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Parity Reversion in Real Exchange Rates: Fast, Slow, or Not at All?

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Parity Reversion in Real Exchange Rates: Fast, Slow, or Not at All?

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Abstract

Consensus estimates put the half-life of deviations from purchasing power parity (PPP) at about four years (Rogoff, 1996). However, conventional least squares estimates of half-lives are biased downward. Accordingly, as a preferred measure of the persistence of real exchange rate shocks, this study uses median-unbiased estimators of the half-life of deviations from parity, which correct for the downward bias of conventional estimators. The paper tests for PPP using real effective exchange rate data for 90 developed and developing countries in the post-Bretton Woods period. Support for PPP is found, as the majority of countries experience finite deviations of real exchange rates from parity. The speed of parity reversion is found to be typically much faster for developed countries than for developing countries, and to be considerably faster for countries with flexible nominal exchange rate regimes in comparison with countries having fixed nominal exchange rate regimes.

JEL Classification Numbers: C22, F31, F41

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I. Introduction

Exchange rates have been at the center of policy and academic debates in developing and developed countries, especially since the floating of developed-country exchange rates in the early 1970s, which marked the commencement of the post-Bretton Woods period. Despite these debates, several key empirical questions regarding the stylized facts of exchange rates remain largely unresolved, particularly in the little-researched area of real exchange rate behavior in developing countries (Edwards and Savastano, 2000). Several of these exchange rate-related questions are tackled in this paper: Do real exchange rates really display parity-reverting behavior? Is purchasing power parity (PPP) an appropriate (very) long-run benchmark for assessing real exchange rate developments? Does the behavior of real exchange rates differ between developed and developing countries? Are there important differences in the extent of parity reversion when comparing countries with fixed nominal exchange rate regimes and those with flexible nominal exchange rate regimes? Can we explain cross-country heterogeneity of parity reversion using countries' structural characteristics, and if so, which characteristics appear to be important? In this study we examine the empirical support for PPP, through time-series analysis of the persistence of shocks to the real effective exchange rates of 90 developed and developing countries in the post-Bretton Woods period.

The theory of PPP, in its most rudimentary form, states that there is an equilibrium level to which exchange rates converge, such that foreign currencies should have the same purchasing power.[2] Therefore, long-run PPP is inconsistent with a unit root in real exchange rates. The reason for this is that a shock to a unit root process will have permanent effects on all future values of the series, potentially without bound. Unfortunately, formal statistical tests that compare the unit root (UR) model against the alternative of a stationary autoregressive (AR) model, typically lead to a failure to reject the hypothesis of a unit root.

There are a number of econometric problems with using the UR/AR model to test for PPP. First, least squares estimators of AR models bias empirical results in favor of finding PPP. This downward bias becomes particularly acute when the AR parameter is close to unity. As lower values of the AR parameter imply faster speeds of adjustment following a shock, this will also result in a downward bias to estimates of half-lives of shocks. This near unit root bias is likely to be particularly relevant for real exchange rates, as they are often found to be stationary, yet exhibit shocks which are highly persistent.

[2] The version of PPP with the longest pedigree is that of relative PPP, which states that the exchange rate will be proportional to the ratio of money price levels (including traded and nontraded goods) between countries, that is to the relative purchasing power of national currencies (see Wickham, 1993). As the present study of PPP employs data on price indices rather than price levels, we are examining the relevance of relative PPP—the notion that the percentage change in the nominal effective exchange rate should compensate for the inflation differential between the home country and a weighted average of partner-countries. The equilibrium real exchange is constant if relative PPP holds, so that deviations from PPP must be stationary for PPP to be valid. For surveys on PPP and exchange rate economics, see Froot and Rogoff (1995), Sarno and Taylor (2002), and Taylor (2003).

Second, unit root tests tend to be uniformative as to the speed of party reversion, because a rejection of the unit root null could still be consistent with a stationary model of real exchanges rates that has highly persistent shocks. This leads to problems about how to interpret results, and often yields arbitrary conclusions that are dependent on the predisposition of the researcher.

The contributions to the literature of this paper are fourfold. First, the median-unbiased estimator of Andrews and Chen (1994) is used to obtain estimates of the AR parameter in the real exchange rate data. Second, these estimates of the AR parameter and associated impulse response functions are used to calculate an unbiased scalar measure of the average duration (in terms of half-lives) of typical real exchange rate shocks. Third, using Andrews' (1993) unbiased model-selection rule we can be more definitive about our willingness to draw conclusions as to the presence (or absence) of parity reversion of real exchange rates in the post-Bretton Woods period. In particular, the use of median-unbiased estimation methods enables us to adopt Andrews' model-selection rule, whereby finite (permanent) half-lives constructed from bias-corrected estimation indicate that real exchange rate shocks are mean reverting (or permanent). Fourth, to the best of our knowledge, this study is the first to apply median-unbiased estimation methods to developing country real exchange rate data.[3]

Our main results may be summarized briefly. First, using post-Bretton Woods data on the real effective exchange rates of 90 industrial and developing countries and least squares estimation of unit root models, we replicate Rogoff's (1996) consensus estimate of the half-life of deviations from PPP of between three to five years. Second, using median-unbiased estimates (robust to serial correlation) we find that the half-lives of deviations of real exchange rates from PPP are typically *longer* than the previous consensus, with cross-country average (median) half-lives lasting about eight years. In particular, we find that for 40 countries in our sample, deviations of the real exchange rate from parity are best viewed as being permanent. However, the majority (50) of countries have finite half-life estimates—this evidence of real exchange rate revision to parity is consistent with PPP holding in the post-Bretton Woods period. Third, the median half-life of parity deviations for industrial countries (eight years) is much shorter than that for developing countries (permanent). Fourth, the median half-life of parity deviation for countries with fixed nominal exchange rate regimes (permanent) is considerably longer than that for countries with flexible nominal exchange rate regimes (six years). Fifth, two key structural characteristics that affect the persistence of shocks to real exchange rates are differences in a country's inflation experience and the extent of nominal exchange rate volatility. Sixth, the existing literature on measuring parity reversion in developing countries has largely concentrated on the Latin American experience (which is dominated by high inflation countries). This sample-selection bias has resulted in the (erroneous) finding that parity reversion in developing countries is faster than that observed for developed countries.

[3] Recent work on testing PPP and the stationarity of real exchange rates in developing countries includes Montiel (1997), and Bahmani-Oskooee (1995), among others; those concentrating on Latin American countries include Devereaux and Connolly (1996), and Calvo, Reinhart and Végh (1996).

The remainder of the paper is as follows. Section II sets out the median-unbiased procedure for estimation of AR/UR models, which are used to examine the persistence of real exchange rate shocks. The data used in the study is described in Section III. Section IV presents the main empirical findings of biased and median-unbiased point estimates of the persistence of real exchange rate shocks, while Section V examines the key determinants of cross-country heterogeneity in the duration of real exchange rate deviations from purchasing power parity. Section VI provides some concluding comments.

