On cross-country differences in the persistence of real exchange rates

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Abstract: Previous findings of long-run purchasing power parity come mainly from data for industrial countries, raising the issue of whether the results suffer sample-selection bias and exaggerate the general relevance of parity reversion. This study uncovers substantial cross-country heterogeneity in the persistence of deviations from parity. The results show that it is more likely, rather than less likely, to find parity reversion for developing countries than industrial countries. Although some cross-country persistence variations may partly reflect country differences in structural characteristics such as inflation experience and government spending, a considerable portion of those variations appears unaccounted for.

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1. Introduction

Arguing that the post-Bretton Woods period may be far too short to reveal PPP reversion, many studies explore long historical data and find evidence of parity reversion in real exchange rates (e.g., Abuaf and Jorion, 1990; Cheung and Lai, 1993; Culver and Papell, 1995; Diebold et al., 1991; Glen, 1992; Lothian and Taylor, 1996). The long-horizon approach, however, is susceptible to specific sample-selection bias, referred to as survivorship bias (Froot and Rogoff, 1995). Because of data availability, long-horizon studies of PPP investigate primarily industrial countries. In contrast, empirical evidence on PPP for developing countries is notably limited. For countries undergoing significant income growth from a low level, relative prices of tradables and nontradables can change substantially, inducing nonstationarity in real exchange rate dynamics. As a result, parity reversion may likely fail to work for this type of countries. A question has been raised about whether existing results from long-horizon studies, given their focuses on major industrial countries, may overstate the general significance of empirical support for long-run PPP. Clearly, economists like to use PPP as a frame of reference not just for industrial countries, but for the rest of the world as well. To resolve the issue apparently requires a large-scale study of different types of countries.

This study evaluates the significance of survivorship bias in PPP analysis by conducting an extensive time-series analysis of the persistence in dollar-based real exchange rates for 94 countries. Several related questions of interest are: Does the behavior of real exchange rates indeed differ between developing countries and industrial countries? If it does, is it less or more likely to find stationarity in real exchange rates for developing countries than industrial countries? Do the results based on industrial countries exaggerate the actual extent of empirical support for parity reversion?

The survivorship bias issue highlights a more basic issue concerning possible differences in the behavior of real exchange rates among countries or groups of countries. If the cross-country differences are substantial, empirical modeling of real exchange rate dynamics should take such cross-sectional heterogeneity into account and consider countries individually. Three different forms of parity-reverting dynamics are allowed for in this study; they include persistent autoregressive (AR) dynamics, persistent fractional dynamics, and trend-break/stationary dynamics. In doing this, the analysis permits diverse rates of parity reversion across real exchange rates.

The allowance for different reverting dynamics under the alternative hypothesis also addresses a commonly known problem associated with generic unit-root tests, namely, their low power against relevant stationary alternatives (Stock (1994) provides an excellent survey of the related issues). The three types of alternatives entertained here have been considered, albeit separately, in some earlier PPP studies to account for the empirical difficulty in detecting parity reversion (see Cheung and Lai (1998) for testing of persistent AR dynamics; Cheung and Lai (1993) and Diebold et al. (1991) for fractional analysis; and Culver and Papell (1995) and Perron and Vogelsang (1992) for trend-break analysis). To be sure, there are theoretical reasons to suggest that allowing for persistent alternatives is desirable. For example, intertemporal smoothing of traded goods consumption (Rogoff, 1992) or cross-country wealth redistribution effects (Obstfeld and Rogoff, 1995) may generate highly persistent dynamics for real exchange rates. Trend-break/stationarity alternatives, on the other hand, can be relevant for countries experiencing substantial changes in differential productivity growth in tradables and nontradables -- the oft-called Balassa-Samuelson effect (Balassa, 1964; Samuelson, 1964). Empirical results show that while much broader evidence in favor of parity reversion than that uncovered by prior studies can be obtained, no one single form of reverting dynamics is sufficient to capture the behavior of real exchange rates in all cases.

To gain more insights into the behavior of real exchange rates, this study conducts extensive analysis documenting variations in the persistence (measured by half-lives of shocks to parity) of real exchange rates across countries. Previous PPP studies for industrial countries typically report estimates of a half-life of 3 to 5 years. This study finds a different and much wider range of half-life estimates for developing countries. The study identifies, indeed, the presence of systematic differences in the persistence of PPP deviations across countries of different geographic regions, different levels of economic development (using both the World Bank's and IMF's classifications), and different exchange rate arrangements. The study further explores whether the observed disparities in the persistence can be linked to cross-country differences in the inflation experience, productivity growth, trade openness and government spending. The authors are unaware of any such systematic investigation of real exchange rate persistence using similar techniques, considering so many different factors, and over such a wide set of countries. This comprehensive analysis enables us to uncover patterns of different real exchange rate behavior across countries.

The country-specific approach here, which highlights cross-country heterogeneity, should be contrasted with the cross-section approach adopted by recent panel data studies of PPP. To search for more support for parity reversion in real exchange rates, a growing body of literature has turned to panel data methods and away from country-by-country analysis (Abauf and Jorion, 1990; Frankel and Rose, 1996; Wei and Parsley, 1995; Wu, 1996). Unlike long-horizon studies, which extend the sample period, panel studies advocate pooling data across currencies to increase statistical power. Panel unit-root tests are often implemented under the specification of at least some cross-country homogeneity in time-series dynamics. Papell (1997) considers panel tests that allow for heterogenous lag structures across data series. This author observes that panel results can be sensitive to the panel size as well as the country grouping. O'Connell (1998) further points out possible bias in panel tests due to cross-sectional dependence induced by calculating different real rates relative to the dollar. Moreover, panel unit-root tests examine the null hypothesis of a unit root for all pooled currencies. Rejections of the null does not necessarily imply that the currencies being pooled all contain no unit root. The rejections may just reflect the parity-reverting behavior of a possibly small subgroup of currencies only, so they cast little light on the question of actually how broad the relevance of long-run PPP is. Taylor and Sarno (1998) illustrate, indeed, that joint nonstationarity of a group of real exchange rates may be rejected when only one of these series is mean-reverting. In addition, since most panel PPP studies are still based on cross-country data sets comprising largely industrial countries, the issue of survivorship bias remains to be resolved.

2. Preliminary empirical analysis

The PPP theory suggests the presence of a long-run equilibrium relationship between national price levels of two countries when expressed in common currency units. In allowing for short-run deviations, an empirical representation of the PPP relationship is

$$p_t = p_t^* + s_t + u_t \tag{1}$$

where p_t and p_t^* are, respectively, the logarithms of the domestic and foreign price indexes; s_t is the logarithm of the spot exchange rate (domestic price of foreign currency); and u_t is an error term capturing deviations from PPP. For PPP to hold in the long run, the real exchange rate, measured by $y_t \equiv s_t + p_t^* - p_t$, should be stationary and not governed by permanent shocks.

The real exchange rate can be decomposed into various components:

$$s_t + p_t^* - p_t = s_t + p_t^{T*} - p_t^T + (1 - \alpha^*)(p_t^{N*} - p_t^{T*}) - (1 - \alpha)(p_t^N - p_t^T) \qquad (2)$$

where p_t^{T*} and p_t^{N*} are, respectively, the logarithms of the foreign tradables and nontradables price indexes; p_t^T and p_t^N are, respectively, the logarithms of the domestic tradables and nontradables price indexes; α^* is the geometric average weight of foreign tradables in the overall foreign price index; and α is the geometric average weight of domestic tradables in the overall domestic price index. This decomposition illustrates that even when goods arbitrage holds so that $s_t + p_t^{T*} - p_t^T$ is stationary, it is possible to reject long-run PPP if the price ratio between tradables and nontradables is not stationary. In this regard, Froot and Rogoff (1995) observe that the real exchange rate process is prone to instability for some developing countries because rapid income growth often induces drastic changes in the relative price structure between tradables and nontradables. This specific observation will be examined later in our empirical analysis.

