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REAL EXCHANGE RATE BEHAVIOUR: EVIDENCE FROM BLACK MARKETS

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SUMMARY

The behaviour of real exchange rates (relative to the US dollar) is examined using monthly data obtained from the black markets for foreign exchange of eight Asian developing countries. The data span is 31 years. The black market real exchange rates do not show excess volatility during the recent float which is in sharp contrast to the results reported elsewhere. Unit root tests in heterogeneous panels and variance ratio tests confirm their stationarity. Thus, we find support for PPP but not for the 'survivorship' bias (Froot and Rogoff, 1995). There is little evidence of segmented trends. Issues raised by Rogoff (1996) — of whether PPP would hold across countries with differing growth experience — and Lothian and Taylor (1996) — of whether the degree of relative price volatility may bias results in favour of mean reverting real exchange rates — are addressed. Copyright © 2000 John Wiley & Sons, Ltd.

1. INTRODUCTION

The theory of purchasing power parity (PPP) is one of the most widely tested propositions in economics. The overall findings can be summarized as follows. Studies based on microeconomic data wholly reject PPP.[1] Rogoff (1996, p. 644) calls this set of evidence 'the abject failure' of the law of one price. Time-series studies based on aggregate price indices for the modern float era also largely reject PPP and suggest that the real exchange rate behaves as a random walk.[2] The random walk hypothesis implies that shocks to the real exchange rate are persistent and that there is no tendency for PPP to hold in the short run or in the long run. Rogoff (1996, p. 655) summarizes this set of findings as 'something of an embarrassment' to the PPP hypothesis and argues that every 'reasonable' theoretical model suggests a mean reverting real exchange rate in the long run.[3]

The short time span of the modern float has led economists to suspect that the rejection of a mean reverting real exchange rate is due to the low power of the tests.[4] Attempts have been made to address this issue by (1) adopting more powerful econometric tests,[5] and (2) extending either

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[1] Isard (1977), Richardson (1978), Giovannini (1988), Knetter (1989), and Engle and Rogers (1996) are some of the studies of relevance here.
[2] Breuer (1996), Taylor (1995) and Rogoff (1996) provide excellent surveys of the empirical literature on PPP. Jorion and Sweeney (1996) is among the few studies that find mean reversion in European real exchange rates.
[3] Roll (1979), and Adler and Lehmann (1983) have shown the theoretical possibility of a random walk real exchange rate. However, for the criticisms of this view see Cheung and Lai (1993), Abuaf and Jorion (1990) and Diebold *et al.* (1991).
[4] Among others, Campbell and Perron (1991), Hakkio and Rush (1991), Frankel and Rose (1996) and Lothian and Taylor (1996) make this point. Frankel (1986, 1990) shows that the post-Bretton Woods period is far too short to reliably reject the random walk hypothesis.
[5] See Abuaf and Jorion (1990), Diebold *et al.* (1991) and Fisher and Park (1991).

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the time span or the cross-sectional dimension of the data. Studies that use long-span data find, in the most cases, slowly mean reverting real exchange rates which implies that PPP holds in the long run.[6] The half-life of a shock to the real exchange rate is estimated somewhere at between 2·5 to 5·0 years. Cross-country panel studies also find mean reverting real exchange rates and suggest a half-life of PPP deviations of between 2·7 to 4·0 years.[7] Thus, a general consensus in the literature is that the long-horizon time-series and panel studies produce results which support PPP.

The observed empirical regularities are not without concern, however. The findings of long-horizon studies are questioned on the grounds that they mix different exchange rate and/or policy regimes as well as periods of war and peace. Rogoff (1996) argues that they blend fixed and floating rate data; yet the econometric consequences of mixing data from two regimes is unclear. Frankel and Rose (1996) point out that a typical 100–200-year sample includes several fixed, floating, and intermediate regimes with potentially serious structural shifts and argue that the nature of PPP adjustment may vary across the exchange rate regimes. In short, these concerns suggest possible structural breaks in the long-horizon data. In this paper we address this issue by formally testing for structural breaks in real exchange rates.

The deficiencies in the existing panel based unit-root tests are recently pointed out by Im *et al.* (1997). In particular: (1) traditional panel tests do not allow for dynamic heterogeneity across groups which results in inconsistent parameter estimates; and (2) tests proposed by Levin and Lin (1993) which have received wide usage (e.g. Frankel and Rose, 1996; Papell, 1997; Wu, 1996) have low power. We employ LM-bar and T-bar tests proposed by Im *et al.* (1997) which address these shortcomings.

Froot and Rogoff (1995) raise the issue of 'survivorship' bias. The argument is essentially what has become known as the Balassa–Samuelson effect which posits that rich countries will have higher exchange rate adjusted price levels than poor countries; hence deviations from PPP between rich and poor countries could be persistent.[8] Since the long-span studies which support PPP analyse industrialized countries' data (mainly because long time-series for poor countries are not available) for which the Balassa–Samuelson effect may be less pronounced, the question is whether their findings are exaggerated by the 'survivorship' bias. Indeed, Froot and Rogoff (1995) could not reject the random walk behaviour of the Argentine peso relative to the US dollar over more than 70 years of data. This prompted Rogoff (1996, p. 657) to ask whether PPP would hold across countries with differing growth experience. According to Froot and Rogoff (1995, p. 1660) this issue has 'not been studied extensively'. One of the motivations of this paper is to address this question precisely.

McNown and Wallace (1989) report mean reverting real exchange rates between currency pairs of four high inflationary developing countries and the US dollar during the modern float. However, Lothian and Taylor (1996) point out that the data set used by McNown and Wallace exhibits excess volatility in the relative price ratios. Besides, in high-inflation countries, monetary

[6] Frankel (1986), Edison (1987), Abuaf and Jorion (1990), Kim (1990), Glen (1992), Diebold *et al.* (1991), Grilli and Kaminsky (1991) and Lothian and Taylor (1996), are some of the major long horizon studies of PPP which report mean reverting real exchange rates. On the other hand, Alder and Lehmann (1983), Froot and Rogoff (1995) and Froot *et al.* (1995) do not find mean reverting real exchange rates even with long-span data.
[7] Typical studies in this line are Hakkio (1984), Frankel and Rose (1996), Oh (1996), Wu (1996) and Lothian (1997) all of which find mean reverting real exchange rates.
[8] See Balassa (1964) and Samuelson (1964). The argument is that the cross-country productivity differences can induce persistence in the real exchange rates between rich and poor countries.

growth is likely to overshadow real factors which may bias evidence in support of PPP.[9] According to Lothian and Taylor (1996) further investigations into a data set with low relative price volatility is required before these results (that real exchange rates between rich and poor countries are stationary) can be generalized.

We investigate the mean reverting behaviour of real exchange rates using a new data set from developing countries. The monthly black market real exchange rates of eight Asian currencies, i.e. Indian Rupee (IRupee), Malaysian Ringgit, Myanmar Kyat, Pakistani Rupee (PRupee), Philippine Peso, Sri Lankan Rupee (SRupee), Taiwanese dollar (TDollar) and Thai Bhat are used.[10] The data span is 31 years. To our knowledge, empirical investigation of real exchange rate persistence using black market rates of this length has not been previously undertaken.[11] Thus, this study extends the test of PPP into a new direction.

Black market real exchange rates are relative to the US dollar; hence they shed light on the issue of 'survivorship' bias. In this data set relative price ratios are less volatile than the nominal exchange rates (see Section 3). This addresses the concern raised by Lothian and Taylor. Mean reversion is tested by implementing unit-root tests in a heterogeneous panel framework as proposed by Im *et al.* (1997). This test is more powerful than alternative panel tests. The panel unit-root test is supplemented by a variance ratio test (Cochrane, 1988). In order to allow for a possible difference in the mean reverting behaviour of real exchange rates between pre- and post-1973 periods, we test separately for the pre-float and float sub-periods in addition to the entire period. Structural breaks in real exchange rates are formally tested through a sequential approach which identifies break dates endogenously and, where identified, unit-root tests are carried out allowing for the break.

One of the stylized facts in the literature is that the real exchange rate behaves differently under alternative (fixed versus floating) nominal exchange rate regimes. Among others, Mussa (1986) reports excess real exchange rate volatility under floating rates compared to fixed rates. However, Grilli and Kaminsky (1991) show that nominal exchange rate regimes may not be so critical to the behaviour of real exchange rates. We examine this issue by computing (1) the volatility of real exchange rates and (2) the Wald–Wolfowitz run tests. The latter test allows us to compare whether the real exchange rates between pre-float and float periods descend from an identical population; this provides additional insight into the behaviour of the real exchange rate across the two regimes.

The rest of the paper is organised as follows. Section 2 briefly describes the PPP hypothesis; Section 3 provides an overview of data; Section 4 discusses real exchange rate volatility and reports Wald-Wolfowitz tests; Section 5 discusses heterogeneous panel unit–root tests; Section 6

[9] When similar data sets (exhibiting high volatility in the price ratio) were used Taylor and McMahon (1988) also find mean reverting real exchange rates between currency pairs of industrial countries.

[10] In developing countries official exchange rates in general are controlled by the authorities whereas black market rates are fully market determined. Nevertheless they may not be the exact float counterparts of developing countries since a portion of foreign currency demand is met through the official transactions. It is argued that black market rates are more realistic (possibly reflecting the true value of domestic currency concerned) than official rates in these economies. In this sense these rates can be treated as proxies of the developing country floats. The International Monetary Fund's policy prescription for exchange rate liberalisation in developing countries is motivated by this very fact. Agenor (1992) provides an extensive survey of the black market literature and indicates that the volume of foreign exchange transactions in these markets tends to be a high proportion of the total volume of trade of these countries.