II. Biased and Unbiased Measures of Parity Reversion Speeds

The existence of long-run PPP is inconsistent with unit roots (infinite half-lives of parity deviation) in the real exchange rate process. This notion has stimulated the growth of a large literature, using various tests, to resolve whether PPP holds in the post-Bretton Woods period. However, standard unit root tests focus only on whether such shocks are mean-reverting (finite persistence) or are not (permanent). For economists, long-run PPP tells us more than the absence of a unit root—it also means whether there is a sufficient degree of mean reversion in exchange rates (over the horizon of interest) to validate the theoretical predictions of models based on the PPP assumption. For example, using the Dornbusch (1976) overshooting model, which has plausible assumptions about nominal wage and price rigidities, we would expect substantial convergence of real exchange rates to PPP over one-to-two years. Rather than use unit root tests to evaluate PPP, it is preferable to use a scalar measure of the speed of reversion of real exchange rate shocks. Recent papers examining the post-Bretton Woods period have used estimates of the half-life of deviations from PPP to calculate the extent of such reversions (Andrews (1993), Andrews and Chen (1994), and Cheung and Lai (2000b)).[4]

Downwardly-biased autoregressive parameters

Analysis of the downward bias in AR models has a long history. Orcutt (1948) showed that least squares estimates of lagged dependent variable coefficients (such as the autoregressive parameter (α) in the AR(1) or Dickey-Fuller regression) will be biased towards zero in small samples. Marriott and Pope (1954) established the mean-bias of the least squares estimator for the stationary AR(1) model, while Shaman and Stine (1988) extended this to the AR(p) model. While least squares will be the best linear unbiased estimator under the Gauss-Markov theorem, in the autoregressive case the assumptions of this theorem are violated, as lagged values of the dependent variable cannot be fixed in repeated sampling, nor can they be treated as distributed independently of the error term for all lags. Marriott and Pope (1954) showed that, ignoring second-order terms, the expected value of the least squares estimate of

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[4] Biased and median-unbiased point estimates of the half-life of shocks to economic time series have also been used: by Cashin, Liang and McDermott (2000) in modeling the persistence of shocks to world commodity prices; by McDermott (1996), Cashin and McDermott (2003) and Murray and Papell (2002) in modeling the persistence of parity deviations in developed-country real exchange rates; and by Cashin, McDermott and Pattillo (2004) in examining the persistence of terms of trade shocks.

the true α in the AR(1) model can be approximated by: $E(\hat{\alpha}) = \alpha - (1+3\alpha)/N$, where $N = T - 1$. Using simulation methods, Orcutt and Winokur (1969) find that, for $T = 40$ and true $\alpha = 1$, the least squares mean bias is $E(\hat{\alpha}) - \alpha = 0.129$. Similarly, the simulation calculations of Andrews (1993) reveal that the least squares median bias of the AR(1) model, again for $T = 40$ and true $\alpha = 1$, is slightly smaller at 0.107. In general, the larger is the true value of α, the larger is the least squares bias, and so the bias is largest in the unit root case. This bias shrinks as the sample size grows, as the estimate converges to its true population value.

The downward bias in least squares estimates of the autoregressive parameter arises because there is an asymmetry in the distribution of estimators of the autoregressive parameter in AR models. The distribution is skewed to the left, resulting in the median exceeding the mean. As a result, the median is a better measure of central tendency than the mean in least squares estimates of AR models.

We use the AR(p) model to measure the degree of persistence in real exchange rates. The AR(p) model (also known as an Augmented Dickey-Fuller regression) takes the form

$$q_t = \mu + \alpha\, q_{t-1} + \sum_{i=1}^{p-1} \psi_i\, \Delta q_{t-i} + \varepsilon_t \qquad \text{for } t = 1,\ldots,T\,, \tag{1}$$

where the observed real exchange rate series is q_t: $t = -p,\ldots,T$, μ the intercept, α the autoregressive parameter (where $\alpha \in (-1,1]$), and ε_t are the innovations of the model. The lagged first differences are included to control for the presence of serial correlation (which are typical in economic time series).[5] Andrews and Chen (1994) show how to perform approximately median-unbiased estimation of autoregressive parameters in Augmented Dickey-Fuller regressions. The bias correction delivers an impartiality property to the decision making process, because there is an equal chance of under- or overestimating the AR parameter. Moreover, an unbiased estimate of α will allow us to calculate an unbiased scalar estimate of persistence—the half-life of a unit shock.

Bias-correcting estimates of the autoregressive parameter and the model selection rule

Andrews (1993) and Andrews and Chen (1994) present a method for median-bias correcting the least squares estimator. To calculate the median-unbiased estimator of α, suppose $\hat{\alpha}$ is an estimator of the true α whose median function ($m(\alpha)$) is uniquely defined $\forall\alpha \in (-1,1]$. Then $\hat{\alpha}_u$ (the median unbiased estimator of α) is defined as:

[5] In examining for the presence of serial correlation, the general-to-specific lag selection procedure of Ng and Perron (1995) and Hall (1994) is used, with the maximum lag set to twelve. Starting with the maximum lag, first-differences of the logarithm of the REER (q_t) were sequentially removed from the AR model until the last lag was statistically significant (at the 5 percent level). At that point all lag lengths smaller than or equal to $p - 1$ are included in the AR(p) regression of equation (1).

$$\hat{\alpha}_u = \begin{cases} 1 \text{ if } \hat{\alpha} > m(1), \\ m^{-1}(\hat{\alpha}) \text{ if } m(-1) < \hat{\alpha} \leq m(1), \\ -1 \text{ if } \hat{\alpha} \leq m(-1), \end{cases} \tag{2}$$

where $m(-1) = lim_{\alpha \to -1}\, m(\alpha)$, and m^{-1}:$(m(-1), m(1)] \to (-1,1]$ is the inverse function of $m(.)$ that satisfies $m^{-1}(m(\alpha)) = \alpha$ for $\alpha \in (-1,1]$. That is, if we have a function that for each true value of α yields the median value of $\hat{\alpha}$, then we can simply use the inverse function to obtain a median-unbiased estimate of α. For example, if the least squares estimate of α equals 0.8 then we do not use that estimate, but instead use that value of α which results in the least squares estimator having a median of 0.8.

It is also possible to control for the presence of serial correlation by using the unit root procedures suggested by Phillips and Perron (1988)—for an example, see Cashin and McDermott (2003). This would have the benefit of allowing a wider class of serial correlation and heterogeneity of the errors, and would be an advantage when dealing with exchange rate data, since time series models of such data are often found to contain heterogeneous errors. Nonetheless, to calculate a measure of persistence, it is easier to deal with a full parametric model rather than the semi-parametric methods of Phillips and Perron (1988). Moreover, measures of persistence derived from parametric models can be more easily compared with previous studies. Cashin and McDermott (2003) find that half-lives of real exchange rate deviations from parity derived from both Dickey-Fuller regressions and Phillips-Perron regressions are broadly similar. Accordingly, it appears that the trade-off when choosing a simpler method involves a relatively small cost. However, neglecting to deal at all with serial correlation would lead to serious over-estimation of the persistence of shocks to the real exchange rate.

Model selection rule

Testing the PPP proposition is a controversial subject that has not yet been resolved. Most modern time-series tests of PPP are based on testing whether α <1 in AR models of the real exchange rate, such as in equation (1). Unit root tests based on classical hypothesis tests tend to have a null hypothesis where PPP does not hold—that is, the null hypothesis is that PPP deviations are permanent. Further, because of the low power properties of these tests (at least against reasonable alternatives), the tests often fail to reject the null. In particular, slow (albeit positive) reversion of real exchange rates toward PPP yield unit root tests that provide little information against relevant alternative hypotheses (Froot and Rogoff, 1995). The typically large estimated confidence intervals around the point estimate of α make it clear that the level of uncertainty about the value of the “true” α is very high. Nevertheless, practitioners often do use hypothesis tests as a formal model selection rule.[6]

[6] Bayesian estimation of AR models tries to get around this problem by using flat or noninformative priors, thus hoping obtain an impartial procedure, but there remain doubts as to how noninformative the prior really is (Phillips, 1991).