PPP reversion has been documented in many long-horizon studies, which use exclusively data sets containing the post-1973 data as a small proportion only. It is not clear whether the inclusion of data from other historical periods biases unit-root test results. Specifically, the long-horizon findings may simply reflect the presence of parity reversion in the pre-1973 period solely and not in the post-1973 period as well. An interesting effort to address the issue comes from Lothian and Taylor (1996). In examining the dollar/pound (1791-1990) and franc/pound (1803-1990) real rates, these authors observe no significant evidence of a structural change between the pre- and post-1973 periods. If their results can be shown to be generally applicable to other long-horizon series of real exchange rates, they provide a justification for using data information of other historical periods to infer the behavior of real exchange rates for the post-1973 period. Clearly, evidence from the post-1973 period itself should be more definitive still.

Real exchange rates during the post-1973 period for 94 countries vis-à-vis the United States are investigated. Fisher and Park (1991) and Papell (1997), for example, have reported that it is more difficult to detect stationarity in dollar-based real exchange rates than German mark-based rates. Hence, if stationarity can be uncovered from dollar-based rates here, the results will be particularly significant and can strengthen and widen the empirical support for long-run PPP in general. In this study, all data series for constructing real exchange rates, including monthly averages of nominal exchange rates and monthly national price levels measured by consumer price indices, are taken from the IMF's *International Financial Statistics* data CD-ROM. The majority of these series cover the sample period from April 1973 through December 1994, though a small number of them are limited by data availability and have somewhat shorter sample periods.

Descriptions of the data samples and the countries under study are contained in Appendix. All the countries are identified for operational and analytical purposes as either industrial countries or developing countries, using the IMF classification given in *International Financial Statistics*. The category of developing countries under the IMF classification appears rather broad and is far from a homogeneous group. To exploit more data information, another country grouping based on income levels is considered. Every country is categorized according to the World Bank's (1989) definitions as low-income, medium-income, or high-income. Low-income countries are those with per capita GNP of $480 or less in 1987; middle-income countries are those with per capita GNP of more than $480 but less than $6,000 in 1987; and high-income countries are those with per capita GNP in excess of $6,000 in 1987. In empirical work the low-income and middle-income countries are sometimes referred conveniently to as developing countries, though the country composition can differ slightly from that under the IMF classification. The high-income countries as a group, on the other hand, overlaps the IMF's group of industrial countries substantially. Since changes in per capita GNP naturally occur over time, the income classification criteria have been adjusted periodically by the World Bank (WB). Specifically, the levels of per capita GNP dividing the different categories have been adjusting upward over the years. With the adjustments in criteria, the country composition of each income group under the WB classification has stayed stable over time.

To serve as a benchmark for comparison, all series of (log) real exchange rates are first tested for a unit root using a commonly used unit-root test, the augmented Dickey-Fuller (ADF) test. The ADF test involves estimating the following regression:

$$(1 - L)y_t = \mu_0 + \mu_1 t + \beta_0 y_{t-1} + \Sigma_{j=1}^{p} \beta_j (1 - L) y_{t-j} + \epsilon_t \tag{3}$$

where L is the lag operator and ϵ_t is the error term. The null hypothesis of a unit root is represented by $\beta_0 = 0$. The ADF statistic is given by the usual t-statistic for the β_0 coefficient.

Table 1 contains the results of the ADF test. Both ADF tests with and without a time trend are conducted. For cases of an insignificant time trend, results of the ADF test without a time trend are reported. For cases in which the time trend is significant at the 10% level or better, however, results of the test with a time trend are reported. Froot and Rogoff (1995) and Obstfeld (1993) have shown that the productivity growth differential between tradables and nontradables can lead to a time trend in the real exchange rate. The lag order used for the ADF test is determined using a data-dependent procedure based on the Akaike information criterion (AIC). To conserve space, results for the rejection cases only are reported in view of the rather large number of cases examined.

According to the ADF results, the evidence on long-run PPP is far from favorable. Out of the 94 cases under consideration, in only 17 cases can the unit-root null be rejected at the 10% level or better. Hence, as reported in other PPP studies, the ADF test uncovers not much evidence of parity reversion. The apparent failure to reject a unit root in most cases does not automatically lead to an outright rejection of the PPP hypothesis, nonetheless. It may simply confirm the low-power problem known to be associated with conventional unit-root tests. To handle the problem, several different statistical techniques are applied simultaneously in this study. This strategy explicitly allows for different types of subtle parity-reverting dynamics in real exchange rates. It is found that parity reversion can manifest itself in various forms and occur at widely different speeds across countries.

3. Testing of different mean-reverting dynamics

Unlike the univariate test approach here, some recent PPP studies have used multivariate unit-root tests with pooling of cross-country data. Taylor and Sarno (1998) point out a potential problem with the multivariate test approach. The null hypothesis in multivariate unit-root tests is typically that all the series of real exchange rates in the panel are not stationary. Since the null hypothesis will be violated even if merely one of the series is stationary, rejection of the null hypothesis renders little help to researchers in determining how many of the series under consideration are stationary. These authors recommend an alternative test procedure under which the null hypothesis is violated only when all the series in the panel are indeed stationary. The procedure turns out to be a special application of the widely used Johansen's cointegration test and has a known limiting $\chi^2(1)$ distribution (for the actual test specification, see Taylor and Sarno, 1998). This special Johansen-type test is performed on our real exchange rate data. Because of the degrees-of-freedom limitation, the test is applied to different subgroups of countries: the G-7 countries, European countries (excluding some of the G-7 members), African countries, Asian countries, and South and North American countries. In no cases can the null hypothesis that at least one of the panel series is nonstationary be rejected at a 10% significance level or better. We next go back to univariate test procedures, which allow for different parity-reverting dyanmics, to examine further how broad the empirical relevance of long-run PPP is.

In search of better support for parity reversion, the analysis proceeds by focusing on those series for which the ADF test fails to reject a unit root. The ADF test is known to have low power against persistent but mean-reverting dynamics under either local AR alternatives or fractional alternatives. In addition, the ADF test is constructed under the maintained hypothesis of a single, linear deterministic trend. When a trend break occurs, the ADF test suffers misspecification and has low power against stationary/broken-trend alternatives. To allow for these various mean-reverting alternatives, three different unit-root tests are performed on each series, and they are outlined below.

3.1. Against persistent local AR alternatives

Elliott, Rothenberg, and Stock (1996) analyze the sequence of Neyman-Pearson tests of the null hypothesis H_0: $\rho = 1$ against the local alternative H_a: $\rho = 1 + \bar{c}/T$, where ρ is the largest AR root and $\bar{c} < 0$. Based on asymptotic power calculation, a modified Dickey-Fuller test, called the DF-GLS test, achieves good power gains over traditional unit-root tests (see also Stock, 1994).