[11] Bahamani-Oskooe (1993) tests PPP using Iranian black market data. Our data span is much longer and the country coverage and approach are different. For other aspects of black market exchange rates see Fishelson (1988), Kiguel and O'Connel (1995) and Agenor and Taylor (1993).

provides variance ratio tests; Section 7 deals with the tests for structural break; and Section 8 summarizes the main findings and offers some conclusions.

2. PURCHASING POWER PARITY

Under absolute PPP the nominal exchange rate is proportional to a ratio of domestic to foreign price levels. In a logarithmic form:[12]

$$s_t = p_t - p_t^* \tag{1}$$

where s_t denotes the nominal exchange rate expressed as the home currency price of a unit of foreign currency, p_t the domestic price level, p_t^* the foreign price level and the t subscript denotes time. Several theoretical models e.g. the open economy version of the quantity theory (Lucas, 1982), the monetary model of exchange rate determination[13] and literature on target zones are based on the assumption that PPP holds continuously. However, the sticky price model (Dornbusch, 1976) allows for transitory divergence of nominal exchange rates from PPP. The well-known complications of transportation costs, trade barriers, taxation, exchange market intervention, and index number comparisons also generate deviations in the actual exchange rate away from its PPP level. Relative PPP is defined as:

$$\Delta s_t = \Delta p_t - \Delta p_t^* \tag{2}$$

Relative PPP states that the percentage change in the nominal exchange rate should be equal to the inflation differential between home and foreign country. The real exchange rate (q_t) is given by:

$$q_t = s_t - p_t + p_t^* \tag{3}$$

From equations (1) and (3) it is evident that the real exchange rate represents a deviation from purchasing power parity. If PPP holds continuously then the real exchange rate should either be zero or a fixed constant reflecting differences in the units of measurement.

Basically three specifications are common in the recent literature of PPP testing. These are: (1) a trivariate relationship between the nominal exchange rate, the domestic price level and the foreign price level; (2) a bivariate relationship between the nominal exchange rate and the domestic to foreign price ratio; and (3) a univariate analysis of the real exchange rate. A typical trivariate specification can be postulated as:

$$s_t = \alpha + \beta_1 p_t + \beta_2 p_t^* + u_t \tag{4}$$

where α and β_i ($i = 1, 2$) are regression parameters and u_t is an error term. PPP holds if the joint restriction $\beta_1 = -\beta_2 = 1$ is not rejected.[14] Equation (4) is the most general specification since it

[12] This derivation of absolute PPP relies on the assumption that the price of each traded good at home and in the foreign country carries equal weight and that they are summable across all traded goods. For further details see Officer (1976) and Rogoff (1996).

[13] On this see the collection of papers in Frenkel and Johnson (1978).

[14] Whether it is absolute PPP or relative PPP depends on price data used. If price data are expressed in levels then it would be absolute PPP. On the other hand, if price indices are used as regressors then it would be relative PPP. For more on this, see Crownover *et al.* (1996).

does not impose any of these restrictions *a priori*. Imposing symmetry on the price coefficients in equation (4) gives rise to a bivariate specification:

$$s_t = \alpha + \beta(p_t - p_t^*) + u_t \tag{5}$$

In equation (5) PPP would be confirmed if $\beta = 1$. Finally, the third specification directly examines the time series behaviour of the real exchange rate as given by equation (3).

Recent studies which choose either equation (4) or equation (5) test for a cointegrating relationship. A finding of a cointegrating relation is interpreted as evidence in favour of PPP since it implies that s, p, and p^* in equation (4) and s and $p - p^*$ in equation (5) show a stationary long-run relationship. Breuer (1994) and Edison *et al.* (1994), however, argue that a test of cointegration in equations (4) or (5) is not really a test of purchasing power parity since it does not impose the proportionality implied by the PPP theorem. A cointegrating relationship does not require cointegrating parameters (β_1 and β_2) to be proportional, $[1 \quad -1]$. The real exchange rate approach, on the other hand, imposes the restrictions of symmetry and proportionality implied by PPP. A stationary real exchange rate implies that PPP holds in the long run, i.e. deviations from PPP are transitory. In this study we follow this approach and analyse the behaviour of black market real exchange rates.[15]

3. AN OVERVIEW OF THE DATA

The monthly series on black market nominal exchange rates for the sample currencies are extracted from the various issues of *Pick's Currency Year Book* and *World Currency Year Book*. The sample period is 1958:1–1989:6.[16] We split the sample, the sub-period beginning April 1973 covering the modern float. The sample consists of 378 data points for the entire period; 183 for the pre-float and 195 for the float sub-periods.[17] Nominal exchange rates are units of domestic currencies per unit of US dollar. The monthly consumer price indices, collected from various issues of International Financial Statistics, are used to derive the real rates.[18] All real exchange rate series are plotted in Figure 1.

Two features are common in these plots. First, all real exchange rates are trending. Second, plots are far from being smooth, which indicates the volatility of black market real exchange rates. All rates (except Kyat) appreciate following the first oil shock and then revert gradually. In particular, IRupee, Ringgit, PRupee, Bhat and TDollar show sharp appreciation in the

[15] There is a lack of unanimity on the issue of whether one should impose proportionality and symmetry on the data set and test for PPP using the real exchange rate or whether the tests should be conducted in an unrestricted cointegrating framework. Breuer (1994), Edison *et al.* (1994) and Papell (1997) argue for the real exchange rate approach, whereas Taylor (1988), Taylor and McMahon (1988) and Cheung and Lai (1993) argue that the symmetry and proportionality restrictions may not be data consistent. Patel (1990) shows that the weights used in the construction of price indices across countries must be identical for these restrictions to be consistent. We acknowledge these contributions; however, it is often difficult to find ideal measures of variables which conform to the theoretical postulate. One can perhaps argue that even if one aims at the prices of a few homogeneous goods the quote price and the transaction price could differ significantly in the real world. In this study, we follow the real exchange rate approach.

[16] Data for Taiwan start from 1959:1.

[17] In this paper pre-float and Bretton Woods sub-periods are used interchangeably. We use float sub-period and post-Bretton Woods likewise.

[18] Tests of PPP using both the consumer price index (*cpi*) and the wholesale price index (*wpi*) are not uncommon. Diebold *et al.* (1991) point out that the *cpi*-based real exchange rate represents the relative price of a consumption basket whereas the *wpi* based represents the relative price of a production basket. Lack of data prevented us from using the wholesale price index. The *cpi* indices are converted to base year 1990 = 100.