Andrews (1993) offers a solution to this impasse by introducing a statistical procedure whereby the probability of selecting the "true" model is at least as large as the probability of selecting the false model.[7] Unbiased model-selection procedures have an impartiality property that may be useful if the selection of one model or another (such as the trend stationary or unit root model) is a contentious issue. Suppose the problem is to select one of two models defined by $\alpha \in I_a$ and $\alpha \in I_b$, where I_a and I_b are intervals partitioning the parameter space (-1, 1] for α, with $I_a = (-1, 1)$ and $I_b = \{1\}$. Then the unbiased model selection rule would indicate that model I_m should be chosen if $\hat{\alpha}_u \in I_m$, for $m = a, b$. This is also a valid level 0.50 (unbiased) test of the H_0: $\alpha \in I_a$ versus H_1: $\alpha \in I_b$. Importantly, the median-unbiased estimator $\hat{\alpha}_u$ is the lower and upper bounds of the two one-sided 0.5 confidence intervals for the true α when $m(.)$ is strictly increasing (Andrews 1993, p.152). These confidence intervals have the property that their probabilities of encompassing the true α are one-half.

In a Monte Carlo study of the AR(p) model, Andrews and Chen (1994, p.194) demonstrate that the unbiased model-selection rule has a probability of correctly selecting the unit-root model (when the true α =1) of about 0.5 (that is, it has an equal chance of accepting or rejecting the unit root null). This is much lower than the corresponding probability for a (two-sided) level 0.10 test or (one-sided) level 0.05 test of a unit-root null hypothesis, as the unbiasedness condition does not give a bias in favor of the unit root model. The greater size of Andrews' unbiased model selection rule, in comparison with conventional tests, increases the probability of rejecting the unit root null. This indicates that if the true $\alpha < 1$, then the probability of a type II error is smaller for Andrews' model selection rule than for conventional tests, especially for the near unit root case. See Cashin and McDermott (2003) for additional details.

Calculating half-lives

Our interest in this paper concerns the persistence of shocks to economic time series. In this connection, the impulse response function of a time series $\{q_t: t=1,2,...\}$ measures the effect of a unit shock occurring at time t (that is, $\varepsilon_t \rightarrow \varepsilon_t+1$ in equation (1)) on the values of q_t at the future time periods t+1, t+2,... . This function quantifies the persistence of shocks to individual time series. For an AR(p) model the impulse response function is given by

$$IR(h) = f_{11}^{(h)} \qquad \text{for } h = 0, 1, 2, \dots, \tag{3}$$

where $f_{11}^{(h)}$ denotes the (1,1) element of $\mathbf{F}^h$ and where $\mathbf{F}$ is the ($p \times p$) matrix

[7] The unbiased model selection procedure based on the median-unbiased estimate of an AR(1) model is an exact test. However, the unbiased model selection procedure based on the median-unbiased estimate of the AR(p) model is an approximate test. This is because the distribution of $\hat{\alpha}_u$ calculated from the AR(p) model depends on the true values of the ψ_i terms in equation (1), which are unknown. Andrews and Chen (1994) demonstrate that the approximately median-unbiased point estimates of α in the AR(p) model are very close to being median-unbiased.

$$F \equiv \begin{bmatrix} \alpha_1 & \alpha_2 & \alpha_3 & \cdots & \alpha_{p-1} & \alpha_p \\ 1 & 0 & 0 & \cdots & 0 & 0 \\ 0 & 1 & 0 & \cdots & 0 & 0 \\ \vdots & \vdots & \vdots & \cdots & \vdots & \vdots \\ 0 & 0 & 0 & \cdots & 1 & 0 \end{bmatrix}. \tag{4}$$

However, rather than consider the whole impulse response function to gauge the degree of persistence, we use a scalar measure of persistence that summarizes the impulse response function: the half-life of a unit shock (l_h). When the parity reversion process is characterized by nonmonotonic adjustment, the approximately median-unbiased estimate of the half-life for AR(p) models (such as the Augmented Dickey-Fuller regression) can be calculated from the impulse response functions of equation (3), with the half-life defined as the duration of time it takes for a unit impulse to dissipate permanently by one half from the occurrence of the initial shock (Cheung and Lai, 2000b). That is, the half-life is given by $IR(l_h) \leq ½$ such that $l_{h+k} < ½$ for $k = 1,2,...$.

As with the estimation of α, the median-unbiased half-lives can be interpreted as follows. Using the Andrews unbiased model-selection rule, there is a 50 percent probability that the confidence interval from zero to the estimated median half-life contains the true half-life of a shock to any given time series, and a 50 percent probability that the confidence interval from the estimated median half-life to infinity contains the true half-life of a shock to any given time series.

We could construct two-sided 90 percent (or one-sided 95 percent) confidence intervals for the half-life estimates, as per Cashin, Liang and McDermott (2000), Cheung and Lai (2000b), and Murray and Papell (2002). Because such intervals are typically very wide, the null hypothesis that the half-life is permanent is rarely rejected. Such a finding leaves us in the awkward position of having to provisionally accept that PPP does not hold, or resigning ourselves to the fact that we cannot make a decision at all because the degree of uncertainty attached to the estimate of the half-life is too high. Murray and Papell (2002) favor the latter option.

However, empirical researchers will typically (for better or worse) make a judgment as to whether the model is acceptable or not. Failure to reject the null of a unit root often invokes the argument that the test was not powerful enough and researchers then proceed as if the null is false. This approach seems unsatisfactory and is one that will often deliver different conclusions from different researchers, because each is now using an arbitrary decision rule—an undisclosed and implicit decision rule at that. Instead, a more natural procedure is to use the unbiased-model selection rule of Andrews (1993), which offers an objective means to determine which model best represents the data. Moreover, because the rule is clear and explicit different researchers can use the rule to verify results in an objective manner. Use of Andrews' (1993) unbiased model-selection rule and median-unbiased estimation means that we have used a 0.5 (unbiased) test of the null hypothesis of a finite half-life versus the alternative hypothesis of a permanent half-life (see Cashin and McDermott, 2003).

There are other methods we could have used to examine the cross-country experiences of the persistence of parity deviations. One approach is to use long-span time series to analyze the persistence of real exchange rates. However, this approach is clouded by the differing nominal exchange rate regimes used by any given country over time. In order to provide a useful test of the validity of the notion of purchasing power parity in the post-Bretton Woods period, a test needs to be derived using data specifically from that period. Another approach would be to use multivariate generalizations of unit root tests of the real exchange rate (through panel unit root analyses of long-run PPP). This approach also has substantial problems. In particular, the use of panel unit root tests has been criticized because authors have typically presumed that rejection of the *joint* null hypothesis of unit root (nonreversion) behavior of the whole panel of real exchange rates implies that *all* real exchange rates are stationary. However, in actuality rejection of the null only implies that *at least one* of the real exchange rates is stationary (or mean reverting). A third approach would be to use nonlinear methods as per Taylor and Sarno (1998). In some sense this approach complements the approach used in this paper.