The DF-GLS$^{\tau}$ test that allows for a linear time trend entails the following regression:

$$(1 - L)y_t^{\tau} = \phi_0 y_{t-1}^{\tau} + \Sigma_{j=1}^{p} \phi_j (1 - L) y_{t-j}^{\tau} + w_t \tag{4}$$

where w_t is the error term; and y_t^{τ}, the locally detrended series under the local alternative, is

$$y_t^{\tau} = y_t - z_t \gamma \tag{5}$$

with γ being the least squares regression coefficient of $\tilde{y}_t$ on $\tilde{z}_t$, for which $\tilde{y}_t = (y_1, (1 - \bar{\rho}L)y_2, \ldots, (1 - \bar{\rho}L)y_T)'$ and $\tilde{z}_t = (z_1, (1 - \bar{\rho}L)z_2, \ldots, (1 - \bar{\rho}L)z_T)'$ for $z_t = (1, t)$. The DF-GLS$^{\tau}$ statistic is given by the *t*-ratio, testing H_0: $\phi_0 = 0$ against H_a: $\phi_0 < 0$. The same procedure applies to the case of no time trend (DF-GLS$^{\mu}$), except that y_t^{τ} is replaced with the locally demeaned series y_t^{μ} and $z_t = 1$.

3.2. Against persistent fractional alternatives

Fractionally integrated processes can display slow mean reversion, not captured by usual stationary processes. A fractionally integrated process is in general represented by

$$C(L)(1 - L)^d y_t = B(L)v_t \tag{6}$$

where $C(L) = 1 - c_1L - \ldots - c_pL^p$, $B(L) = 1 + b_1L + \ldots + b_qL^q$, all roots of $C(L)$ and $D(L)$ are stable, and v_t is the random error term. By considering noninteger values of the integration order d, fractional integration analysis avoids the knife-edged unit root/no unit root distinction and accommodates a broader range of mean-reverting dynamics than standard unit-root analyses. The low-frequency behavior of y_t is parameterized by d. Specifically, mean reversion occurs so long as $d < 1$. This condition offers the basis for fractional tests of parity reversion in y_t.

The d parameter can be estimated using a frequency-domain maximum likelihood procedure. Following Fox and Taqqu (1986) and Cheung and Diebold (1994), we exploit the property that maximization of the likelihood function is asymptotically equivalent to minimization of

$$\Sigma_{j=1}^{T-1} I_y(2\pi j/T)/f_y(2\pi j/T;\, \xi) \tag{7}$$

with respect to $\xi = (d, c_1, \ldots, c_p, b_1, \ldots, b_q)$, where $I_y(\lambda)$ is the periodogram of y at frequency λ, and $f_y(\lambda, \xi) = |1 - e^{-i\lambda}|^{-2d}|C^{-1}(e^{-i\lambda})B(e^{-i\lambda})|^2$ is proportional to the spectral density of y at frequency λ. The resulting estimator for d is consistent and asymptotically normal.

3.3. Against stationary/trend-break alternatives

To account for possible structural shifts in the real exchange rate behavior, sequential unit-root tests devised by Banerjee, Lumsdaine, and Stock (1992), henceforth BLS, are performed. The BLS sequential tests extend the ADF test by accounting for a possible jump or shift in trend in the data process, without knowing a prior the break date. Consider the following representation:

$$(1 - L)y_t = \mu_0 + \mu_1 t + \mu_2 d_t(k) + \alpha_0 y_{t-1} + \Sigma_{j=1}^{p} \alpha_j (1 - L)y_{t-j} + \zeta_t \tag{8}$$

where $d_t(k)$ is a dummy variable and ζ_t is the error term. When a trend shift is allowed for at time k, $d_t(k) = (t - k)I(t > k)$, with $I(\cdot)$ being the indicator function. Alternatively, when a mean shift (or a break in the trend) is allowed for at time k, $d_t(k) = I(t > k)$. For the usual ADF test, $d_t(k) = 0$. A sequence of statistics, $\tau_{DF}(k)$, indexed by k can be generated by varying k over the sample. BLS discuss several versions of the mean-shift or trend-shift sequential test. One is the minimal sequential test, and its test statistic is defined by

$$\mathrm{DF}_{min} \equiv \min_{r \leq k \leq T\text{-}r} \tau_{DF}(k) \tag{9}$$

for the sample size, T, and a trimming parameter, r. Following BLS, r is set equal to [.15T]. Another is based on the statistic, $\mathrm{DF}_{max\text{-}F} = \tau_{DF}(k^*)$, computed at where $F(k^*) = F_{max}$, for which

$$F_{max} \equiv \max_{r \leq k \leq T\text{-}r} F(k) \tag{10}$$

is constructed using sequential F-tests for the hypothesis of $\mu_2 = 0$ in equation (8).

4. Summary of unit-root test results

For the 77 series of real exchange rates for which the ADF test cannot reject a unit root, each one of them is subjected to the three different types of tests discussed earlier. Again, results for nonrejection cases are not reported to save space.

Results of the DF-GLS test are presented in Table 2. The AR lag order considered by the DF-GLS test is selected using the AIC, like the ADF test. Confirming the possible gain in power from using the DF-GLS test, the results indicate that significant evidence in favor of stationary alternatives can be found in 13 new cases. Hence, the use of the DF-GLS test helps find additional support for parity reversion, though the unit-root null still cannot be rejected in the majority of cases.

Table 3 displays the results of fractional integration analysis. Maximum likelihood estimates of the integration order, d, are obtained using the Davidson-Fletcher-Powell algorithm and based on the model specifications selected by the AIC, with both p and q being permitted to be less than or equal to three in model (6). Similar parsimonious model restrictions on p and q are often made in the empirical literature on fractional time series to reduce computation burden, given that lag selection for fractional models can be computationally demanding. Indeed, low orders of p and q (less than three in most cases) are found to be generally adequate in capturing the dynamics of real exchange rates. In 28 cases, the null of an exact unit root ($d = 1$) can be rejected in favor of the alternatives of a fractional unit root ($d < 1$). Since seven of these rejection cases are also obtained from the DF-GLS test, the fractional integration test reveals supportive evidence of parity reversion in 21 more cases, in addition to those reported earlier. The changes in results illustrate the low power of standard unit-root tests against fractional alternatives.

Table 4 gives the results of unit-root tests that allow for structural instability under the alternatives. The results based on the DF_{min} statistic are qualitatively the same as those using the $\mathrm{DF}_{max\text{-}F}$ statistic.

Significant evidence rejecting the unit-root null can be found in 22 cases, when either a mean shift or a trend shift is included under the alternative hypothesis. After subtracting away some overlapping cases in view of the other test results, the BLS sequential tests produce a net of 15 additional rejection cases. Interestingly, these trend-break cases involve primarily developing countries. The special pattern of these findings will be discussed more later.

To provide an overall picture, Table 5 sums up all the rejection results reported so far. When all the test results are combined together, the number of real exchange rate series found to display parity reversion increases considerably from that detected using merely the ADF test. In total, evidence in favor of parity reversion can be uncovered in 66 series of real exchange rates (over 70% of all the series examined). The most frequent type of cases seem to involve fractional dynamics, but no single model can sufficiently capture the dynamic behavior of all different real exchange rates.

It should be noted that since various unit-root tests have been applied to the data, the overall statistical significance of the findings needs further examination. Specifically, a Monte Carlo experiment is carried out to evaluate the overall test size for applying sequentially the ADF, DF-GLS, fractional, and structural-break tests, as conducted in our earlier analysis. The Monte Carlo results are summarized as follows: When 5% tests are applied, the overall empirical size is given by 15.6% and 13.9% for tests with and without a time trend, respectively. When 10% tests are used, the overall empirical size is given by 29.3% and 25.5% for tests with and without a time trend, respectively. These Monte Carlo results do not invalidate our empirical findings, given that evidence in favor of parity reversion can be uncovered in over 70% of the cases examined. Nevertheless, the findings reported earlier should be interpreted with such qualification in mind.