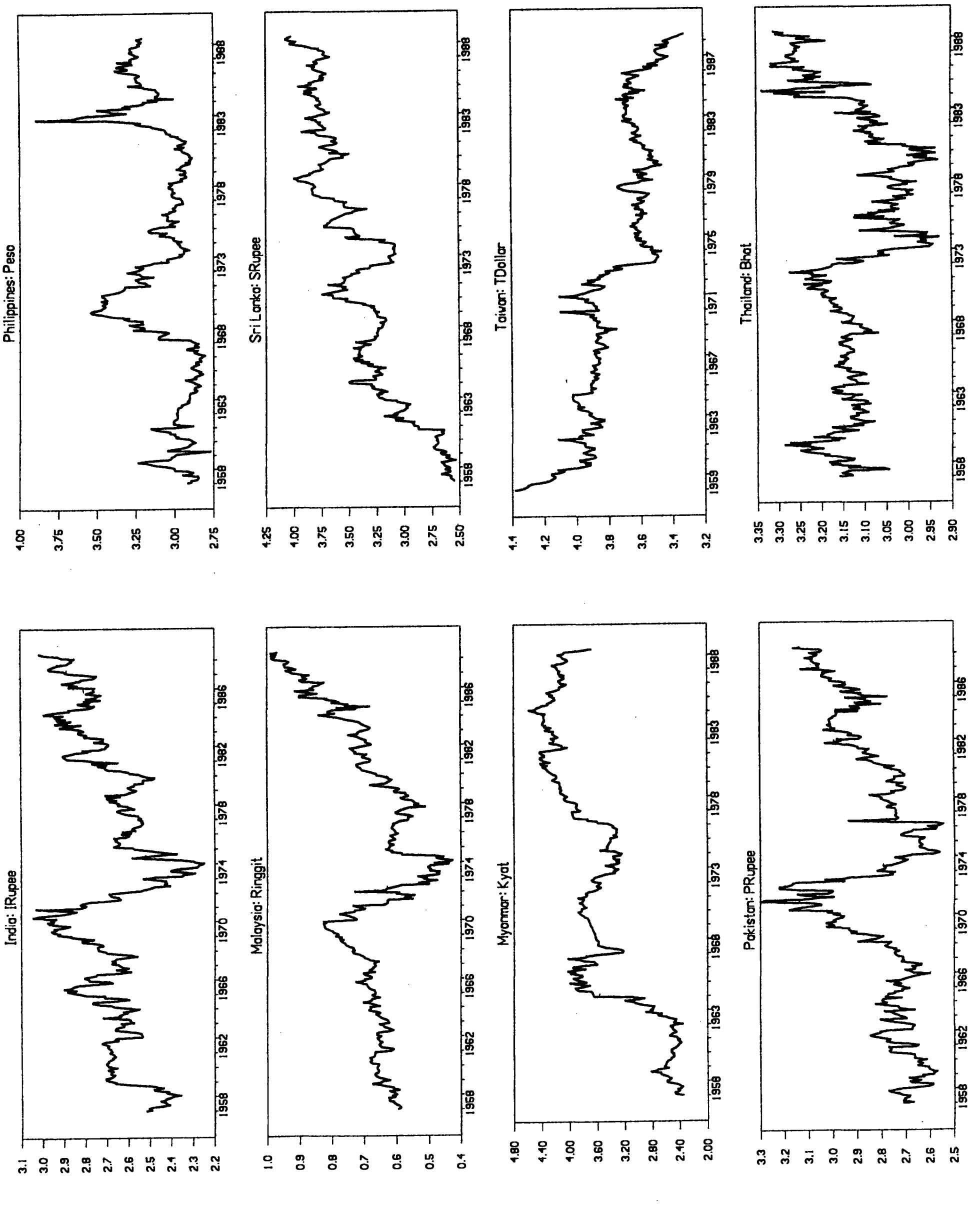

Figure 1. Plots of real exchange rates

aftermath of the first oil shock; Peso and SRupee appreciate relatively moderately, whereas Kyat remains virtually unchanged. It is also noteworthy that most currencies went through a period of gradual depreciation *vis-à-vis* the US dollar during the mid to late 1960s, a trend which reappeared in the 1980s. TDollar is the only currency which appreciated against the US dollar in the sample period; the rest depreciated over time. Peso shows a blip (sharp devaluation) in 1982. These plots indicate segmented trends. However, formal tests (reported below) reveal far fewer structural breaks than there appear visually. Another interesting aspect of these plots is that there are no discernible patterns of volatility between pre- and post-1973 periods.

4. EXCHANGE RATE REGIMES AND REAL EXCHANGE RATE VOLATILITY

Table I reports the sample mean and standard deviations of (1) the monthly absolute rate of change of the real and nominal exchange rates and relative price ratios, and (2) the monthly rate of change of real exchange rates. We find little evidence of excess volatility of real exchange rates during the recent float. Three rates (PRupee, SRupee, and TDollar) exhibit identical magnitudes of volatility (mean value of the absolute rate of change of the real exchange rate) between pre- and post-1973 periods; one (Kyat) shows a reduction; and the remaining four rates exhibit very similar volatility across two regimes (the largest increase in volatility, which is 1·56 times, is recorded for Bhat).

We formally tested if the mean values of the absolute rate of change of real exchange rates were statistically the same across fixed and floating regimes.[19] Even though the magnitudes were small, tests revealed that for IRupee, Ringgit and Bhat the increments were statistically significant; for Kyat the decline was significant; and for the remaining four rates volatility across the two regimes were the same statistically. Thus, the majority of black market real exchange rates (five out of eight) did not exhibit a rise in volatility during the recent float. Grilli and Kaminsky (1991), using the same measure of volatility, report a rise of at least four to five times in the volatility of the pound–USdollar real exchange rate during the current float compared to the Bretton Woods period. Our results do not concur with theirs and indicate that real exchange rates of poor countries may be less volatile than their rich counterparts.

Mussa (1986) reported that the variance of the major industrialized countries' real exchange rates increased by eight to eighty times during current float compared to the Bretton Woods period. In order to make our results comparable to Mussa's we also report the standard deviation of the rate of change of real exchange rates. Contrary to Mussa's finding, we find very similar volatility of real exchange rates across pre- and post-1973 periods. Kyat and TDollar show a reduced volatility; other rates show small increments during the float period; the largest increment is experienced by Bhat (1·52 times) which is nowhere near that reported by Mussa.[20] Overall, black market real exchange rates do not exhibit the extent of increase in volatility reported for the currency pairs of industrialized countries during the recent float.

[19] The test is performed by calculating the standard error of the mean (absolute growth rate): $\sigma_\mu = S/\sqrt{T}$; where s denotes the standard error of the absolute growth rate and T is the sample size.

[20] We also computed the standard deviation of real exchange rates in levels and found that they provide an even stronger picture. Five rates (Kyat, PRupee, Peso, SRupee and TDollar) exhibit a reduced standard deviation during the float period; the remaining rates show a small increase in their standard deviations. These results are not reported here to conserve space. However, they are available on request.

Table I. Exchange rates and relative price volatility

	$\|\Delta rex\|$		$\|\Delta ex\|$		$\|\Delta pr\|$		Δrex	
	μ	σ	μ	σ	μ	σ	μ	σ
IRupee:								
Full-sample	0·033	0·030	0·031	0·030	0·009	0·007	0·001	0·0443
Bretton Woods	0·031	0·031	0·030	0·030	0·009	0·007	0·000	0·0441
Float	0·034	0·029	0·033	0·030	0·008	0·007	0·003	0·0447
Ringgit:								
Full-sample	0·015	0·018	0·013	0·018	0·006	0·006	0·010	0·024
Bretton Woods	0·013	0·018	0·008	0·017	0·008	0·008	0·002	0·022
Float	0·015	0·018	0·013	0·018	0·006	0·006	0·001	0·024
Kyat:								
Full-sample	0·049	0·056	0·040	0·049	0·022	0·038	0·004	0·074
Bretton Woods	0·056	0·069	0·046	0·055	0·024	0·049	0·007	0·088
Float	0·042	0·041	0·035	0·043	0·020	0·023	0·000	0·059
PRupee:								
Full-sample	0·031	0·035	0·029	0·034	0·009	0·009	0·001	0·047
Bretton Woods	0·031	0·034	0·030	0·033	0·009	0·008	0·003	0·046
Float	0·031	0·037	0·028	0·036	0·009	0·011	0·000	0·049
Peso:								
Full-sample	0·032	0·043	0·027	0·041	0·011	0·017	0·001	0·054
Bretton Woods	0·031	0·038	0·024	0·036	0·014	0·021	0·002	0·049
Float	0·033	0·048	0·031	0·046	0·009	0·012	0·000	0·058
SRupee:								
Full-sample	0·037	0·041	0·033	0·041	0·013	0·021	0·004	0·056
Bretton Woods	0·037	0·041	0·037	0·041	0·005	0·005	0·004	0·055
Float	0·037	0·042	0·029	0·041	0·021	0·027	0·004	0·056
TDollar:								
Full-sample	0·029	0·027	0·025	0·025	0·011	0·012	− 0·003	0·039
Bretton Woods	0·029	0·031	0·023	0·029	0·012	0·012	− 0·003	0·042
Float	0·029	0·023	0·026	0·021	0·009	0·013	− 0·003	0·036
Bhat:								
Full-sample	0·023	0·023	0·019	0·022	0·009	0·012	0·001	0·032
Bretton Woods	0·018	0·017	0·010	0·012	0·013	0·015	0·000	0·025
Float	0·028	0·026	0·027	0·026	0·006	0·006	0·001	0·038

μ indicates the mean value; σ indicates standard error, and $|\Delta x|$ denotes the absolute rate of change· *rex*, *ex* and *pr* stand for real exchange rate, nominal exchange rate and relative price ratio, respectively. Δrex is the rate of change in the real exchange rate.

Now we turn to the volatility of relative price ratios. In this data set, the relative price ratios are less volatile than the nominal exchange rates: the mean value of the monthly absolute rate of change of the nominal exchange rate is always greater than that of the relative price ratio except in one count (Bhat shows a marginally higher volatility of relative price ratio compared to that of the nominal exchange rate during the pre-float). Given this low volatility of relative price ratios, our findings should provide new insights into the behaviour of PPP deviations for poor countries.

4.1 Wald–Wolfowitz Tests

The above analysis suggests little difference in the volatility of black market real exchange rates between Bretton Woods and post-Bretton Woods periods. This is an interesting finding in that it goes against what Grilli and Kaminsky (1991, p. 192–3) call one of the 'basic tenets of the international finance literature' that the behaviour of the real exchange rate essentially depends on nominal exchange rate regimes. Indeed, Grilli and Kaminsky also provide evidence which refutes the notion that nominal exchange rate regimes are crucial to the behaviour of real exchange rates.

In order to shed further light on this issue we implement the Wald–Wolfowitz (1940) test which discriminates between the underlying distributions of pre-float and float real exchange rates. This is a non-parametric test which tests the null hypothesis that two random samples (i.e. pre-float and float real exchange rates) descend from populations with an identical distribution against the alternative that they may differ in central tendency, dispersion, or skewness. Thus, it provides a broader insight into the characteristics of pre-float and float real exchange rates than is yielded by volatility alone.

Let n_1 and n_2 be observations from two independent random samples from two populations. The test can be implemented in the following steps: (1) form a single set of $n_1 + n_2$ observations by combining both random samples and arrange them in order of increasing magnitude; (2) generate runs based on the ordered sequence of $n_1 + n_2$ observations. Define R as the total number of such observed runs; the mean and variance of R are:

$$\mathrm{E}(R) = \frac{2n_1 n_2}{n_1 + n_2} + 1, \quad \text{and} \quad \mathrm{Var}\ (R) = \frac{2n_1 n_2(n_1 n_2 - n_1 - n_2)}{(n_1 + n_2)^2(n_1 - n_2 - 1)}$$

Wald and Wolfowitz show that

$$Z = \frac{R - \mathrm{E}(R)}{\sqrt{\mathrm{Var}\ (R)}} \sim N(0, 1)$$

hence Z can be used to test whether two independent samples follow identical populations in distribution.