III. DATA

In this and the following section we will investigate the properties of real exchange rate persistence. The data used to estimate the near unit root model are monthly time series of the real exchange rate obtained from the International Monetary Fund's *International Financial Statistics* (*IFS*) over the sample 1973:3 to 2002:3 (the post-Bretton Woods period), which gives a total of 348 observations—see Appendix I for additional details. Appendix I also lists the derivation and description of the variables used in Section V to explain the cross-country heterogeneity of the duration of real exchange rate shocks.

The definition of the real exchange rate is the real effective exchange rate (REER) based on consumer prices (line *rec*), for which 20 industrial and 70 developing countries were selected. As such, we will examine the behavior of REER based on: (i) the nominal effective exchange rate, which is the trade-weighted average of bilateral exchange rates vis-à-vis trading partners' currencies; (ii) the domestic price level, which is the consumer price index; and (iii) the foreign price level, which is the trade-weighted average of trading partners' consumer price indices. We analyze effective rather than bilateral real exchange rates as the effective rate measures the international competitiveness of a country against all its trade partners, and helps to avoid potential biases associated with the choice of base country in bilateral real exchange rate analyses.

The REER indices measure how nominal effective exchange rates, adjusted for price differentials between the home country and its trading partners, have moved over a period of time. The CPI-based REER indicator is calculated as a weighted geometric average of the level of consumer prices in the home country relative to that of its trading partners, expressed in a common currency. The Fund's CPI-based REER indicator (base 1995=100) of country i is defined as $q_i = \prod_{j \neq i} \left[\frac{P_i \, R_i}{P_j \, R_j} \right]^{W_{ij}}$, where j is an index that runs over country i's trade partner (or competitor) countries; W_{ij} is the competitiveness weight attached by country i to country j, which are based on 1988–90 average data on the composition of trade in manufacturing, non-oil primary commodities and tourism services[8]; P_i and P_j are the seasonally-adjusted consumer price indices in countries *i* and *j*; and R_i and R_j are the nominal exchange rates of countries *i* and *j*'s currencies in U.S. dollars. As shown by McDermott (1996), alternative measures of the real exchange rate, such as real bilateral exchange rates based on consumer prices, and the *IFS*'s REER based on normalized unit labor costs, are both highly correlated with the *IFS*'s CPI-based REER index.

The REER data for six representative countries are set out in Figure 1—an increase in the REER series indicates a real appreciation of the country's currency.[9] Several features of the data stand out. First, a cursory inspection of the REER series indicates that most countries have real exchange rates that appear to exhibit symptoms of drift or nonstationarity. There appear to be substantial and sustained deviations from PPP (that is, nonstationarity in the REER). The evolution of REER appears to be a highly persistent, slow-moving process; for most countries the REER does not appear to cycle about any particular equilibrium value, with the possible exception of Sweden. Second, sharp movements in the REER during the 1980s and 1990s are a relatively frequent occurrence, especially for developing countries such as Indonesia, Nigeria and Mexico. We now describe the results for our analysis of the persistence of parity deviations for the REER series, using both biased least squares and median-unbiased estimators.

IV. Empirical Results

In this section we present our estimates of the persistence of parity deviations. We then examine the cross-country heterogeneity of the persistence results, by grouping the countries using several key characteristics, such as income level, type of nominal exchange rate regime, type of dominant exportable, and geographic location.

Biased least squares estimates of half-lives of parity reversion

The results for the half-life of the duration of shocks to the REER, calculated from the least squares estimates of α in ADF regression of equation (1), are set out in Table 1. Across all countries, the average (median) half-life of parity reversion is 4.04 years. This result is consistent with Rogoff's (1996) consensus of half-lives of parity reversion of between 36 to 60 months (three-to-five years), and with Cheung and Lai's (2000b) finding of median half-lives for industrial countries in the post-Bretton Woods period of 3.3 years.

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[8] W_{ij} can be interpreted as the sum over all markets of a gauge of the degree of competition between producers of country *i* and *j*, divided by the sum over all markets of a gauge of the degree of competition between producers of country *i* and all other producers.

[9] The 20 industrial countries and 70 developing countries are listed in Table 1. A decline (depreciation) in a country's REER index indicates a rise in its international competitiveness (defined as the relative price of domestic tradable goods in terms of foreign tradables). For a detailed explanation and critique of how the Fund's REER indices are constructed, see Zanello and Desruelle (1997) and Wickham (1993).

Median-unbiased estimates of half-lives of parity reversion

The half-lives of PPP deviations calculated above (using the least squares estimator) are reasonably close to past studies but are likely to be biased downward (and in favor of finding that PPP holds in the REER data). Consequently, we remove this bias by calculating median-unbiased estimates for the autoregressive parameter in equation (1).[10] [11]

The median-unbiased estimates of the AR parameter in ADF regressions are reported in Table 1. In comparison with their least squares counterparts, they are typically much longer, ranging from as little as one month (Bolivia) to infinity (Belgium and Sweden, among others). Across all countries, the average (median) half-life of parity reversion is 8.17 years, in excess of the average least-squares AR(p) half-life of 4.04 years.[12] This implies a rate of parity reversion of only 8 percent per year, rather than the 16 percent per year calculated using least squares.

Using the Andrews unbiased model-selection rule we find 50 of the countries are subject to finitely-persistent shocks to their REER (which is consistent with the reversion of REER to parity), while 40 of the countries experience permanent shocks to their REER series. The interpretation of this rule is that for any given country there is a 50 percent probability that the interval from zero to the estimated median-unbiased half-life contains the true half-life. There is also a 50 percent probability that the interval from the estimated median-unbiased half-life to infinity contains the true half-life. Let us take the examples of Iceland (short-lived half-life) and Togo (infinite half-life). For Iceland, while there is a 50 percent probability that the interval from zero to 1.1 years contains the true half-life, there is also a 50 percent probability that the interval from 1.1 years to infinity contains the true half-life. For Togo, while there is a 50 percent probability that the interval with a finite upper bound contains the

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[10] The method of bias-correction in the AR(p) regression of equation (1) works as follows. To obtain median-unbiased estimates of the parameters α, $\psi_1, \ldots, \psi_{p\text{-}1}$, μ (and associated impulse responses), Andrews and Chen's (1994, pp. 191–192) iterative procedure is used. This involves: compute the least squares estimates of equation (1), and obtain the least squares estimates of $\psi_1, \ldots, \psi_{p\text{-}1}$. Treat these estimates as though they were the true values of $\psi_1, \ldots, \psi_{p\text{-}1}$, and calculate the bias-corrected estimator of α, denoted as $\hat{\alpha}_u$, using equation (2). Then, treat $\hat{\alpha}_u$ as if it was the true value of α, compute a second round of least squares estimates of $\psi_1, \ldots, \psi_{p\text{-}1}$ in equation (1), and calculate a second-round bias-corrected estimator of α, $\hat{\alpha}_u$, again using equation (2). This procedure continues until convergence occurs, with these final, approximately median-unbiased estimators of the true parameter values denoted as $\hat{\alpha}_u, \hat{\psi}_{1,u}, \ldots, \hat{\psi}_{p-1,u}, \hat{\mu}_u$.

[11] Median-unbiased estimates of the half-life of a shock (in this case for T=348 observations) were determined using quantile functions of $\hat{\alpha}$ generated by numerical simulation (using 2,500 iterations), following the method suggested by Andrews and Chen (1994), for the ADF regression of equation (1).