In studying these unit-root rejection results more closely, Table 5 also reports the rejection patterns under different country groupings, as categorized by several criteria.

4.1. By geographic regions

One of the advantages of a wide-scale, cross-country study is that potentially interesting sample information can be revealed by classifying the statistical results systematically into groups according to specific criteria. One classification studied is by geographic location (or, specifically, by continent): Africa, Asia, North America, South America, Europe, and Oceania. This choice is motivated by the recent trend of economic blocs in Europe, America, and Asia. Countries in a given region are likely to share similar social and economic conditions.

Panel A of Table 5 summarizes the findings of parity reversion for different country groups by geographic location. The hypothesis of equality in the rejection proportion across country groups is tested using a χ^2 test (see, e.g., Conover, 1980), and the hypothesis is strongly rejected by the data. This implies that country groups with geographical differences have significantly different unit-root rejection proportions. An observable feature of the results is that European countries give the lowest unit-root rejection frequency, whereas North American countries yield the highest unit-root rejection frequency. Such difference in the rejection proportion, however, cannot be explained just by location proximity to the U.S. since African countries show a higher unit-root rejection frequency than South American countries. On the other hand, a common characteristic of the European countries is their relatively high income levels. This leads us to an alternative grouping.

4.2. By levels of economic development

The central issue under investigation concerns the possibly different behavior of real exchange rates between industrial countries and developing countries. Prior studies on PPP reversion consider primarily industrial countries. This leads to the question of potential sample-selection bias or, more specifically, of whether the findings for industrial countries exaggerate the actual extent to which PPP reversion is relevant across countries in general. Froot and Rogoff (1995) observe that real exchange rates can be unstable for some developing countries because rapid income growth is often associated with drastic changes in the relative price structure between tradables and nontradables; consequently, it may be most difficult to find PPP reversion for this type of countries. These authors discuss a two-sector model of the Balassa-Samuelson effect, illustrating how trend breaks in real exchange rates can be linked to changes in the relative price structure between tradables and nontradables. Specifically, changes in the middle term of equation (2) can be represented by

$$d(p_t^{N*} - p_t^{T*}) \;=\; \mu_{LN}[d(\ln(A_T)) - d(\ln(A_N))]/\mu_{LT} \tag{11}$$

where $d(\cdot)$ is the differential operator; μ_{LT} and μ_{LN} are the respective labor's shares in tradables and nontradables production; and A_T and A_N are the corresponding productivity indexes for tradables and nontradables production. This suggests that the finding of a trend break in the real exchange rate can be interpreted as evidence of a structural change in productivity growth in tradables relative to nontradables.

The cross-country results from this study enable us to evaluate whether it is less or more likely to find parity reversion in real exchange rates for developing as compared to industrial countries. In addition to the IMF classification (industrial/developing), the WB classification (high-income/medium-income/low-income) is used for grouping countries. The unit-root rejection results for these various country groups are summarized in Panel B of Table 5.

Interestingly, the cross-country results indicate that developing countries tend to have a higher, rather than lower, unit-root rejection rate than industrial countries. The difference in the rejection proportion is statistically significant at the 5% level. Accordingly, using exclusively data for industrial countries in PPP studies does bias unit-root test results. Nevertheless, instead of overstating the extent of empirical relevance of parity reversion, results for industrial countries actually understate its general relevance. Consistent with Froot and Rogoff's (1995) observation, moreover, the breakdown of the rejection results among tests supports that the hypothesis of a structural shift in the real exchange rate (as indicated by the BLS test results) is more pertinent to developing countries than industrial countries.

The WB classification based on income levels reveals a similar unit-root rejection pattern as the IMF classification. The likelihood of rejecting a unit root in real exchange rates tends to be higher for lower-income countries. Again, the hypothesis of equality in the rejection proportion across country groups is strongly rejected by the data. In addition, the trend-break hypothesis appears more relevant to lower-income countries than higher-income countries.

Later in this study, we will examine whether the observed pattern of results under either IMF or WB classification reflects any systematic differences in the persistence of PPP deviations across relevant country groups. Before going into such analysis, however, the pattern of unit-root rejections under another classification is analyzed.

4.3. By exchange rate arrangements

Many PPP studies have suggested that it is less difficult to find parity reversion in fixed-rate data than flexible-rate data. Among the countries considered in this study, a small proportion of them have had their currencies pegged to the U.S. dollar. It is instructive to check based on our cross-country results if it is true that it is relatively easy to uncover parity reversion in fixed-rate data. Information about nominal exchange rate arrangements for individual countries is available from various issues of the IMF's *International Financial Statistics Yearbook* and *Annual Reports on Exchange Arrangements and Exchange Restrictions*. Given that some countries could adopt different exchange arrangements over the years, a currency is considered pegged to the U.S. dollar if the currency was pegged to it for at least two-thirds of the time over our sample period. Accordingly, there are 18 countries matching this definition; they include The Bahamas, Barbados, Dominica, El Salvador, Ethiopia, Grenada, Guatemala, Haiti, Honduras, Liberia, Netherlands Antilles, Panama, Paraguay, St. Kitts and Nevis, St. Lucia, Suriname, Trinidad and Tobago, and Venezuela.

Panel C of Table 5 gives the unit-root rejection results for country groups with currencies pegged and not pegged to the U.S. dollar. The unit-root rejection rate is about 89% for currencies pegged to the U.S. dollar as opposed to 66% for other currencies. In addition, the difference in the rejection rate is found to be statistically significant at the 10% level. These results are consistent with the usual finding that parity reversion is relatively easy to detect in fixed-rate data.

5. Half-lives of shocks to parity

The results reported in Section 4.2 support the presence of sample-selection bias when the data set consists principally of industrial countries. However, the direction of the bias is different from that predicted in the PPP literature: statistical results are biased toward understating, not overstating, the general relevance of parity reversion. These findings can be further understood by studying the persistence of PPP deviations for individual real exchange rates.

Persistence estimates are obtained from cumulative impulse response analysis, as discussed by Campbell and Mankiw (1987). In studying the moving-average (MA) representation of the $(1 - L)y_t$ process, the MA coefficients, $\{a_1, a_2, \ldots\}$, are referred to as impulse responses. The impact of a unit innovation at time t on the relevant variable at time $t + j$ can be shown to be given by $C(j) = 1 + a_1 + a_2$

$+ \ldots + a_j$, with $C(j)$ being called the cumulative impulse responses. Let $C(j)$ be the computed cumulative impulse response for a series of the real exchange rate. The half-life, denoted by ℓ_h, is given by $C(\ell_h) = ½$; it indicates how long it takes for the impact of a unit shock to dissipate by half.

Half-lives of shocks to parity are measured for all series of real exchange rates based on model specifications estimated under unit-root tests. Given that different parity-reverting models have been fitted to each series, it seems reasonable to choose a model specification under which the unit-root null can be rejected. For example, if the unit-root null can be rejected using simply the ADF test, the ADF regression model will be employed to compute the half-life for the relevant series. If the unit-root null can be rejected by the DF-GLS test but not the ADF test, the DF-GLS regression model will be used to compute the half-life for the series. If the unit-root null is rejected by the fractional integration test only, the half-life for the relevant series will then be calculated based on a fractional model. Likewise, if the unit-root null is rejected by the trend-break test solely, the half-life will be computed using the corresponding trend-break model. For the remaining series for which the various tests all fail to reject a unit root, the ADF regression model is used, arbitrarily, to compute the half-life.