We implement this test on pre- and post-1973 real exchange rates and the results are reported in Table II. Tests comfortably reject the null for all real exchange rates at the conventional level of significance. Thus, pre-float and float black market real exchange rates are different in terms of their population characteristics. This finding is similar to that reported in Grilli and Kaminsky

Table II. Wald–Wolfowitz test

Tests/currency	IRupee	Ringgit	Kyat	PRupee	Peso	SRupee	TDollar	Bhat
Wald–Wolfowitz	-2.455^{b}	-6.889^{a}	-15.447^{a}	-4.105^{a}	-5.755^{a}	-13.178^{a}	-8.437^{a}	-12.457^{a}

Under the null of identical distribution of populations the test statistic is approximately distributed as standard normal. The critical values at 1% and 5% are 2·58 and 1·96, respectively. Superscripts a and b indicate rejection of the null at 1% and 5%, respectively.

(1991) in that the inclusion of the Bretton Woods period in their study was crucial for the rejection of the null.

5. HETEROGENEOUS PANEL UNIT-ROOT TESTS

The panel approach substantially improves the power of the unit-root tests compared to those applied to single time-series. Recently, Im *et al.* (1997; hereafter IPS) have proposed unit-root tests for heterogeneous panels—the LM-bar and the T-bar tests—which are more powerful than Levin and Lin's (1993; hereafter LL) test which has received widespread application. Besides, the IPS tests have the following attractive features: (1) they allow for the heterogeneity of the dynamics and error variances across groups; (2) even if errors in different regressions contain common time specific components, the LM-bar and the T-bar tests based on cross-sectionally de-meaned regressions are valid; (3) they are consistent under a general class of fixed alternatives that allows for a fraction of individual groups to have a unit root, and this is more general than the alternative hypothesis of stationary across all groups which is implicitly tested under the LL approach; and finally (4) they may have better small sample properties since their asymptotic validity only requires $N/T \to k$ (k is any finite positive constant) when both N (cross-section) and T (time periods) tend to infinity compared to the more stringent condition, $N/T \to 0$, needed for the LI test. Consider a standard ADF equation in a panel framework.[21]

$$\Delta y_{it} = \mu_i + \beta_i y_{i,t-1} + \sum_{j=1}^{p_i} \rho_{ij} \Delta y_{i,t-j} + \varepsilon_{it} \quad i = 1, \ldots, N;\ j = 1, \ldots, p;\ t = 1, \ldots, T \tag{6}$$

where y_{it} is a stochastic process observed over N cross-sections and T time periods; $\Delta y_{it} = y_{it} - y_{it-1}$; ρ_{ij} are the coefficients associated with the pth-order augmentation which account for possibly heterogeneous error serial correlation across groups; and ε_{it} are assumed to be independently distributed normal variates with zero mean and finite (and possibly) heterogeneous variance, σ_i^2. The null of the unit root across all cross-sections (H_0: $\beta_i = 0$ for all i) is tested against the alternatives which allow a fraction of a cross-section unit to have unit roots (H_1: $\beta_i < 0$, $i = 1, 2, \ldots, N_1$, $\beta_i = 0$, $i = N_{1+1}, N_{1+2}, \ldots, N$). The individual LM statistic, $LM_{iT}(p_i, \rho_i)$, for testing $\beta_i = 0$ is given by:

$$LM_{iT}(p_i, \rho_i) = TR^2 \tag{7}$$

where R^2 is obtained from the regression: $\hat{\omega}_{it} = \varphi + \lambda_i y_{i,t-1} + \Sigma_{j=1}^{p_i} \varphi_{ij} \Delta y_{i,t-j} + \upsilon_{it}$.[22]. The LM-bar statistic across groups $\overline{LM}_{NT}$, is defined as:

$$\overline{LM}_{NT} = \frac{1}{N} \sum_{i=1}^{N} LM_{iT}(p_i, \rho_i) \tag{8}$$

The exact sample critical values are tabulated by IPS. In a finite sample the individual LM

[21] The methodological illustration draws heavily on Im *et al.* (1997). Tests based on the DF equations are available but we directly concentrate on the ADF equations since in our empirical model some DF equations exhibit serially correlated errors. A linear time trend can be included in the equation without loss of generality. Indeed, we report results with and without a time trend in the underlying DF/ADF equation.

[22] $\hat{\omega}_{it}$ is the residual series constructed from the auxiliary regression: $\Delta y_{it} = \delta_0 + \Sigma_{j=1}^{p_i} \pi_{it} \Delta y_{i,t-j} + \omega_{it}$.

statistic, $LM_{iT}(p_i, \rho_i)$, will depend on the nuisance parameter, ρ_i; a modified-standardized LM-bar statistic, $\Psi_{\overline{LM}}$, which performs better in a small sample is given by[23]

$$\Psi_{\overline{LM}} = \frac{\sqrt{N}\{\overline{LM}_{NT} - \frac{1}{N}\sum_{i=1}^{N} \mathrm{E}[LM_{iT}(p_i, 0) \mid \beta_i = 0]\}}{\sqrt{\frac{1}{N}\sum_{i=1}^{N} \mathrm{Var}[LM_{iT}(p_i, 0) \mid \beta_i = 0]}} \tag{9}$$

where E is the expectation operator, and $[LM_{iT}(p_i, 0) \mid \beta_i = 0]$ is the LM statistic for testing $\beta_i = 0$ in the ADF(p) regression (6). The mean and variance of $[LM_{iT}(p_i, 0) \mid \beta_i = 0]$ are tabulated by IPS via stochastic simulations. The analogous modified-standardized T-bar statistic, Ψ_t, which corresponds to equation (9) in testing $\beta_i = 0$ is defined as:

$$\Psi_{\bar{t}} = \frac{\sqrt{N}\{\bar{t}_{NT}(P, \rho) - \frac{1}{N}\sum_{i=1}^{N} \mathrm{E}[t_{iT}(p_i, 0) \mid \beta_i = 0]\}}{\sqrt{\frac{1}{N}\sum_{i=1}^{N} \mathrm{Var}[t_{iT}(p_i, 0) \mid \beta_i = 0]}} \tag{10}$$

where $\bar{t}_{NT}(P, \rho) = \frac{1}{N}\sum_{i=1}^{N} t_{iT}(p_i, \rho)$; $t_{iT}(p_i, \rho_i)$ is the individual Dickey–Fuller t-statistic for testing $\beta_i = 0$ in equation (6). Again, IPS tabulate the mean and variance of $[t_{iT}(p_i, 0) \mid \beta_i = 0]$. Although both the LM-bar and T-bar tests are more powerful than the LL test, the T-bar is shown to perform better than the LM-bar in small samples. An important feature of these tests is that their power is favourably affected by a rise in T compared to an equivalent rise in N. This feature is particularly important in our data set where T is very large compared to N.

When disturbances are correlated across groups the tests outlined above become inapplicable. The appropriate test statistics are LM-bar and T-bar tests based on de-meaned regressions. Assuming that the error term in equation (6) is $\varepsilon_{it} = \theta_t + \eta_{it}$, where θ_t is a time-specific common effect and η_{it} is independently distributed, the de-meaned regression, obtained by subtracting the mean from both sides of equation (6), is:

$$\Delta\tilde{y}_{it} = \tilde{\mu}_i + \beta_i\tilde{y}_{i,t-1} + \sum_{i=1}^{p_i}\gamma_i\Delta\tilde{y} + \tilde{\xi}_{it} \tag{11}$$

where '$\sim$' on top of variables indicates that they are (cross-sectional) mean deducted (see IPS for details). The de-meaned individual LM statistic, $\tilde{LM}_{iT}$, which tests the null $\beta_i = 0$ $(i = 1, 2, \ldots, N)$ in equation (11) is obtained following the TR^2 procedure outlined above. The de-meaned LM-bar statistic, $\tilde{LM}_{NT}$, is defined as[24]

$$\bar{\tilde{LM}}_{NT} = \frac{1}{N}\sum_{i=1}^{N}\tilde{LM}_{iT} \tag{12}$$

[23] As $T \to \infty$, $\Psi_{\overline{LM}}$ is shown to converge to standard normal provided that ε_{it} are independently distributed normal variates with zero mean and finite (possibly heterogeneous) variance.

[24] It should be noted, however, that this de-meaning procedure will not be robust if the effect of the common component varies across groups. This point is acknowledged by Im *et al.* (1997).

