given the extremely short time-span available for analysis, the fact that purchasing power parity is essentially a long run condition and uncertainty associated with the rate of reversion to an unconditional mean. For example, it could well be the case that nominal exchange rate movements due to short-term nominal price rigidities may affect the real exchange rate but different processes may govern the long-run behaviour of the real exchange rate. Possible hypotheses for exploration in this regard are the popular Belassa-Samuelson effect, where a rise in productivity in the tradable goods sector triggers a wage increase in the tradable goods sector, and if the productivity differentials between the two sectors widen, a price increase in the non-tradable goods sector is inevitable. One prediction of this theory is that fast growing economies will experience a real exchange rate appreciation, presuming that the traded goods sector is the locus of productivity increases. This would be a worthwhile hypothesis to explore in the Indian context, particularly as our examination suggests the failure of PPP to hold for

Table 3: Johansen – Cointegration Tests

System	λ_{trace}	Asymptotic Critical Values (5 Per Cent)	Finite Sample Critical Values (5 Per Cent)	Cointegrating Vector Coefficients Normalised with Respect to e_t
e_t, $pcpi^*$, $pcpi$	24.02	42.44	45.36	1.03, 4.04*
e_t, $(pcpi^* - pcpi)$	16.63	25.32	27.76	-0.006
e_t, $pppi^*$, $pwpi$	20.44	29.68	34.18	-1.58, 0.46
e_t, $(pppi^* - pwpi)$	15.63	15.41	16.48	0.74
e_t, $pcpi^*$, $pwpi$	26.92	42.44	47.04	-1.01, 4.00**
e_t, $(pcpi^* - pwpi)$	16.45	15.41	16.48	2.48
$e_t, \frac{pppi^*}{pcpi^*}, \frac{pwpi}{pcpi}$	47.01	42.44	48.88	1.58, -0.68
$e_t, \left(\frac{pppi^*}{pcpi^*} - \frac{pwpi}{pcpi}\right)$	15.44	15.41	16.9	-2.77

All variables are in logarithms. *, **, *** indicates significance at 10 per cent, 5 per cent and 1 per cent levels respectively. Assuming a priori that 12 lags, i e, one year, might be a reasonable dynamic representation of the data generating process, we began with a lag length of 12 months, paring down to parsimonious lag length using the multivariate generalisations of the AIC and SC as specification indicators. Successive lag lengths are reported along side each cointegrating relationship. The asymptotic critical values were adjusted to approximate finite sample properties, using the adjustment factor $CR_\infty[T/(T-nk)]$ proposed by Reinsel and Ahn (1988) where CR_∞ is the asymptotic critical value at the corresponding significance level, T the sample size, 'n' the number of variables in the estimated system and 'k' is the lag parameter.

the detrended relative prices of tradable and nontradable sectors.

A second aspect that could be explored to explain real exchange rate behaviour is fiscal policy effect. The long-run real effects of government spending have been shown to affect the real exchange rate by Froot and Rogoff (1991). It may of course, also be the case that none of these explanations may universally account for real exchange rate changes in the long-run and may complement purchasing power parity. Further, one has to make an allowance for policy effects like currency stabilisation vis-a-vis the dollar, which induce bias in the data. EPW

Notes

[The views expressed here are the author's own and not of the institution to which she belongs. I am grateful to K L Krishna for invaluable help and Mili Gupta for programming assistance. I am, of course, responsible for any errors.]

1 For instance, if the non-traded goods price index is subject to a fixed-weight or new-goods bias, a change in relative prices will generate upward index movements (Froot and Rogoff 1995: 1663).

2 A time trend usually does not feature in PPP tests as it is inconsistent with the hypothesis. Inclusion of a time trend however, controls for the presence of the Balassa-Samuelson effect, which hypothesises differential rates of growth in the tradable and nontradable sectors.

3 This is specified by $w_t = 0$ for $1 \leq t \leq p + 1$; $w_t = (t - p - 1)/(T - 2p)$ for $p + 1 < t \leq T - p$, and $w_t = 1$ for $T - p < t \leq T$.

4 In fact, the finite sample bias magnifies as the dimension of the estimated system and the lag-order increase [Cheung and Lai 1993].

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