The results of half-life estimation are summarized and displayed in Figure 1. For the group of industrial countries alone, the half-life estimates, which range mainly from 2 to 5 years, are similar to those reported in prior studies, which typically report a half-life of 3 to 5 years for PPP deviations. In contrast, the half-life estimates for developing countries appear much more dispersed than those for industrial countries. Most of the half-lives for developing countries are, nonetheless, less than 3 years. Accordingly, the persistence in PPP deviations tends to be lower for developing countries than for industrial countries. Such difference in the persistence of PPP deviations is consistent with the different unit-root rejection rates observed between the two country groups.

Under the WB classification, country groups of different income levels also show some systematic pattern of differences in persistence. In most cases, lower-income countries are more likely to have lower persistence in PPP deviations than higher-income countries. This pattern may account for the higher unit-root rejection rate observed for the lower-income group.

6. Empirical determinants of persistence

The foregoing results pose a new question: Why have PPP deviations for industrial or high-income countries as a group been more persistent, instead of less persistent, than those for other country groups? A possible line of research is to explore whether the observed pattern of persistence can be linked to some systematic differences in structural characteristics across countries.

Table 6 provides the median statistics for the half-life estimates and several measures of structural characteristics across country groups. All the data used are obtained from the IMF's *International Financial Statistics* data CD-ROM. Being restricted by data availability, a small number of countries are omitted when productivity growth, openness or government spending is considered. In terms of the persistence in PPP deviations, the group medians, coupled with their statistically significant differences, evidently confirm the pattern observed from Figure 1: industrial countries are more likely to display higher persistence in PPP deviations than developing countries; high-income countries tend to have higher persistence in PPP deviations than low-income countries, with medium-income countries standing in between.

In analyzing whether the observed pattern of persistence in PPP deviations is correlated with the countries' structural characteristics, the significance of the correlation is measured and formally tested based on our cross-country data. Since the usual distributional assumption on normality appears not tenable here, a nonparametric test is conducted. The test is also desirable for its robustness to extreme observations. For a pair of variables, say (x_1, x_2), the test statistic is

$$r_s = 1 - 6\Sigma_{j=1}^{N}[R(x_{1j}) - R(x_{2j})]^2/[N\,(N^2 - 1)] \tag{12}$$

where r_s is known as the Spearman rank correlation coefficient, N is the number of countries under consideration, x_{aj} $(a = 1, 2)$ is the rank of the jth observation of series x_a, and $R(x_{aj}) = 1$ if x_{aj} is the smallest observed value of x_a. In general, $-1 \leq r_s \leq 1$. When the variables are independent, $r_s = 0$. Asymptotically, $r_s N^{1/2}$ is distributed as the standard normal.

6.1. Inflation

Inflation experience is one of the countries' characteristics to be considered. The long-run PPP relationship can be viewed as an equilibrium condition of money neutrality in an international framework. If price movements are dominated by monetary shocks, there is a strong reason to expect parity reversion to prevail. Indeed, PPP has been known to hold well for high-inflation countries (Frenkel, 1978; McNown and Wallace, 1989). It is therefore instructive to see if there is any significant correlation between the cross-country differences in the persistence of PPP deviations and the differentials in inflation rates. To analyze this, the average inflation rate over our sample period is computed for each country.

In comparison to the group medians for the persistence in PPP deviations, Table 6 indicates that the group medians for the average inflation rate display an inverse pattern. During the sample period, industrial countries generally have experienced lower inflation than developing countries; high-income countries tend to have lower inflation than low-income countries, though low-income and medium-income countries differ not much in this respect. The rank correlation coefficient between the half-life of PPP deviations and the inflation rate is computed to be -.203 (see Table 7) with an approximate p-value of .049. The coefficient is statistically significant and negative, i.e., economies of higher inflation rates tends to be associated with lower persistence in PPP deviations.

6.2. Productivity growth

Productivity is another country characteristic to be studied. This supply-side factor is specifically focused by the Balassa-Samuelson hypothesis, which highlights the potential effects of productivity levels on real exchange rates. The study by Balassa (1964), for instance, explores a possible relationship between the level of the real exchange rate and the level of real per capita income. The present analysis concerns the persistence of PPP deviations, not the level of the real exchange rate. A question then is how much the observed pattern of persistence in PPP deviations can be explained by the differentials in productivity growth across countries. Similar to Balassa (1964) in measuring general productivity levels, the average rates of growth in per capita real GDP are calculated as a proxy for productivity growth for individual countries.

The group medians for the average rate of productivity growth across different country groups are reported in Table 6. Interestingly, industrial countries do not tend to have lower productivity growth than developing countries. Such a finding can be consistent with endogenous growth theory, which has suggested that productivity changes need not settle down at a steady state of slow growth in industrialized economies. We observe also that developing countries are far from homogeneous as a group, as indirectly shown under the WB classification. Although medium-income countries generally show higher productivity growth than high-income countries, low-income countries actually show the opposite. Nonetheless, the differences in the group medians turn out to be statistically insignificant. Furthermore, the rank correlation coefficient between the half-life of PPP deviations and the productivity growth rate is given by .111 with an approximate p-value of .337, suggesting that the correlation coefficient, albeit positive, is not statistically significant at any usual level of significance. Hence, the cross-country differentials in productivity growth can explain little the observed pattern of the persistence of PPP deviations.

It should be noted that the empirical failure to find a significant relationship in the foregoing analysis does not represent an outright rejection of the Balassa-Samuelson effect. Specifically, using the rate of change in per capita real GDP as a proxy for productivity growth may be too crude and highly inadequate. Indeed, the empirical evidence from prior work has been mixed (see Rogoff's (1996) review of the evidence on the Balassa-Samuelson effect), though some supportive evidence can be found when disaggregated data on sectoral productivity are used (e.g., DeGregorio, Giovannini and Wolf, 1994). The findings here are thus restricted by the lack of disaggregated productivity data for most of the countries under study.

6.3. Trade openness

A basic idea underlying the PPP theory is that goods market arbitrage can induce parity in prices for a sufficiently broad range of goods. Accordingly, deviations from PPP are corrected over time through adjustments in the trade flow. It appears instructive to investigate if the openness of the economy can influence the speed of parity reversion and therefore the persistence of PPP deviations. In this study, the openness of an economy is measured as the ratio of the average of imports and exports to the size of GDP of the economy.

The group medians for the openness measure under different country groupings are provided in Table 6 (the second to last column). The differences in group medians between industrial and developing countries under the IMF classification are not statistically significant. Under the WB classification, by contrast, the group medians can be noticeably different -- medium-income countries tend to have much more open economies, and low-income countries tend to have much less open economies. On the other hand, the rank correlation coefficient between the half-life of PPP deviations and the openness measure is estimated to be .005 with an approximate p-value of .966. Evidently, the correlation coefficient differs insignificantly from zero; indeed, the coefficient estimate practically equals zero. It follows that the differences in openness can explain little the observed pattern of the persistence in PPP deviations across countries.

6.4. Government spending

The last variable examined is government spending, a demand-side factor considered in some structural models of PPP deviations (Frenkel and Razin, 1987; Froot and Rogoff, 1991; Rogoff, 1992). Analytically, government spending can affect the relative demand for and thus the relate price of tradables and nontradables, especially in the short run. Froot and Rogoff (1991) point out that in comparison with private spending, government spending tends to fall more heavily on nontradables. Based on data for EMS countries, these authors find a significant effect of government spending on the real exchange rate. DeGregorio, Giovannini and Wolf (1994) also detect a significant positive relationship between government spending and (p_t^{N*} - p_t^{T*}) for a panel of OECD countries. Rogoff (1992) notes, however, that the allocation effects of government spending may not be long-lived if production factors are perfectly mobile across sectors over the long run. On the other hand, Alesina and Perotti (1995) show that government spending, if financed by distortionary taxes, can have long-run real effects.