Table III. Augmented Dickey–Fuller test

	Entire sample (958:1–1989:6) t_t	Bretton Woods (1958:1–1973:3) t_t	Post-Bretton Woods (1973:4–1989:6) t_t
IRupee	−2·696(0)	−2·065(0)	−3·288(0)
$Q(8)$	0·766	0·819	0·592
Ringgit	−1·469(2)	−2·120(1)	−3·397(2)
$Q(8)$	0·148	0·268	0·688
Kyat	−1·528(0)	−1·714(0)	−2·001(0)
$Q(8)$	0·749	0·856	0·795
PRupee	−3·354(0)	−1·987(2)	−5·030(0)[a]
$Q(8)$	0·514	0·340	0·667
Peso	−3·094(0)	−2·068(0)	−2·648(1)
$Q(8)$	0·140	0·378	0·233
SRupee	−3·221(0)	−2·114(0)	−2·136(0)
$Q(8)$	0·373	0·828	0·472
TDollar	−4·246(0)[a]	−4·333(0)[b]	−3·720(1)
$Q(8)$	0·155	0·468	0·278
Bhat	−2·440(1)	−3·361(1)	−3·169(1)
$Q(8)$	0·264	0·128	0·857

The estimated model is $\Delta q_t = \mu + \gamma t + \phi_1 q_{t-1} + \Sigma_{i=1}^{p} \Delta q_{t-1} + \mu_t$. Superscripts a and b indicate rejection of the null at 1% and 5%. The 1% and 5% critical values are −3·99 and −3·43, respectively. Unit root results without time trends are qualitatively the same. $Q(8)$ are the 8th-order Ljung–Box tests for serial correlation; P-values under the null of no serial correlation are reported. The ADF tests are reported only when the DF residuals exhibit serial correlation. The upper lag-length of the ADF equation are specified following Campbell and Perron (1991) and insignificant ones are dropped following Engle *et al.* (1993). The latter argue for dropping the insignificant lags to preserve the power of unit root tests. The returns of the black market real exchange rates are all stationary (results are available on request).

Likewise, the de-meaned T-bar statistic is obtained by averaging the individual ADF t-ratios estimated using equation (11).[25]

Prior to the panel unit-root tests, we report the standard DF/ADF tests in Table III. The low power of Dickey–Fuller tests (Dickey and Fuller, 1979, 1981) is well known. However, the motivation here is to see if the standard non-rejection of a unit root in the real exchange rate shown elsewhere by these tests stands in this data set as well.

Indeed, the overall picture from the DF/ADF tests is the standard non-rejection of a unit root in real exchange rates. Out of the 24 test cases (eight real exchange rates over three sample periods), the null of a unit root is rejected on only three occasions: the TDollar over the full sample and Bretton Woods sub-sample and the PRupee for the post-Bretton Woods period. The

[25] The cross-sectional de-meaning introduces dependence across the errors of de-meaning regressions, $\hat{\xi}_{it}$. However, the test statistic, $\bar{\tilde{LM}}_{NT}$, remains asymptotically valid if the idiosyncratic effects, ν_{it} are serially uncorrelated. The standardized de-meaned LM-bar statistic,

$$\Gamma_{\tilde{LM}} = \frac{\sqrt{N}\{\bar{\tilde{LM}}_{NT} - \mathrm{E}(\eta_T)\}}{\sqrt{\mathrm{Var}(\eta_T)}}$$

converges to a standard normal in the limit (for details see Im *et al.* (1997)).

rest are all non-stationary. The Ljung–Box tests indicate absence of error serial correlation in the reported equations.

As a precursor to the panel unit-root tests, we tested for the contemporaneous correlation among DF/ADF residuals across groups by employing the Lagrange Multiplier test, λ_{LM}, due to Breusch and Pagan (1980). The test statistic is given by:

$$\lambda_{LM} = T\sum_{i=2}^{n}\sum_{j=1}^{i-1} r_{ij}^2 \tag{13}$$

where r_{ij}^2 is the squared ijth correlation coefficient of residuals. Under the null of no correlation across group residuals the test statistic, λ_{LM}, is asymptotically χ^2 distributed with $N(N-1)/2$ degrees of freedom. The computed $\hat{\lambda}_{LM}$ from the underlying DF and ADF(2) regressions are 555·895 and 544·680 for the full sample, 201·366 and 200·270 for the pre-float, and 306·629 and 172·838 for the float period. For 28 degrees of freedom, the 1% critical value for the chi-squared distribution is 48·3. Thus, the $\hat{\lambda}_{LM}$ statistics are highly significant and reject the null in all cases. This finding is robust to higher-order augmentations [ADF(4)].

Significant error correlation across groups implies that the modified-standardised LM-bar and T-bar tests are not valid; instead the de-meaned LM-bar and T-bar are the appropriate tests. We report the results of the de-meaned LM-bar and T-bar tests in Table IV. The first row shows results based on the de-meaned DF regressions; those on the second row are based on the augmentation structure that accounts for the few cases of residual serial correlation found in the underlying de-meaned DF equations (see footnote to Table IV). The de-meaned T-bar statistics reject the null of a unit-root in all cases at a very high level of precision. The de-meaned LM-bar does the same in all cases but one: it cannot reject the null for the Bretton Woods period when the linear time trend is excluded from the underlying regressions. It is important to note that the T-bar test, which has better power performance than the LM-bar test, provides a clear rejection of non-stationary panels. Thus, the overall finding is that our panels of black market real exchange rates are stationary, which supports PPP.

Table IV. Unit root tests in heterogeneous panels

	Bretton Woods		Post-Bretton Woods		Full Sample	
	Constant only	Constant and trend	Constant only	Constant and trend	Constant only	Constant and trend
LM-bar:	3·855	7·502[b]	6·223[a]	7·923[b]	7·241[a]	10·700[a]
	3·775	#	5·929[a]	7·562[b]	6·361[a]	10·017[a]
T-bar:	−1·913[a]	−2·689[a]	−2·332[a]	−2·611[a]	−2·632[a]	−3·237[a]
	−1·859[a]	#	−2·260[a]	−2·526[a]	−2·513[a]	−2·993[a]

Results in the first row are based on the de-meaned DF regressions; those on the second row are based on the de-meaned DF/ADF regressions which account for the error serial correlation. The currency names and the required augmentations(.) are as follows. For the full sample: Ringgit(1) and Bhat(1). For the post-Bretton Woods period: Bhat(1). For the Bretton Woods period: Ringgit(1) and Bhat(1) only when the linear time trend is excluded. # implies no serial correlation in the DF equation and hence no results based on augmentation are reported. Appropriate critical values for all the exact tests were kindly made available by Kyung So Im. To save space we do not report these statistics but they are available on request. Superscripts a and b indicate rejection of the null at 1% and 5%, respectively.

IPS emphasizes the correct choice of the order of the underlying ADF regressions since it affects both the power and the size of these tests. In effect, there is a trade-off; underestimation of order causes the sizes of all tests to go to zero whereas over-fitting the model reduces their power. In order to allow for a possible under-fitting, tests were carried out by estimating uniform ADF(2) regressions and the results remain qualitatively the same.[26]

6. VARIANCE RATIO TEST

Cochrane (1988) introduced this test and implemented it to examine the degree of persistence in US real GNP.[27] Glen (1992) and Grilli and Kaminsky (1991) use the variance ratio to examine the persistence in real exchange rates. The variance ratio (V^k) is given by:

$$V^k = \frac{\mathrm{Var}(y_t - y_{t-k})}{k * \mathrm{Var}(y_t - y_{t-1})} \tag{14}$$

where k is the lag length, $\mathrm{Var}(y_t - y_{t-k})$ and $\mathrm{Var}(y_t - y_{t-1})$ are the variances of the kth difference and the first difference of a time series (y_t). The intuition behind the variance ratio test is as follows. For any stationary series, the variance of the k lagged differences approaches twice the variance of the series. If a series follows a random walk, then the variance of its k differences grows linearly with k such that: $\mathrm{Var}(y_t - y_{t-k}) = k * \mathrm{Var}(y_t - y_{t-1})$. Therefore, a test of random walk is equivalent to testing the null hypothesis that $1/k$ times the variance of the k differences over the variance of the first difference, i.e. the variance ratio (V^k), is equal to one. However, the estimated V^k have been shown to converge to $(T-k)/T$ rather than one.[28] The estimate of V^k is consistent; nevertheless it is subject to small sample bias. Cochrane suggests a degrees-of-freedom correction in small samples which is achieved by multiplying V^k by a factor, $T/(T-k+1)$.[29] The asymptotic standard error of V^k is given by $V^k/(0{\cdot}75 * T/k + 1)^{0{\cdot}5}$. The long difference ($k$) is set sufficiently large as to minimize the bias resulting from excluding the higher-order autocorrelation. However, k should not be unduly large since V^k is zero for $k = T-1$.

Table V reports the estimates of Cochrane's degrees of freedom adjusted variance ratio (V^k) with their asymptotic standard errors and random walk benchmarks. Panel A contains selected V^k for the entire period, and panels B and C for the two sub-periods. For the entire period, the estimated V^k for all currencies are consistently below their random walk benchmarks with the sole exception of Kyat. They are all significantly different from zero for k ranging from 1 to 70. Interestingly, the cut-off lag length is 70 for all the seven real exchange rates.[30] Further, the variance ratios for PRupee, TDollar and Bhat are significantly less than unity whereas for IRupee, Peso, Ringgit and SRupee they are not statistically different from unity. Gucht *et al.* (1996) and Huizinga (1987) emphasize the overall pattern of the successive estimates of these

[26] When ADF(4) are estimated across groups, results still remain qualitatively the same for the entire period but there are fewer rejections of the unit root null for the sub-samples. The latter non-rejections may be attributed to the loss of power due to over-fitting, a concern voiced by IPS. For the sake of brevity we do not report these results but they are available on request.

[27] Lo and MacKinlay (1988) also develop a similar variance ratio test.

[28] $(T-k)/T$ is known as the random walk benchmark, where T is the sample size.

[29] It is shown that $V^k = 1 + 2\Sigma_{j=1}^{k}[1 - j/(k+1)]\rho_j$, where ρ_j is the jth autocorrelation of ΔS_t. Thus, V^k can also be estimated by replacing the ρ_j by the sample autocorrelations. For a stationary process and a random walk process both measures of persistence produce the same results.