[12] In calculating group and all-country median half-lives of deviation from parity, a permanent deviation is defined as one which is at least as long as the span of the data. In our sample, the data spans 29 years, so permanent deviation from parity (infinite shock) is set to equal 30 years in duration.

true half-life, there is a 50 percent probability that the true half-life will be infinite (Table 1). Using the Andrews unbiased model-selection rule, the finite (Iceland) and infinite (Togo) estimates of the half-lives indicate that while shocks to Iceland's REER are transitory, shocks to Togo's REER are best viewed as being permanent.

We conclude that the majority of countries in our sample have real exchange rates which do revert (albeit sometimes slowly) to PPP and thus that PPP holds in the post-Bretton Woods period. Our conclusion differs from that of some previous authors because we use an unbiased model-selection rule that has a probability of correctly selecting the unit root model (when the true α =1) of about 0.5. This contrasts with the Monte Carlo experiments of Sarno and Taylor (2002), who show standard method yields a probability of rejecting the null hypothesis of a unit root in the real exchange rate (when the rate is in fact mean reverting) of only 0.11.

Impulse response functions and half-lives of shocks to parity

The half-lives of shocks to parity have been calculated from the impulse response functions of equation (3), with the half-life defined as the time it takes for a unit impulse to dissipate permanently by one half from the occurrence of the initial shock (Cheung and Lai, 2000b). That is, the half-life is given by $IR(l_h) = ½$; it indicates the duration of time for the impact of a unit shock to permanently dissipate by half. The AR(p) model allows for shocks to dissipate in a nonmonotonic fashion, and several representative country cases are discussed in this section. In particular, we will examine the nature of each country's adjustment to shocks to their REER (that is, the shape of the impulse response functions).

Figure 2 sets out the impulse response functions derived from the least squares and median-unbiased estimation of equation (3), for three country pairs: Sweden and Indonesia (which exhibit permanent shocks to their real exchange rate); the United States and Nigeria (which exhibit hump-shaped IRF, yet with finite half-lives); and the United Kingdom and Mexico (which exhibit monotonic decay to real exchange rate shocks and finite half-lives). The half-life can be read from the intersection of the relevant IRF with the line indicating persistence of 0.5. The impact of high-frequency noise can be seen in the first few lags of the hump-shaped IRFs of the United States and Nigeria. The hump-shaped IRFs indicate that real exchange rate shocks are initially magnified. In both the cases of the United States and Nigeria, the magnification of the real exchange rate shock can last up to 1½ years, with the IRF taking about four years to return to unity.[13] As pointed out by Cheung and Lai (2000b), this nonmonotonicity in the response of the impulse response function to shocks results in a significant extension to the process of parity reversion (shock dissipation). Clearly, it is

[13] Following Cheung and Lai (2000b), the effects of non-monotonic reversion to parity can also be illustrated by calculating the half-life of real exchange rate shocks *after* the period of initial magnification. For both the United States and Nigeria, these modified half-life estimates are much shorter than the standard half-life of parity reversion, indicating that once the initial magnification period is completed, the impulse response functions dissipate rather rapidly. The duration of post-magnification half-lives can be calculated from Table 1, as the difference between the unbiased half-life (column 3) and the time to peak (column (4)).

important to control for this kind of serial correlation in real exchange rates—while the AR(p) regressions can do so, standard AR(1) regressions assume that shocks decay monotonically, and so would drastically underestimate the half-life of real exchange rate shocks.

In contrast to the above results, Taylor (2001) argues that there are two sources of *upward* bias in the conventional estimation of the speed of parity reversion: first, temporal aggregation bias, whereby sampling data at low frequencies does not allow one to identify a high-frequency adjustment process; and second, the linear AR(1) specification of the standard (Dickey-Fuller) unit root model, which assumes that reversion occurs monotonically, regardless of how far the process is from parity. However, as the present paper uses monthly REER data, the temporal aggregation bias is likely to be minimal. In addition, our use of AR(p) models allows for shocks to the REER to decline at a rate that is not necessarily constant.[14] [15]

Cross-country heterogeneity of half-lives of real exchange rate shocks

The results of our estimation of the median-unbiased half-life of reversion to PPP are shown in the histograms of Figures 3A to 3D. For our sample of 90 countries, there are 50 countries where the half-life estimates range from 1 to 14 years, and 40 countries that experience permanent real exchange rate shocks (half-lives of 29 years or more). Examining industrial countries alone, the average (median) half-life is about 8 years, which is twice as long as the downwardly biased estimates of previous studies (Rogoff, 1996; Cheung and Lai, 2000a).[16] In contrast, the half-life estimates for developing countries appear to be evenly spread with most of the estimated deviations from parity being permanent. Accordingly, median-unbiased estimates of half-lives tend to be much longer for developing countries than for industrial countries. This finding of relatively slower parity reversion for developing countries contradicts the results of Cheung and Lai (2000a). However, this finding is consistent with Froot and Rogoff (1995), who expected that the rapid economic growth often associated with low-income countries would induce such drastic changes in the relative price of tradables and

[14] While median-unbiased estimation allowing for AR(p) behavior does engender a nonmonotonic impulse response function, the shape of that impulse response function will be unaffected by the size of the shock to the real exchange rate. In contrast, methods that are robust to nonlinear adjustment to equilibrium allow for such an effect, yet do not correct for downwardly biased autoregressive parameters (see Taylor, Peel and Sarno, 2001, for a discussion of these issues).

[15] Chen and Rogoff (2003) and Cashin, Céspedes and Sahay (2004) point out that there are important real factors (such as real commodity-export prices) which might be expected to affect the equilibrium real exchange rate of both advanced and developing countries. Both studies find that the half-life of reversion of the real exchange rate to its (constant) long-run average level is much longer than the half-life of the reversion of the real exchange rate to its (time-varying) long-run equilibrium with real commodity prices. In this view, the long-run reversion of real exchange rates to purchasing power parity is a first approximation only (see Taylor, 2003).

[16] This result is also consistent with the median-unbiased half-lives calculated for industrial countries by Cashin and McDermott (2003).

nontradables that the likelihood of parity reversion holding in any given country would rise with the level of income. Finally, the hypothesis that the real exchange rate is mean reverting in developing countries receives some support—for 32 of 70 developing countries the half-life of parity reversion is finite (Table 1). In contrast, the hypothesis of mean reversion of real exchange rates for developed countries receives much stronger support—18 of 20 developed countries have real exchange rates that typically experience finite shocks.

When countries are classified by their dominant exportable, the half-life estimates for (nonfuel) primary commodity exporters appear rather dispersed (Figure 3A). This is consistent with earlier findings as to the heterogeneous persistence of terms of trade shocks affecting African commodity-exporting countries (see Cashin, McDermott and Pattillo, 2004). The average half-life for nonfuel primary-product exporters is found to be about 22 years—while finite, this implies only very slow speed of parity reversion.

The regional grouping of developing countries reveals that shock persistence of African and Western Hemisphere countries is evenly spread between finite and permanent shocks; in contrast, permanent shocks dominate in Asian countries (Figures 3A and 3B). The existing literature on parity reversion in developing countries has a sample-selection problem, as it has typically concentrated on analyses of high-inflation Latin American countries. Consistent with that literature, we find that parity deviations in the often-studied countries of the Western Hemisphere (especially Brazil, Mexico, Chile, and Argentina) are rather short-lived (with half-lives all less than four years). Not surprisingly, the parity reversion half-lives for Heavily Indebted Poor Countries (HIPCs), which are mostly African, resemble the African results—the average half-life for HIPC countries is found to be about 11 years.