A question here is whether the persistence pattern for PPP deviations is correlated with the cross-country differences in government spending. The group medians for government spending as a share of GDP are given in Table 6 (the last column). Systematic differences in the government spending/GDP ratio across country groups can be observed under both IMF and WB classifications. Industrial countries tend to have a higher government spending/GDP ratio than developing countries. Similarly, the government spending/GDP ratio tends to be lower in moving from high-income countries to medium-income countries and then from medium-income countries to low-income countries. The rank correlation coefficient between the half-life and the government spending ratio is computed to be .238 with an approximate p-value of .039, so the correlation coefficient is significantly greater than zero. This indicates that the persistence of PPP deviations has a significant and positive relationship with the government spending ratio across countries.

6.5. Section summary

The above analysis has investigated whether the diverse persistence observed between different real exchange rates can be attributable to some systematic structural differences across countries. In general, the empirical findings are mixed. Among the various structural characteristics examined, cross-country differences in both inflation experience and government spending are found to have a statistically significant relationship with the observed persistence pattern. A caveat is that the correlation in either case appears weak in magnitude. Accordingly, a substantial portion of the cross-country differences in the persistence still cannot be explained. Future research on the possible determinants of the persistence of PPP deviations is warranted.

7. Conclusion

The issue of survivorship bias in parity reversion analysis of real exchange rates has been explored. Prior studies on PPP examine primarily real exchange rates for industrial countries, which have high per capita income levels not typical for the rest of the world. This raises the question of whether previously reported findings of long-run PPP suffer sample-selection bias and exaggerate the general relevance of parity reversion. In the absence of direct evidence for non-industrial countries, the empirical evidence on PPP reported so far appears limited and fails to establish the actual extent of empirical support for parity reversion.

This study provides an extensive cross-country analysis of PPP reversion, using data on dollar-based real exchange rates for 94 countries. The analysis uncovers significant heterogeneity in the behavior of real exchange rates across countries or groups of countries. Various forms of parity-reverting dynamics are detected, and substantial variations in the persistence of PPP deviations are also observed. Interestingly, the results show that it is more likely, rather than less likely, to find parity reversion for developing countries than industrial countries. Hence, the use of data for industrial countries does bias

PPP test results, but it understates -- not overstates -- the general relevance of parity reversion. Moreover, although some of the cross-country variations in the persistence of PPP deviations may partly reflect country differences in structural characteristics such as inflation experience and government spending, a substantial portion of those cross-country variations appears still unaccounted for. A recent study by Taylor and Peel (1998) explores possible nonlinearity in the speed of PPP reversion. Further research on the nature of the dynamic process at which the real exchange rate adjusts to a shock may prove fruitful in providing new insights into the persistence issue.

A final remark is in order. The basic argument of Froot and Rogoff (1995) is that countries for which long spans of data are available have always been rich so that analysis based on the real exchange rates of these countries is likely to understate the Balassa-Samuelson effect. The findings in this study may not be inconsistent with the Froot-Rogoff argument (the authors owe this observation to an anonymous referee). As Edwards (1989) has observed, productivity growth can have both demand and supply effects. Productivity growth has a positive income effect, raising demand for nontradables. On the other hand, productivity growth can be factor augmenting, creating positive supply effects similar to an increase in factor availability under the Rybczynski principle. When the supply effects more than offset the demand effects, it generates a tendency toward real depreciation, possibly operating in the opposite direction to the Balassa-Samuelson effect .

Appendix
Data description and country classification

Country	Sample Size	Sample Period	Classification IMF	Classification WB	Classification Geog	Country	Sample Size	Sample Period	Classification IMF	Classification WB	Classification Geog
United States	261	4/73 - 12/94	I	H	NA	Thailand	261	4/73 - 12/94	D	M	AS
United Kingdom	261	4/73 - 12/94	I	H	EU	Gambia	261	4/73 - 12/94	D	L	AF
Austria	261	4/73 - 12/94	I	H	EU	Ghana	261	4/73 - 12/94	D	L	AF
Belgium	261	4/73 - 12/94	I	H	EU	Cote D'Ivoire	261	4/73 - 12/94	D	M	AF
Denmark	261	4/73 - 12/94	I	H	EU	Kenya	261	4/73 - 12/94	D	L	AF
France	261	4/73 - 12/94	I	H	EU	Madagascar	261	4/73 - 12/94	D	L	AF
Germany	261	4/73 - 12/94	I	H	EU	Mauritius	261	4/73 - 12/94	D	M	AF
Italy	261	4/73 - 12/94	I	H	EU	Morocco	261	4/73 - 12/94	D	M	AF
Luxembourg	261	4/73 - 12/94	I	H	EU	Niger	261	4/73 - 12/94	D	L	AF
Netherlands	261	4/73 - 12/94	I	H	EU	Nigeria	261	4/73 - 12/94	D	L	AF
Norway	261	4/73 - 12/94	I	H	EU	Seychelles	261	4/73 - 12/94	D	M	AF
Sweden	261	4/73 - 12/94	I	H	EU	Senegal	261	4/73 - 12/94	D	M	AF
Switzerland	261	4/73 - 12/94	I	H	EU	Swaziland	261	4/73 - 12/94	D	M	AF
Canada	261	4/73 - 12/94	I	H	NA	Burkina Faso	261	4/73 - 12/94	D	L	AF
Japan	261	4/73 - 12/94	I	H	AS	Fiji	261	4/73 - 12/94	D	M	OC
Finland	261	4/73 - 12/94	I	H	EU	Western Samoa	261	4/73 - 12/94	D	M	OC
Greece	261	4/73 - 12/94	I	M	EU	Chile	235	6/75 - 12/94	D	M	SA
Malta	261	4/73 - 12/94	D	M	EU	Costa Rica	243	10/74 - 12/94	D	M	NA
Portugal	261	4/73 - 12/94	I	M	EU	Panama	246	7/74 - 12/94	D	M	NA
Spain	261	4/73 - 12/94	I	H	EU	Uruguay	232	9/75 - 12/94	D	M	SA
Turkey	261	4/73 - 12/94	D	M	EU	Bahamas, The	261	4/73 - 12/94	D	H	NA
South Africa	261	4/73 - 12/94	D	M	AF	Grenada	228	1/76 - 12/94	D	M	NA
Colombia	261	4/73 - 12/94	D	M	SA	Guyana	237	4/75 - 12/92	D	L	SA
Ecuador	261	4/73 - 12/94	D	M	SA	St. Kitts & Nevis	192	1/79 - 12/94	D	M	NA
El Salvador	261	4/73 - 12/94	D	M	NA	St. Lucia	258	4/73 - 9/94	D	M	NA
Guatemala	261	4/73 - 12/94	D	M	NA	Suriname	253	4/73 - 4/94	D	M	SA