[30] To save space not all the variance ratios are reported. However, the complete set of results indicate this.

Table V. Variance ratio test

K (months)	Variance ratio (V^k) IRupee	Ringgit	Kyat	PRupee	Peso	SRupee	Tdollar	Bhat	RW
Panel A: Entire period (1958:1–1989:6)									
1	1	1	1	1	1	1	1	1	0·997
	(0·084)	(0·084)	(0·084)	0·084)	(0·084)	(0·084)	(0·084)	(0·084)	
5	0·953	0·732	0·983	0·647	0·801	0·890	0·712	0·516	0·987
	(0·139)	(0·106)	(0·143)	(0·094)	(0·116)	(0·129)	(0·104)	(0·075)	
10	0·809	0·633	0·894	0·496	0·791	0·876	0·613	0·389	0·974
	(0·159)	(0·125)	(0·179)	(0·098)	(0·156)	(0·173)	(0·121)	(0·077)	
15	0·901	0·572	0·998	0·479	0·796	0·881	0·616	0·370	0·960
	(0·214)	(0·136)	(0·237)	(0·114)	(0·189)	(0·209)	(0·146)	(0·088)	
25	0·842	0·682	1·148	0·556	0·699	0·804	0·535	0·323	0·934
	(0·255)	(0·206)	(0·348)	(0·168)	(0·212)	(0·244)	(0·162)	(0·098)	
50	0·584	0·629	1·149	0·508	0·638	0·500	0·411	0·282	0·868
	(0·248)	(0·267)	(0·487)	(0·216)	(0·271)	(0·212)	(0·174)	(0·120)	
100	0·510	0·718	0·870	0·334	0·386	0·358	0·261	0·295	0·735
	(0·305)	(0·429)	(0·519)	(0·199)	(0·231)	(0·214)	(0·156)	(0·176)	
125	0·434	0·667	0·533	0·271	0·320	0·308	0·242	0·252	0·669
	(0·289)	(0·445)	(0·356)	(0·181)	(0·213)	(0·205)	(0·161)	(0·168)	
150	0·407	0·657	0·209	0·259	0·203	0·212	0·128	0·204	0·603
	(0·297)	(0·480)	(0·153)	(0·189)	(0·148)	(0·155)	(0·094)	(0·149)	
190	0·129	0·324	0·308	0·078	0·132	0·200	0·157	0·107	0·497
	(0·106)	(0·266)	(0·252)	(0·064)	(0·108)	(0·164)	(0·129)	(0·088)	
Panel B: Pre-Bretton Woods period (1958:1–1973:3)									
1	1	1	1	1	1	1	1	1	0·995
	(0·121)	(0·121)	(0·121)	(0·121)	(0·121)	(0·121)	(0·121)	(0·121)	
5	0·924	0·633	0·934	0·613	0·815	0·785	0·661	0·503	0·973
	(0·193)	(0·132)	(0·195)	(0·128)	(0·170)	(0·164)	(0·143)	(0·105)	
10	0·796	0·558	0·878	0·390	0·815	0·782	0·545	0·362	0·945
	(0·225)	(0·158)	(0·247)	(0·110)	(0·231)	(0·221)	(0·160)	(0·102)	
15	0·884	0·569	1·037	0·354	0·868	0·738	0·545	0·324	0·918
	(0·330)	(0·194)	(0·354)	(0·121)	(0·296)	(0·252)	(0·193)	(0·111)	
25	0·747	0·602	1·278	0·364	0·901	0·580	0·386	0·261	0·863
	(0·325)	(0·262)	(0·556)	(0·159)	(0·392)	(0·252)	(0·174)	(0·114)	
50	0·173	0·232	0·174	0·389	0·887	0·482	0·315	0·148	0·727
	(0·105)	(0·141)	(0·715)	(0·237)	(0·541)	(0·296)	(0·199)	(0·090)	
60	0·139	0·197	0·988	0·387	0·879	0·534	0·189	0·123	0·672
	(0·093)	(0·131)	(0·685)	(0·258)	(0·586)	(0·356)	(0·131)	(0·082)	
75	0·303	0·249	0·852	0·294	0·778	0·492	0·185	0·107	0·590
	(0·225)	(0·185)	(0·634)	(0·218)	(0·579)	(0·366)	(0·143)	(0·080)	
90	0·334	0·249	0·550	0·239	0·624	0·385	0·267	0·151	0·508
	(0·272)	(0·203)	(0·448)	(0·195)	(0·508)	(0·314)	(0·225)	(0·123)	
Panel C: Post-Bretton Woods period (1973:4–1989:6)									
1	1	1	1	1	1	1	1	1	0·994
	(0·117)	(0·117)	(0·117)	(0·117)	(0·117)	(0·117)	(0·117)	(0·117)	
5	0·926	0·759	1·119	0·600	0·811	0·939	0·792	0·527	0·974
	(0·187)	(0·154)	(0·227)	(0·121)	(0·164)	(0·190)	(0·160)	(0·107)	
10	0·751	0·597	0·975	0·385	0·788	0·856	0·718	0·378	0·949
	(0·206)	(0·164)	(0·267)	(0·106)	(0·216)	(0·235)	(0·197)	(0·104)	

Table continued over page

Table V. Continued

K (months)	Variance ratio (V^k) IRupee	Ringgit	Kyat	PRupee	Peso	SRupee	Tdollar	Bhat	RW
15	0·735	0·434	0·983	0·391	0·767	0·833	0·599	0·311	0·923
	(0·243)	(0·144)	(0·325)	(0·129)	(0·254)	(0·276)	(0·198)	(0·103)	
25	0·401	0·393	0·966	0·388	0·558	0·668	0·463	0·195	0·872
	(0·169)	(0·166)	(0·407)	(0·164)	(0·235)	(0·282)	(0·195)	(0·082)	
50	0·314	0·184	1·276	0·143	0·401	0·532	0·402	0·216	0·744
	(0·186)	(0·109)	(0·753)	(0·085)	(0·237)	(0·314)	(0·238)	(0·127)	
65	0·130	0·169	1·324	0·129	0·381	0·542	0·279	0·188	0·667
	(0·088)	(0·113)	(0·889)	(0·087)	(0·256)	(0·364)	(0·187)	(0·126)	
80	0·175	0·200	1·383	0·197	0·301	0·314	0·190	0·223	0·590
	(0·130)	(0·153)	(1·029)	(0·147)	(0·223)	(0·234)	(0·141)	(0·166)	
100	0·119	0·229	1·362	0·166	0·195	0·464	0·252	0·149	0·487
	(0·099)	(0·191)	(1·132)	(0·138)	(0·162)	(0·385)	(0·209)	(0·124)	

The reported V^k range from first lag to a half of the sample. For the entire period the half sample size value of k is approximated by $k = 190$. For the pre-float and the float sub-periods they are 90 and 100 months, respectively. The asymptotic standard errors are in parentheses. RW stands for random walk benchmark.

ratios rather than their individual significance. From this perspective seven of the real exchange rates are stationary. Their variance ratios gradually decline towards zero as k increases showing a typical behaviour of trend stationarity. Beyond the lag length of 70, standard errors become large and statistically all V^k are zero.

For Kyat the V^k exceeds the random walk benchmark beginning with the 15th lag and remains so for a considerable length of time (15–110 months) and then falls below the random walk benchmark. Thus, it requires long differences to identify the mean reverting behaviour of Kyat. Variance ratios of Kyat are not significantly different from unity from k ranging from 1 to 70 and they are all insignificant thereafter.

The two sub-periods show similar results. The same seven real exchange rates are stationary and their variance ratios exhibit a similar pattern for the pre-float sub-period. They are consistently below the random walk benchmark and are significantly different from zero for $k = 1$–35. All V^k for $k > 35$ are insignificant. PRupee, TDollar and Bhat indicate their variance ratios to be significantly less than unity whereas IRupee, Peso, Ringgit and SRupee indicate that they are not. The variance ratio for Kyat exceeds the random walk bench mark at the 15th lag and remains so throughout the sample period. Statistically it is not different from unity for k ranging from 1 to 35 and turns insignificant thereafter. The float sub-period shows qualitatively identical results.