Net debtor developing countries have also been grouped by source of external financing—for both private external financing and official external financing, real exchange rate shocks are typically permanent. Country groups of different income levels also exhibit a systematic pattern of differences in the persistence of real exchange rate shocks (Figure 3C)—our results tend to confirm previous work finding an inverse relationship between income level and shock persistence (see Cheung and Lai, 2000a).

We also examine the persistence of parity deviations after classifying countries by the type of nominal exchange rate regime. We classify country real exchange rates by type of nominal exchange rate regime, using: (i) the IMF's (1998) *de jure* classification, which is based on the publicly stated commitment of the authorities of the country in question; and (ii) the *de facto* classification of Reinhart and Rogoff (2002), which is based on the observed behavior of market-determined real exchange rates, including that of active parallel exchange rate markets. It is important to examine both *de facto* and *de jure* exchange rates, because through exchange market intervention and/or monetary policy, the authorities of any given country can transform a *de jure* flexible exchange rate regime into a *de facto* pegged regime. Similarly, active parallel markets can transform *de jure* pegged official exchange rates into *de facto* flexible regimes.

When all 90 countries are categorized by the IMF's *de jure* exchange rate classification rules, it is clear that pegged exchange rate countries typically experience permanent deviations from parity, while more flexible exchange rate countries have much more dispersed half-lives of deviations from parity (Figure 3D). Very similar results are derived when exchange rates are classified using the Reinhart-Rogoff (2002) *de facto* exchange rate classification. The *de facto* classification reveals that while permanent real exchange rate shocks dominate the experience of pegged exchange rate countries, parity deviations for the majority of flexible exchange rate countries are finite, and typically range between two to ten years.

V. What Is Causing the Heterogeneous Persistence of Deviations from Parity?

Underpinning the fundamental idea of long-run PPP is that arbitrage in goods ensures the parity condition is satisfied across a range of goods over a certain time horizon. A key question to be answered is what might be causing the heterogeneous duration of exchange rate deviations from parity. In particular, in this section we will investigate the differences in the speed of parity reversion when comparing advanced (industrial) and developing countries, and when comparing countries with fixed and flexible nominal exchange rate regimes. Following Cheung and Lai (2000a), we examine whether the observed pattern of the persistence of real exchange rate shocks can be linked to any systematic differences in structural characteristics across countries.

Given that the usual distributional assumption of normality is not likely to hold for the distribution of either the half-lives of deviations from parity or the structural characteristics of countries, we use several nonparametric tests to examine the statistical significance of various hypotheses. These concern equality of the average duration of parity deviations across country groups; equality of the variability of the average duration of parity deviations across country groups; and whether the pattern of persistence of parity deviations is correlated with countries' structural characteristics.

First, we implement a nonparametric test of the equality of the median half-life of deviations from parity across country groups (the Wilcoxon-Mann-Whitney test). The results (listed in column (4) of Table 2)) indicate that the null hypothesis of equal median half-lives is rejected, at the 5 percent significance level, for the advanced versus developing countries. Further, the null hypothesis of equal median half-lives is rejected, at the 5 percent significance level, for the pegged versus flexible exchange rate countries (classified under either the IMF (1998) or Reinhart-Rogoff (2002) classification schemes). Accordingly, we conclude that the average half-life of real exchange rate deviations from parity for advanced (industrial) countries are shorter in duration than those for developing countries. In addition, the average half-life for countries with pegged nominal exchange rates are significantly longer in duration than those for countries with flexible nominal exchange rates. When examining the persistence of real exchange rate shocks, the statistical significance of the difference of the group medians confirms the pattern exhibited in Figures 3A–3D—industrial countries display less persistence than developing countries; and countries with fixed nominal exchange rate regimes display more persistence than countries with flexible nominal exchange rate regimes.

Second, we examine whether the differences in the sample variances of the half-lives across country groups are statistically significant, by implementing the Brown-Forsythe test. The results (listed in column (5) of Table 2)) indicate that the null hypothesis of equal variances of the half-lives is rejected, at the 5 percent significance level, for the industrial countries in comparison with developing countries. Interestingly, the null of equality of the variances of the half-lives across pegged and flexible exchange rate regimes cannot be rejected. Accordingly, there is evidence that the duration of half-lives of parity deviations for developing countries is more variable than the duration of half-lives of parity deviations for industrial countries.

Third, we use our cross-country data to examine whether each country's persistence of deviations from parity is correlated with its structural characteristics. As normality is unlikely to hold here, we again use a nonparametric test—the Spearman rank correlation test (see Cheung and Lai (2000a) and Conover (1999) for details). The Spearman rank correlation statistic measures whether there is a significant relation between the persistence of parity deviations and key structural characteristics (such as cross-country differentials in inflation and trade openness). The null hypothesis of the Spearman rank correlation test is that there is no rank correlation between the persistence of parity deviations and differentials in each country's structural characteristics. In explaining empirically cross-country heterogeneity in the duration of real exchange rate deviations from parity, we will evaluate several fundamental characteristics (see Appendix I for details), including: cross-country differentials in inflation (INF); nominal exchange rate arrangements (VOFFER and VPARER); openness to trade (TGDP); productivity growth (PCGDP); and the share of government spending in the economy (GGDP).

Inflation

If national price movements were dominated by nominal (monetary) shocks, then parity reversion would be expected to be rather fast. Given the presence of nominal rigidities, a higher inflation rate may bring about more frequent adjustment of goods prices, and accordingly shrink the duration of deviations from parity. Indeed, previous work has indicated that PPP typically holds for high-inflation countries (Frenkel, 1978; McNown and Wallace, 1989). To analyze whether relative inflation is related to cross-country differences in persistence of parity deviations, we construct the average (median) inflation rate over the sample period (INF) for each country. The rank correlation coefficient between the half-life of the real exchange rate and the rate of inflation is -0.281, with an approximate p-value of 0.007 (Table 3). Using the critical values of Zar (1972), the null of no rank correlation is decisively rejected, with the rank correlation being negative and statistically significant (at the 1 percent level). This indicates that across advanced and developing countries, there is an inverse relationship between inflation and the persistence of real exchange rate shocks. That is, countries with higher inflation rates tend to have shorter-lived deviations of their real exchange rates from PPP, which suggests that parity reversion is fast when price movements largely reflect monetary shocks.

Nominal exchange rate volatility

Mussa (1986) stressed that exchange rates behaved differently under alternative exchange rate regimes, finding that the post-Bretton Woods float of major currencies had induced large real exchange rate variability in many industrial countries. Greater flexibility in nominal exchange rates would be expected to increase the speed of parity reversion of real exchange rates, by encouraging more frequent adjustment of goods prices. However, as noted above it is important to measure the variability of nominal exchange rates using both the official exchange rate (VOFFER), and the exchange rate determined in parallel markets (VPARER). In doing so we use both the official and parallel exchange rate market data of Reinhart and Rogoff (2002), for the period 1973–98, with variability measured as the standard deviation of the (monthly) rate of change of the (log) series.

The rank correlation coefficient between the half-life of the real exchange rate and the variability of the official nominal exchange rate is -0.379, with an approximate *p*-value of 0.002 (Table 3). Similarly, the rank correlation coefficient between the half-life the real exchange rate and the variability of the parallel-market nominal exchange rate is -0.282, with an approximate *p*-value of 0.02. Both correlations are negative and statistically significant (at the 1 percent level). This indicates that across advanced and developing countries, there is an inverse relationship between nominal exchange rate variability and the persistence of real exchange rate shocks. That is, countries with more variable nominal exchange rates tend to have shorter-lived deviations of their real exchange rates from PPP. The above results are broadly consistent with the conclusions of Goldfajn and Valdes (1999), who find that overvaluation of real exchange rates were typically corrected by changes in the nominal exchange rate rather than changes in inflation differentials.