Country	Sample Size	Sample Period	Classification IMF	Classification WB	Classification Geog	Country	Sample Size	Sample Period	Classification IMF	Classification WB	Classification Geog
Haiti	261	4/73 - 12/94	D	L	NA	Bahrain	234	7/75 - 12/94	D	H	AS
Honduras	261	4/73 - 12/94	D	M	NA	Israel	180	1/80 - 12/94	D	H	AS
Mexico	261	4/73 - 12/94	D	M	NA	Jordan	228	1/76 - 12/94	D	M	AS
Paraguay	261	4/73 - 12/94	D	M	SA	Saudi Arabia	155	2/80 - 12/92	D	H	AS
Venezuela	261	4/73 - 12/94	D	M	SA	Bangladesh	246	7/74 - 12/94	D	L	AS
Barbados	261	4/73 - 12/94	D	M	NA	Nepal	256	4/73 - 7/94	D	L	AS
Dominica	261	4/73 - 12/94	D	M	NA	Botswana	235	6/75 - 12/94	D	M	AF
Jamaica	261	4/73 - 12/94	D	M	NA	Burundi	252	1/74 - 12/94	D	L	AF
Netherlands Antilles	261	4/73 - 12/94	D	M	NA	Congo	256	4/73 - 7/94	D	M	AF
Trinidad & Tobago	261	4/73 - 12/94	D	M	NA	Ethiopia	259	4/73 - 10/94	D	L	AF
Cyprus	261	4/73 - 12/94	D	M	EU	Liberia	207	4/73 - 6/90	D	L	AF
Egypt	261	4/73 - 12/94	D	M	AF	Zimbabwe	204	1/78 - 12/94	D	M	AF
Myanmar	261	4/73 - 12/94	D	L	AS	Rwanda	249	4/73 - 12/93	D	L	AF
Sri Lanka	261	4/73 - 12/94	D	L	AS	Gabon	190	4/73 - 1/89	D	M	AF
Hong Kong	261	4/73 - 12/94	D	H	AS	Somalia	100	4/73 - 12/89	D	L	AF
India	261	4/73 - 12/94	D	L	AS	Sudan	255	4/73 - 6/94	D	M	AF
Indonesia	261	4/73 - 12/94	D	L	OC	Togo	249	4/73 - 12/93	D	L	AF
Korea	261	4/73 - 12/94	D	M	AS	Uganda	168	1/81 - 12/94	D	L	AF
Malaysia	261	4/73 - 12/94	D	M	AS	Zambia	261	4/73 - 12/94	D	L	AF
Pakistan	261	4/73 - 12/94	D	L	AS	Solomon Islands	204	1/78 - 12/94	D	M	OC
Philippines	261	4/73 - 12/94	D	M	AS	Hungary	228	1/76 - 12/94	D	M	EU
Singapore	261	4/73 - 12/94	D	H	AS						

Notes: For the IMF's classification of countries (the column "IMF"), I = industrial countries and D = developing countries. For the World Bank's classification of countries (the column "WB"), H = high-income countries, M = medium-income countries, and L = low-income countries. For the geographical classification (the column "Geog"), AF = Africa, AS = Asia, NA = North America, SA = South America, EU = Europe, and OC = Oceania.

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Table 1
Unit-root rejections based on the ADF Test

Exchange rate series	Test type	Lag	ADF statistic		Exchange rate series	Test type	Lag	ADF statistic
				5				
El Salvador	With trend	4	-3.84**	5	Seychelles	With trend	12	-3.29*
Guatemala	With trend	1	-3.16*	5	St. Lucia	No trend	2	-2.72*
Barbados	With trend	5	-4.22**	5	Bahrain	With trend	2	-4.07**
Dominica	No trend	1	-2.80*	5	Nepal	With trend	12	-3.28*
Jamaica	With trend	5	-3.17*	5	Zimbabwe	With trend	2	-3.19*
Netherlands Antilles	With trend	2	-4.56**	5	Sudan	No trend	5	-3.52**
Indonesia	With trend	2	-3.34*	5	Uganda	With trend	2	-4.26**
Pakistan	With trend	2	-4.28**	5	Solomon Islands	With trend	4	-4.65**
Madagascar	With trend	2	-3.40*	5				
				5				

Notes: In cases in which the ADF test with a time trend is used, the time trend variable is found to be significant at the 10% level or better. The AR lag order used for the corresponding test is selected using the AIC (a maximum lag order of 12 is allowed for). Statistical significance is indicated by a single asterisk (*) for the 10% level and a double asterisk (**) for the 5% level.

Table 2
Additional unit-root rejections using the DF-GLS test

Exchange rate series	Test type	Lag	DF-GLS statistic	Exchange rate series	Test type	Lag	DF-GLS statistic
United Kingdom	No trend	4	-1.91*	Venezuela	With trend	1	-2.72*
France	No trend	4	-2.10**	Egypt	No trend	12	-2.07**
Germany	No trend	2	-1.83*	Gambia	With trend	1	-3.00**
Italy	No trend	2	-1.98*	Swaziland	No trend	9	-1.95*
Netherlands	No trend	2	-1.78*	Uruguay	No trend	3	-1.70*
Norway	No trend	2	-2.01*	Congo	No trend	1	-2.01**
Haiti	No trend	1	-2.26**				

Notes: In cases in which a test with a time trend is used, the time trend variable is significant at the 10% level or better. The AR lag order used for the corresponding test is chosen using the AIC (a maximum lag order of 12 has been allowed for). Critical values are based on Cheung and Lai (1995). Statistical significance is indicated by a single asterisk (*) for the 10% level and a double asterisk (**) for the 5% level.

Table 3
Additional unit-root rejections based on fractional analysis

Exchange rate series	Model code	d - 1	t-statistic for $d = 1$	Exchange rate series	Model code	d - 1	t-statistic for $d = 1$
United Kingdom	15	-1.0434	-3.3333**	Niger	1	-0.1034	-3.3200**
France	12	-0.7762	-9.6563**	Senegal	3	-0.9020	-6.2588**
Norway	12	-0.0867	-8.0710**	Burkina Faso	6	-1.0151	-7.8244**
Switzerland	16	-0.9465	-17.2543**	Fiji	16	-0.1240	-5.2168**
Japan	12	-0.6395	-23.0973**	Western Samoa	12	-0.9468	-19.9395**
Malta	2	-0.6554	-6.2039**	Panama	11	-0.3642	-4.2356**
South Africa	15	-0.2331	-2.2059**	Uruguay	9	-0.1697	-5.0015**
Honduras	12	-0.8031	-6.7932**	St. Kitts & Nevis	15	-0.1523	-2.2070**
Mexico	11	-0.1828	-8.3912**	Israel	12	-0.3786	-3.8882**
Venezuela	2	-0.3287	-2.6797**	Congo	1	-0.0492	-17.3486**
Trinidad & Tobago	5	-0.1537	-2.4308**	Ethiopia	6	-0.9695	-7.9186**
India	12	-0.6693	-13.0318**	Somalia	2	-0.2030	-1.4489*
Gambia	2	-0.9675	-10.7101**	Togo	6	-0.9270	-16.5096**
Kenya	12	-0.6702	-5.6342**	Zambia	12	-1.1161	-28.6253**

Notes: The fractional integration order, d, is estimated using Fox and Taqqu's (1986) frequency-domain maximum likelihood method. The t-statistics are computed for the hypothesis of $d = 1$, which is tested the one-sided alternative hypothesis of parity reversion, $d < 1$. Statistical significance is indicated by a single asterisk (*) for the 10% level and a double asterisk (**) for the 5% level.