Since V^k is identically zero for $k = T - 1$, Campbell and Mankiw (1987) and Gucht *et al* (1996) suggest that the values of V^k for k ranging from $T/3$ to $T/2$ are particularly interesting from the perspective of the random walk component. Following Gucht *et al.* we report V^k for $k = T/3$, $k = 0{\cdot}4T$, and $k = T/2$ to indicate the random walk component of each real exchange rate.[31] They are, respectively, approximated by V^{125}, V^{150} and V^{190} for the entire period, by V^{60}, V^{75} and V^{90} for the pre-float, and by V^{65}, V^{80} and V^{100} for the float sub-period. Judged from the midpoint estimates, the random walk component lies between a maximum of 0·657 (Ringgit) to a

[31] $0{\cdot}4T$ is the approximate midpoint of the interval $T/3$ and $T/2$. Cheung and Lai (1993) report V^k for $0{\cdot}6T$.

minimum of 0·128 (TDollar) for the entire period; 0·85 (Kyat) and 0·11 (Bhat) for the pre-float, and 1·383 (Kyat) and 0·157 (IRupee) for the float sub-period. The overall evidence from the variance ratio test is that all sample real exchange rates are mean reverting for the entire period though Kyat takes long differences to identify this behaviour. However, the magnitude of the random walk component is different across currencies. The initial shock is not magnified and tapers off gradually and smoothly. We do not find the humped pattern reported in Huizinga (1987) and Gucht *et al.* (1996).[32] It also suggests that the mean reverting behaviour of the real exchange rate is qualitatively the same between pre-float and float sub-periods except for one real exchange rate (Kyat).[33]

7. TESTS OF STRUCTURAL BREAK

The issue of whether there is a structural break in the real exchange rates is an important one and has been raised by several economists. The long-span studies which support PPP mix different exchange rates and/or policy regimes; hence the concern about the structural shift across regimes and its implications on the test results. We address this issue by (1) direct tests of a structural break in real exchange rates, and (2) a test for the unit-root conditional on the identified structural break. We apply sequential tests (Banerjee *et al.* 1992; Zivot and Andrews, 1992) which estimate the break-date endogenously. Consider the following models:

$$\Delta y_t = \mu_{10} + \mu_{11} d_t(\lambda)^* + \mu_{12} t + \alpha_1 y_{t-1} + \beta(L)\Delta y_{t-1} + e_{1t} \tag{15}$$

$$\Delta y_t = \mu_{20} + \mu_{21} d_t(\lambda) + \mu_{22} t + \alpha_2 y_{t-1} + \gamma(L)\Delta y_{t-1} + e_{2t} \tag{16}$$

$$\Delta y_t = \mu_{30} + \mu_{31} d_t(\lambda) + \mu_{32} d_t(\lambda)^* + \mu_{33} t + \alpha_3 y_{t-1} + \psi(L)\Delta y_{t-1} + e_{3t} \tag{17}$$

$$(t = 1, \ldots, T)$$

where $\lambda = k_0, k_0 + 1, \ldots, T - k_0$; k_0 is the initial start-up sample defined as $k_0 = \delta_0 T$, and δ_0 is the trimming parameter.[34] Model (15) captures the shift in the trend of the process, y_t; the dummy regressor, $d(\lambda)^* = (t - \lambda)1(t > \lambda)$, and 0 otherwise. Perron (1989, 1990) calls it a 'changing growth' model. Equation (16) specifies a shift in the mean of the process; the dummy regressor, $d_t(\lambda) = 1(t > \lambda)$, and 0 otherwise. This is Perron's 'crash' model. Banerjee *et al.* (1992) consider only these two cases. The equation (17) which allows simultaneous shifts in the level and slope of the trend function is considered by Zivot and Andrews (1992) and Christiano (1992).

The sequential approach involves estimating equations (15)–(17) using the full T observations for each possible break date, $\lambda = k_0, k_{0+1}, \ldots, T - k_0$. We compute two statistics. The first is the minimum sequential Augmented Dickey–Fuller t statistic, $\min t_\alpha = \min_{k0 \leqslant \lambda \leqslant T-k0} t_{\text{ADF}}(\lambda/T)$, for testing $\alpha_i = 0$ ($i = 1, 2, 3$) over the $T - 2k_0$ regressions. The no-break unit-root null is rejected if the min t_α exceeds (in absolute level) the corresponding critical value. Next, a sequence of F-statistics for testing the no-break null hypothesis (i.e. $\mu_{11} = 0$ in equation (15); $\mu_{21} = 0$ in equation (16); and $\mu_{31} = \mu_{32} = 0$ in equation (17)) against each of the three trend-break alternatives are computed. The $F(r, T - r - L - 3)$ statistics are computed sequentially with the

[32] Only Kyat shows a modest humped pattern.
[33] Results from the variance ratio test indicate that for Kyat the entire period seems to be long enough to capture its mean reverting behaviour, but not the sub-periods.
[34] Zivot and Andrews (1992) show lack of unique solution if end points are included.

numerator $r = 1$ for one parameter restriction associated with model (15) or (16) and $r = 2$ for model (17); the denominator equals the degrees of freedom of the unrestricted model. If the maximum of the F-statistics, $F_T^{\max} = \max_{k0 \leqslant \lambda \geqslant T-k0} F_T(\lambda/T)$, exceeds the corresponding critical value then the null is rejected. The estimated break point, $T_B = \hat{\lambda}$, corresponds to the date for which the F-statistic (t-statistic) is maximized (minimized) under the null $\alpha_i = 0$ against the alternative of $\alpha < 0$; $\hat{\lambda}/T$ gives the proportional location of the break date in the sample. For models (15) and (16) critical values are obtained from Banerjee *et al.* (1992); for model (17) Zivot and Andrews (1992) tabulate critical values for min t_α and Christiano (1992) tabulates for $F_T^{\max}$. Once the break dates are identified, the conditional unit-root tests are performed by estimating the following equation:

$$\Delta y_t = \mu_{40} + \mu_{41} d_t(\hat{\lambda}) + \mu_{42} d_t(\hat{\lambda})^* + \mu_{43} t + \alpha_4 y_{t-1} + \varphi(L)\Delta y_{t-1} + e_t, \quad t = 1, \ldots, T \tag{18}$$

The dummy variables are now indexed according to the estimated break date; the hat on the parameter (λ) indicates that the break date is endogenously identified. The null hypothesis of a unit root with one-time endogenous structural break corresponds to $\alpha_4 = 0$; the alternative hypothesis is one-time trend-break stationarity. The null is rejected if ADF$\hat{\tau}_t$ (for testing $\alpha_4 = 0$) is significant. The asymptotic critical values are provided by Zivot and Andrews (1992).

Empirical results of structural break tests are reported in Table VI. Following Banerjee *et al.* (1992), we set $\delta_0 = 0{\cdot}15$ and a lag length, $p = 4$. Results are robust to changes in the trimming parameter and lag lengths which are discussed below. There is little evidence of structural shifts in real exchange rates. Only two real exchange rates (Ringgit and Bhat) show significant mean shifts in their trend functions. Both test statistics ($F_T^{\max}$ and min t_α) reject the no-break unit-root null at the conventional level of significance. There is, however, no evidence of a shift in the slope of the trend function. Interestingly, the same break dates are estimated by either test statistic.[35]

The identified break dates are February 1972 for Ringgit and June 1973 for Bhat. The latter date coincides with the first oil shock and the demise of the Bretton Woods system whereas the break-date for Ringgit precedes these events. Plots of real exchange rates (Figure 1) show that Ringgit and Bhat experienced sharp appreciation during 1972–3 which is well picked up by our results as significant mean shifts. However, IRupee's sharp appreciation during 1972–3, PRupee's depreciation in the early 1970s followed by sharp appreciation around the first oil shock, Peso's sharp devaluation (blip) in 1982, and TDollar's appreciation during 1973–4 do not appear to be statistically significant turning points. Plots of real exchange rates for Kyat and SRupee appear relatively smooth and tests indicate no structural shifts.

The sensitivity of our results with respect to the lag length are examined by re-computing the statistics for p chosen by (a) the Akaike information criterion, (b) the Schwarz information criterion, and (c) setting $p = 8$. The Schwarz criterion chooses $p = 1$ in all cases. The Akaike criterion chooses $p = 3$ for PRupee and Bhat, $p = 9$ for Peso, and $p = 2$ for TDollar. For the remaining real exchange rates it also chooses a lag length of one for each. Results based on these wider lag length selections are qualitatively the same as those reported in Table VI with only one exception: now with $p = 1$, the min ι_t statistic turns out to be marginally insignificant for Ringgit although the $F_t^{\max}$ remains highly significant.

[35] We also applied alternative tests of structural breaks, i.e. the rolling and recursive tests, but did not find any evidence of structural breaks. Banerjee *et al.* (1992, pp. 279–80) point out that recursive and rolling tests have low power.

Table VI. Structural break tests

Currency	Sequential: trend shift model (15)		Sequential: mean shift model (16)		Sequential: mean and trend shift model (17)	
	$F_T^{\max}$	t_{minADF}	$F_T^{\max}$	t_{minADF}	$F_T^{\max}$	t_{minADF}
Irupee	1·470	−2·683	11·050	−3·988	10·649	−4·621
	[84:09]	[62:09]	[72:01]	[72:01]	[72:01]	[73:02]
Ringgit	4·373	−2·324	21·989[b]	−4·680[c]	15·120[b]	−5·249[b]
	[62:08]	[62:08]	[72:02]	[72:02]	[72:02]	[72:02]
Kyat	2·958	−1·929	8·139	−2·804	4·452	2·931
	[73:09]	[62:08]	[63:06]	[63:06]	[63:03]	[63:03]
Prupee	0·638	−2·405	8·813	−3·456	8·795	−3·392
	[72:02]	[70:02]	[73:05]	[73:05]	[73:05]	[73:07]
Peso	3·609	−2·804	9·601	−3·008	10·367	−3·964
	[83:11]	[81:06]	[83:11]	[83:02]	[80:08]	[68:04]
Srupee	0·288	−3·239	2·042	−3·513	2·171	−3·647
	[84:08]	[83:09]	[63:02]	[63:02]	[77:05]	[63:08]
Tdollar	3·916	−3·538	5·347	−3·579	7·136	−4·691
	[80:08]	[80:10]	[81:10]	[81:10]	[73:09]	[74:01]
Bhat	8·950	−3·270	26·270[b]	−5·332[b]	14·565[b]	−5·217[b]
	[62:08]	[62:08]	[73:06]	[73:06]	[73:06]	[73:06]
CV: 5%	16·04	−4·39	18·99	−4·78	12·00	−5·08
10%	13·20	−4·13	16·78	−4·51	11·20	−4·82

Dates corresponding to min t_α test and $F_T^{\max}$ test are reported within brackets [.]. In all cases the lag length, $p = 4$ and the trimming parameter, $\delta_0 = 0{\cdot}15$. CV denotes critical values obtained from Banerjee *et al.* (1992, tables 1–2) for trend shift (model (15)) and mean shift (model (16)) models. Critical values show very little or no change when the sample size increases from 250 to 500. Banerjee *et al.* (1992) suggest that the critical values for 500 observations can be treated as approximate asymptotic values. For model (17) critical values are obtained from two sources: Zivot and Andrews (1992) tabulate for min t_α and Christiano (1992) tabulates for $F_T^{\max}$. Christiano's critical values are for 154 observations and therefore we attach more importance to the inference based on min t_α in this case. Superscripts b and c indicate significance at 5% and 10%, respectively.