Productivity growth and government spending

Productivity growth is a supply-side factor which can affect the persistence of real exchange rate shocks. The Balassa-Samuelson effect has traditionally been the most popular explanation for the persistence of parity deviations. The Balassa-Samuelson hypothesis highlights the potential effects of differential productivity growth (favoring the traded goods sector) on the behavior of real exchange rates, which ultimately raises the relative price of nontraded goods. Importantly, this traded-goods productivity bias is deemed to rise with the wealth of the country. Following Balassa (1964), we proxy for productivity growth by using the rate of growth of per capita real GDP; accordingly, we construct the average (median) per capita growth rate over the sample period for each country (PCGDP).

In the presence of capital and labor which are mobile across sectors in the long run but not the short run, government spending is a demand-side factor which is also hypothesized to affect the speed of parity reversion, by producing a stronger home goods bias. Froot and Rogoff (1991) point out that government spending typically falls more heavily on nontraded goods, thereby bidding up their price relative to the price of tradables. In turn, Bergin and Feenstra (2001) suggest that real exchange rate persistence rises with the share of produced goods that are nontraded. Accordingly, we construct the average (median) for government spending as a share of GDP over the sample period for each country (GGDP).

The rank correlation coefficient between the half-life of the real exchange rate and the growth of per capita GDP is -0.038, with an approximate *p*-value of 0.77 (Table 3). Similarly, the rank correlation coefficient between the half-life of the real exchange rate and government spending (as a share of GDP) is -0.073, with an approximate *p*-value of 0.50. Neither rank correlation is significantly different from zero. We thus conclude that there is little evidence that either higher productivity growth or greater government spending can explain much of the observed pattern of deviations from PPP across countries. Our conclusions are consistent with that of Rogoff (1996).

Trade openness

Rogoff , Froot and Kim (2001) highlight the effect of goods market arbitrage, which if operative can accelerate the speed of parity reversion across goods. Greater trade flows should in principle promote goods market arbitrage, encourage more frequent price adjustment by firms and thereby reduce the persistence of real exchange rate shocks (Faruqee, 1995). As a measure of trade openness, we construct for each country over the sample period the average (median) ratio of exports plus imports to GDP (TGDP). The rank correlation coefficient between the half-life of the real exchange rate and external trade (as a share of GDP) is 0.180, with an approximate *p*-value of 0.10 (Table 3). This rank correlation differs insignificantly from zero. Accordingly, there is little evidence in favor of the goods arbitrage view of PPP—differences in trade openness are not associated with the persistence of parity deviations across countries.[17]

To summarize, we find that cross-country differences in inflation rates and nominal exchange rate variability have a strong and statistically significant (inverse) relationship with the observed pattern of the persistence of deviations from purchasing power parity.[18] Indeed, to the extent that the existing literature on measuring parity reversion in developing countries has largely concentrated on the Latin American experience (which is dominated by high-inflation countries), this sample-selection bias can account for the (erroneous) received wisdom that parity reversion in developing countries is faster than that observed for developed countries. However, once a wider set of developing countries is analyzed, parity reversion in developing countries is typically slower than that observed for developed countries.

VI. CONCLUSION

The validity of purchasing power parity (PPP)—the notion that prices in different countries move towards equality in common currency terms—is of interest to policymakers for two main reasons. First, PPP provides a long-run benchmark for the equilibrium value of exchange rates, and as such is a criterion for evaluating the competitiveness of real exchange rates. Second, PPP has been adopted as a central building block of many theories of

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[17] Cheung et al. (2001) also find a non-negative relationship between openness and the persistence of sectoral real exchange rate deviations. They attribute this result to the inverse relationship between openness and inflation (see Romer, 1993). For our sample of countries, we also find a strong negative rank correlation between openness and inflation (-0.397). Accordingly, this suggests that greater openness is associated with lower inflation, which implies a slower speed of parity reversion.

[18] As a robustness check on the unconditional correlation results, least squares regression of the half-life of parity deviation on the abovementioned structural characteristics was also carried out. As in the bivariate correlations, inflation is the most significant variable for reducing the half-life of parity deviations.

exchange rate determination; the quality of policy advice based on these theories may be dependent on the validity of PPP.

This paper has re-examined whether PPP holds during the post-Bretton Woods period, by investigating the time series properties of the real effective exchange rate of 90 advanced and developing countries. The post-1973 revival of flexible exchange rates spawned a great interest in the empirical relevance of the PPP theory of real exchange rate determination. Previous studies of PPP reversion largely focused on developed countries, and univariate studies of the hypothesis of unit roots in real exchange rates yielded consensus estimates of the half-life of deviations of real exchange rates from PPP of about four years (Rogoff, 1996).

Using least squares estimation of unit root models, we replicate the consensus finding in the literature. However, using median-unbiased estimation techniques that remove the downward bias of least squares, we find that the half-lives of parity reversion are much longer than the consensus estimate, with the cross-country average of unbiased half-lives of deviations from parity lasting about eight years. Our results confirm Rogoff's (1996) 'PPP puzzle'—that while PPP holds for the majority of countries, the speed of reversion of real exchange rates to parity is, in many cases, rather slow.

Importantly, using the median-unbiased estimates of the half-lives of deviations from parity and the Andrews (1993) unbiased model-selection rule, we can be more definitive about our willingness to draw conclusions as to the presence or absence of parity reversion of real exchange rates in the post-Bretton Woods period. We conclude that 50 of the 90 countries in our sample have finite half-lives of parity reversion, which indicates that there is a better than even chance that shocks to their real exchange rates are transitory. Consequently, for these countries we can conclude that there is reversion of real exchange rates to parity, and so PPP holds in the post-Bretton Woods period. Conversely, there is little evidence of parity reversion for the remaining 40 countries in our sample, where shocks to their real exchange rate are best viewed as being permanent.

Our analysis yields evidence of significant heterogeneity of parity reversion across countries and across groups of countries. In our view, the general relevance of parity reversion has been exaggerated by the predilection of existing studies to focus on the reversion experience of developed countries. In addition, those (few) studies of parity reversion for developing countries have suffered from sample-selection bias, in that they typically examined the speed of reversion in high-inflation Latin American countries. We find that parity reversion is more likely to be found for developed countries than for developing countries. Parity reversion is also more likely to be found for countries with flexible nominal exchange rate regimes than for countries with fixed nominal exchange rate regimes. Finally, when we examine the determinants of the observed cross-country heterogeneity in the persistence of reversion of real exchange rates to parity, we find that parity reversion tends to be faster in high-inflation countries than in low-inflation countries, while parity reversion tends to be slower in countries with less nominal exchange rate variability.

Description of the Data

The primary data sources are the IMF's *International Financial Statistics* (IFS) and *Information Notice System* (INS), Reinhart and Rogoff (2002), and the World Bank's (2002) *World Development Indicators*. Below we provide a description of each series used in the paper. There are 90 developing and developed countries in the full sample, which are listed in Table 1. The data are for the sample period 1973–2002 (unless otherwise denoted).