Table 4
Additional unit-root rejections based on BLS sequential tests

Exchange rate series	Shift type	Lag	$DF_{max\text{-}F}$ statistic	DF_{min} statistic	Exchange rate series	Shift type	Lag	$DF_{max\text{-}F}$ statistic	DF_{min} statistic
Ecuador	Mean	1	-6.001**	-6.001**	Mauritius	Mean	2	-4.679*	-4.679*
	Trend	1	-2.462	-2.642		Trend	2	-3.159	-3.182
Haiti	Mean	1	-6.317**	-6.317**	Nigeria	Mean	1	-5.735**	-5.735**
	Trend	1	-2.769	-2.770		Trend	1	-1.595	-1.932
Honduras	Mean	7	-15.615**	-15.615**	Costa Rica	Mean	12	-12.498**	-12.498**
	Trend	7	-5.199**	-5.199**		Trend	12	-4.184*	-4.184*
Mexico	Mean	9	-5.596**	-5.596**	Panama	Mean	10	-5.415**	-5.415**
	Trend	9	-3.489	-3.489		Trend	10	-3.522	-3.522
Venezuela	Mean	3	-5.232**	-5.232**	Uruguay	Mean	3	-5.948**	-5.948**
	Trend	3	-3.898	-3.899		Trend	3	-2.317	-2.430
Myanmar	Mean	3	-1.623	-1.623	Bahamas, The	Mean	1	-3.495	-3.495
	Trend	3	-6.197**	-6.198**		Trend	1	-4.194*	-4.232*
Sri Lanka	Mean	8	-7.325**	-7.325**	Grenada	Mean	2	-1.918	-2.581
	Trend	8	-4.714**	-4.715**		Trend	2	-5.723**	-5.723**
India	Mean	4	-4.632*	-4.632*	Guyana	Mean	2	-5.824**	-5.824**
	Trend	4	-3.877	-3.877		Trend	2	-4.571**	-4.571**
Malaysia	Mean	3	-4.440	-4.440	Bangladesh	Mean	11	-4.663*	-4.663*
	Trend	3	-4.686**	-4.697**		Trend	11	-3.600	-3.600
Singapore	Mean	2	-3.322	-3.431	Liberia	Mean	10	-3.843	-3.843
	Trend	2	-5.791**	-5.791**		Trend	10	-4.204*	-4.465**
Ghana	Mean	2	-4.743*	-4.743*	Rwanda	Mean	2	-6.678**	-6.678**
	Trend	2	-3.003	-3.004		Trend	2	-3.496	-3.496

Notes: The AR lag order specified for the corresponding test is chosen based on the AIC. Critical values for the sequential mean-shift and trend-shift tests ($DF_{max\text{-}F}$ or DF_{min}) are provided by Banerjee, Lumsdaine, and Stock (1992; Table 2). Statistical significance is indicated by a single asterisk (*) for the 10% level and a double asterisk (**) for the 5% level.

Table 5
Summary of unit-root rejection results by various classifications

Country group	Number of countries	Total number of rejections	Breakdown of rejection cases among different tests		
			ADF	DF-GLS/Fractional	Trend-break (BLS)
All	94	66 (70.2%)	17 (18.1%)	34 (36.2%)	15 (16.0%)
Panel A. By geographic region					
Africa	28	23 (82.1%)	5 (17.9%)	13 (46.4%)	5 (17.9%)
Asia	17	11 (64.7%)	3 (17.6%)	3 (17.6%)	5 (29.4%)
North America	17	16 (94.1%)	7 (41.2%)	6 (35.3%)	3 (17.6%)
South America	8	4 (50.0%)	0 (0%)	2 (25.0%)	2 (25.0%)
Oceania	4	4 (100%)	2 (50.0%)	2 (50.0%)	0 (0%)
Europe	20	8 (40.0%)	0 (0%)	8 (100%)	0 (0%)
χ^2-test (5 df) =		18.79**			
Panel B. By economic development level					
IMF classification:					
Industrial	18	8 (44.4%)	0 (0%)	8 (44.4%)	0 (0%)
Developing	76	58 (76.3%)	17 (22.4%)	26 (34.2%)	15 (19.7%)
χ^2-test (1 df) =		7.07**			

Table 5
Continued

Country group	Number of countries	Total number of rejections	Breakdown of rejection cases among different tests		
			ADF	DF-GLS/Fractional	Trend-break (BLS)
WB classification:					
High-income	22	12 (54.5%)	1 (4.5%)	9 (40.9%)	2 (9.1%)
Medium-income	48	31 (64.6%)	11 (22.9%)	15 (31.3%)	5 (10.4%)
Low-income	24	23 (95.8%)	5 (20.8%)	10 (41.7%)	8 (33.3%)
χ^2-test (2 df) =		10.84^{**}			
Panel C. By exchange rate arrangement					
Pegged to US$	18	16 (88.9%)	6 (33.3%)	7 (38.9%)	3 (16.7%)
Not pegged to US$	76	50 (65.8%)	11 (14.5%)	27 (35.5%)	12 (15.8%)
χ^2-test (1 df) =		3.71^{*}			

Notes: The numbers in parentheses give the corresponding rejection frequency. The χ^2-test with the given degree of freedom (df) examines the null hypothesis that the country subgroups under an individual classification have the same rejection proportion, and statistical significance is indicated by a single asterisk (*) for the 10% level and a double asterisk (**) for the 5% level. The test for equality of proportions is a case of the chi-square goodness-of-fit test (see, e.g., Conover, 1980, Ch.4).

Table 6
Some descriptive characteristics by country groups

Country group	Persistence in PPP deviations	Inflation rate	Productivity growth	Economy's openness	Government spending
All	1.93 [94]	9.92 [94]	1.53 [75]	26.55 [77]	27.86 [76]
IMF Classification:					
Industrial	3.31 [18]	6.92 [18]	1.62 [17]	23.52 [18]	37.78 [18]
Developing	1.36 [76]	11.13 [76]	1.46 [58]	29.17 [59]	25.17 [58]
χ^2-test (1 df) =	22.26**	6.87**	1.28	1.89	7.28**
WB Classification:					
High-income	3.15 [22]	6.25 [22]	1.55 [19]	26.27 [20]	37.62 [19]
Medium-income	1.90 [48]	11.51 [48]	1.63 [37]	34.87 [36]	29.87 [38]
Low-income	0.93 [24]	11.13 [24]	0.69 [19]	15.11 [21]	17.52 [19]
χ^2-test (2 df) =	14.99**	15.22**	0.72	6.69**	13.58**

Notes: Persistence in PPP deviations is computed as the half-life (in years) of shocks to parity. The inflation rate is calculated as the average annualized rate (in %) of CPI-based inflation. Productivity growth is measured as the annualized rate (in %) of growth in per capita GDP. The economy's openness is proxied by the ratio (in %) of the average of exports and imports to the level of GDP. Government spending is given as the proportion (in %) of government expenditure in GDP. The figures in the table give the respective median values for individual country groups. The numbers in brackets indicate the number of observations used to compute the corresponding median values. The χ^2-test with the given degree of freedom (df) examines the null hypothesis of equality in median among the relevant subgroups, and statistical significance is indicated by a double asterisk (**) for the 5% level. The median test represents a special application of the chi-square goodness-of-fit test (see, e.g., Conover, 1980, Ch.4).

Table 7
Rank correlation analysis

	Inflation rate	Productivity growth	Economy's openness	Government spending
Persistence in PPP deviations	-.203 (.103) [.049]	.111 (.116) [.337]	.005 (.114) [.966]	.238 (.115) [.038]

Notes: The standard error for each individual estimate of Spearman's rank correlation coefficient is given in the parenthesis, and the approximate p-value corresponding the estimate is given in the bracket.