The initial start-up sample consists of 56 observations with $\delta_0 = 0{\cdot}15$. In order to assess the sensitivity of our results *vis-à-vis* trimming parameters we recomputed the tests setting $\delta_0 = 0{\cdot}10$ and $p = 4$. This allows breaks in the early and late samples, if they exist, to appear. However, the results remain qualitatively the same.[36]

Unit-root tests conditional on the identified mean shift are implemented as specified in equation (18) for the real exchange rates of Ringgit and Bhat. A lag length of 4 is selected. The computed ADF *t*-ratios are −3·137 and −3·033 for Ringgit and Bhat real exchange rates, respectively. The 5% and 10% asymptotic critical values are −5·08 and −4·81. Thus, the null of a unit root with unknown one-time structural break in the real exchange rate cannot be rejected in either case. Results are robust to $p = 8$ and the lag lengths determined by the Akaike and Schwarz information criteria.

[36] We acknowledge that the size and power of these tests may be sensitive to the trimming parameter, δ_0.

The main points of our results are as follows. First, shocks do not impinge on the behaviour of real exchange rates in the same way. This is evident from the fact that only one real exchange rate exhibits a significant mean shift which corresponds to the timing of the first oil shock and/or the end of the Bretton Woods era. The mean shift in Ringgit precedes these events. The other real exchange rates examined here do not appear to be significantly affected by these events. Second, the statistically identified date of break does not necessarily correspond to the visually identified one. On this point, our findings corroborate those of Raj (1992) who also reports that visual turning points do not necessarily correspond to the statistically significant ones. This highlights the problem associated with the exogenous determination of the break date (see Christiano, 1992). We find that the problem of structural breaks is not widespread at least in this data set. Finally, even after allowing for the mean shifts, ADF tests did not alter the non-rejection of a unit-root null.[37]

8. CONCLUSION AND IMPLICATIONS

We have investigated the PPP relationship by employing a unique data set obtained from the black markets for foreign exchange in developing countries. The black market exchange rates are perceived as a proxy of the developing countries' float rates. The data span is 31 years. In our view this study is a useful addition to the menu of empirical literature on PPP which is largely dominated by the currency pairs of industrial countries. The main findings of this paper are as follows.

First, the evidence is that the real exchange rates between rich and poor countries are stationary; it supports PPP but not the Balassa–Samuelson hypothesis — that the real exchange rates between rich and poor countries tend to be persistent. It also implies that the long-span studies may not suffer from survivorship bias, an issue emphasised by Froot and Rogoff (1995). Our results also offer some insight to Rogoff (1996) by showing that there is support for PPP across countries with differing growth experiences. The stationary of real exchange rates is confirmed by the LM-bar and T-bar tests as well as the variance ratio tests. The panel tests reject non-stationarity of real exchange rates in all cases; the variance ratio test does the same with only one exception (the non-stationarity of Kyat is not rejected in two sub-samples).

Most studies of the recent float cannot reject a unit root in real exchange rates between currency pairs of developed countries. Glen (1992) finds non-stationary real exchange rates for nine currencies of industrialized countries relative to the USdollar during the recent float. Grilli and Kaminsky (1991) report a non-stationary pound–USdollar real exchange rate both for the Bretton Woods and post-Bretton Woods periods. Both use the variance ratio test. Likewise, Froot and Rogoff (1995) could not reject the non-stationarity of a developing country's real exchange rate (the Argentine peso relative to the US dollar) over more than 70 years of data.[38] From this standpoint, our finding of stationary black market real exchange rates is interesting. Overall, we find strong support for PPP which may provide new insights. Although recent panel studies (e.g. Frankel and Rose, 1996; Oh, 1996; Wu, 1996) report stationary real exchange rates, consistent with our findings, the concern raised by IPS on the validity of these panel tests remain serious.

[37] This non-rejection of the null of a unit root with one-time endogenous structural break may be due to the low power of the ADF test. How to address structural break while testing for a unit root under the heterogeneous panel framework is not clear and requires further research.
[38] Froot and Rogoff's non-rejection could be due to the low power of the ADF tests that they apply.

Second, we find rapid convergence of black market real exchange rates to PPP. The estimates of the monthly rate of decay for the deviations from PPP are 6·01%, 3·36% and 2·63% for the pre-float, float and entire period, respectively.[39] They imply half-lives for deviations from parity of approximately 11 months for the pre-float, 21 months for the float, and 25 months for the full sample. These estimates of half-life are shorter than the 2·5 to 5·0 years reported elsewhere in the literature (e.g. Rogoff, 1996). It is also interesting to note that the rate of convergence during the fixed regime is almost twice that during the float. Such an outcome is plausible (see Wei and Parsley, 1995, p. 19).

Third, our results show that the stationarity of real exchange rates between rich and poor countries is not confined to high-inflation cases alone where relative prices are excessively volatile and monetary growth may overshadow the real factors. Our data set exhibits low relative price volatility and yet confirms the stationarity of real exchange rates. Thus, our findings complement those of McNown and Wallace (1989) and go some way to address the concern of Lothian and Taylor (1996) who caution against generalizing the results based on a data set with high relative price volatility. The low volatility of relative price ratios which we find in our data set is similar to those reported by Mussa (1986).

Fourth, we find that the black market real exchange rates are remarkably stable across the fixed and floating regimes. This contradicts the notion that real exchange rates tend to be more volatile under floating than under fixed nominal exchange rate regimes (e.g. Mussa, 1986). However, Grilli and Kaminsky (1991) argue that it is the real shocks to the economy which determine the behaviour (volatility) of real exchange rates rather than the nominal exchange rate regime *per se*. They attribute the excess volatility of the pound–USdollar real exchange rate during the recent float to the large real shocks that occurred during this period rather than to the floating regime. In similar vein, one can argue that the developing countries analysed here might not have experienced fundamentally different real shocks between pre- and post-1973 periods and hence no major difference in their real exchange rate volatility. In support of this argument one may also point out that the economic liberalization and transformation witnessed by these economies were part and parcel of a gradual process of development which started at different points in time during the mid-1960s to early 1970s. We acknowledge that, in recent times, East Asian economies are going through serious economic hardships (negative shocks); however, this does not affect our results as our sample extends only up to June 1989.

Nevertheless, Wald–Wolfowitz tests revealed that the pre-float and float periods' real exchange rates did not descend from an identical population in distribution; hence from a broader perspective these two samples of real exchange rates bear different characteristics. A similar findings is recorded for pound–USdollar real exchange rate by Grilli and Kaminsky (1991).

Fifth, formal tests for structural breaks reveal little evidence of segmented trends in black market real exchange rates. Significant mean shifts are identified in only two rates. This indicates that (1) the visually identified segmented trend does not necessarily conform to a statistically significant turning point and (2) identifying the break date exogenously (looking at the data plots) could be problematic. However, in view of our sample of 31 years, relative to hundreds of years covered by some typical long-horizon studies, this finding of a lack of a structural break in most real exchange rates should be taken with some caution. Nevertheless, they are sufficiently

[39] The convergence rate for the panel is calculated by averaging β obtained from estimating the equation: $\Delta q_{it} = \mu_i + \beta_i q_{i,t-1} + \Sigma_{j=1}^{2} \rho_{ij} \Delta q_{i,t-j} + \varepsilon_{it}$, where q_i $(i = 1, \ldots, N)$ denotes the real exchange rate. The absolute (mean) value of β gives the rate of decay (for the panel) for deviations from PPP, and the half-life is $\ln(0{\cdot}5)/\ln(1 + \beta)$.

indicative of the fact that the demise of the Bretton Woods system and the first oil shock did not inflict any trend segmentation on the majority of black market real exchange rates analysed here.

Finally, a large number of theoretical models,[40] assume that PPP holds continuously, and that the behaviour (volatility) of the real exchange rate is neutral with respect to the nominal exchange rate regime. Unfortunately, the experience of the recent float reported elsewhere in the empirical literature does not concur with this and suggests: (1) non-rejection of a unit root in real exchange rates, and (2) excess volatility of real exchange rates during the float regime. However, we provide evidence which, to a large extent, does concur with these assumptions and by implication lends support to the predictions of these theoretical models. As Dornbusch and Krugman (1976) put it 'Under the skin of an international economist lies a deep-seated belief in some variant of the PPP theory of exchange rates'. Our analysis using a unique data set supports this conviction.

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[40] Such models include the monetary models of exchange rate determination (see Frenkel and Johnson, 1978), the open economy version of the quantity theory (Lucas, (1982), and the literature on target zones, to name but a few.

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