REER: The real effective exchange rate data are of monthly frequency, for the period March 1973–March 2002; a total of 358 observations. REER is the trade-weighted measure of the seasonally adjusted, CPI-based real effective exchange rate (base 1990=100); obtained from the IMF's INS.

INF: The rate of change of consumer prices (percent per annum); period average of annual data 1973–2002; obtained from the IMF's IFS.

GGDP: General government final consumption spending as a share of GDP; period average of annual data 1973–2002; obtained from World Bank (2002).

PCGDP: Growth of per capita GDP (percent per annum); period average of annual data, 1973–2002; obtained from World Bank (2002).

TGDP: Exports and imports of goods and services (valued in current U.S. dollars) as a share of GDP (valued in current U.S. dollars); period average of annual data, 1973–2002; obtained from World Bank (2002).

VPARER: Volatility of the parallel market exchange rate; measured as the standard deviation of the (monthly) rate of change of the parallel market exchange rate, with the exchange rate measured as the logarithm of the parallel nominal exchange rate (local currency per U.S. dollar), March 1973 to December 1998; obtained from Reinhart-Rogoff (2002).

VOFFER: Volatility of the official exchange rate; measured as the standard deviation of the (monthly) rate of change of the official exchange rate, with the exchange rate measured as the logarithm of the official nominal exchange rate (local currency per U.S. dollar), March 1973 to December 1998; obtained from Reinhart-Rogoff (2002).

Nominal Exchange Rate Regime:

The IMF's *de jure* classification used between 1975–97 in its *Annual Report on Exchange Arrangements and Exchange Restrictions* (AREAER) consisted of ten categories, grouped into: regimes 1–5 are defined as fixed pegs; regimes 6–7 (limited flexibility with respect to a single currency, cooperative arrangements) are intermediate, and regimes 8–10 (including managed floating and independently floating) are flexible arrangements. For each country, the *de jure* exchange rate regime classification is the mode of the annual IMF (AREAER) classification numbers over the period 1975–98. See Bubula and Otker-Robe (2002) for additional details.

The Reinhart-Rogoff (2002) *de facto* classification describes exchange rate regimes as: (i) de facto pegs (including no separate legal tender and currency boards), denoted as regime 1; (ii) limited flexibility (including crawling pegs and narrow crawling bands), denoted as regime 2; (iii) managed floating (including wider crawling bands), denoted as regime 3; (iv) freely floating, denoted as regime 4; and (v) freely falling (where the annualized rate of inflation exceeds 40 percent), denoted as regime 5. For each country, the *de facto* exchange rate regime classification is the mode of the annual Reinhart-Rogoff (2002) classification numbers over the period 1973–98. The following countries had no Reinhart-Rogoff (2002) classification data: Ethiopia, Fiji, Papua New Guinea, Seychelles, Sudan, Samoa, Democratic Republic of Congo, Barbados, Rwanda, Sierra Leone, and Trinidad and Tobago.

Country Group Classifications

The 90 developed and developing countries in our sample (as listed in Table 1) have been classified into various country groups, in order to undertake cross-country comparisons of the persistence of parity deviations. The major sources of classification were: the IMF's *World Economic Outlook* (2002, 2000); the IMF's *Annual Report on Exchange Arrangements and Exchange Restrictions* (AREAER), various issues; the World Bank's (2002) *World Development Indicators* database; Andrews et al. (1999); and Reinhart and Rogoff (2002). The country groups, along with the country members and classification rule, are as follows.

The IMF's *World Economic Outlook* classifies countries into groups, based on certain criteria (see IMF (2002)). The groups are (for non-developing countries): *advanced economies*; and *countries in transition*. Developing countries are classified by their predominant export as: *primary product exporters* (those countries whose exports of agricultural and mineral primary products (Standard Industrial Trade Classification 0, 1, 2, 4, 68) accounted for at least 50 percent of their total export earnings during 1994–98); and *fuel exporters* (those countries whose exports of fuel products (SITC 3) accounted for at least 50 percent of their total export earnings during 1994–98). *Net debtor countries* are defined as developing countries with negative external assets at the end of 1998. Net debtor countries are then differentiated by their main source of external financing—net debtor countries with official financing (including official grants) accounting for more than two-thirds of their total 1994–98 external financing are classified as *official external financing countries*; net debtor countries with private financing (including direct and portfolio investment) accounting for more than two-thirds of their total 1994–98 external financing are classified as *private external financing countries*.

Advanced economies: Australia, Austria, Belgium, Canada, Finland, Germany, Iceland, Ireland, Italy, Japan, Korea, Netherlands, Norway, New Zealand, Portugal, Spain, Sweden, Switzerland, United Kingdom, United States [IMF (2002) classification].

Developing countries: Argentina, Brazil, Cameroon, Central African Republic, Chile, Côte d'Ivoire, Egypt, Ethiopia, Fiji, The Gambia, Guyana, Haiti, Honduras, Hungary, India, Indonesia, Islamic Republic of Iran, Kenya, Madagascar, Malawi, Malaysia, Mauritania, Mauritius, Morocco, Myanmar, Niger, Pakistan, Papua New Guinea, Paraguay, Philippines, Senegal, Seychelles, Sri Lanka, Sudan, Syria, Tanzania, Thailand, Togo, Tunisia, Turkey, Samoa, Democratic Republic of Congo, Zambia, Barbados, Bolivia, Burkina Faso, Chad, Colombia, Republic of Congo, Dominican Republic, Ecuador, El Salvador, Gabon, Ghana, Guatemala, Jamaica, Mexico, Malta, Nepal, Nigeria, Peru, Rwanda, Sierra Leone, Trinidad and Tobago, Uruguay, Venezuela, Costa Rica, Lesotho, Panama, and Uganda [IMF (2002) classification].

Nonfuel primary-product exporting countries: Australia, Canada, Iceland, New Zealand, Central African Republic, Chile, Cote d'Ivoire, The Gambia, Guyana, Honduras, Madagascar, Malawi, Mauritania, Myanmar, Niger, Papua New Guinea, Paraguay, Sudan, Tanzania, Togo, Democratic Republic of Congo, Zambia, Bolivia, Burkina Faso, Chad, Ghana, Peru [IMF (2002) classification].

Fuel-exporting countries: Norway, Islamic Republic of Iran, Republic of Congo, Gabon, Nigeria, Trinidad and Tobago, Venezuela [IMF (2002) classification].

African developing countries: Cameroon, Central African Republic, Côte d'Ivoire, Ethiopia, The Gambia, Kenya, Madagascar, Malawi, Mauritania, Mauritius, Morocco, Niger, Senegal, Seychelles, Sudan, Tanzania, Togo, Tunisia, Democratic Republic of Congo, Zambia, Burkina Faso, Chad, Congo, Gabon, Ghana, Nigeria, Rwanda, Sierra Leone, Lesotho, and Uganda [IMF (2002) classification].

Asian developing countries: Fiji, India, Indonesia, Malaysia, Myanmar, Pakistan, Papua New Guinea, Philippines, Sri Lanka, Thailand, Samoa, and Nepal [IMF (2002) classification].

Western Hemisphere developing countries: Argentina, Brazil, Chile, Guyana, Haiti, Honduras, Paraguay, Barbados, Bolivia, Colombia, Dominican Republic, Ecuador, El Salvador, Guatemala, Jamaica, Mexico, Peru, Trinidad and Tobago, Uruguay, Venezuela, Costa Rica, and Panama [IMF (2003) classification].

Middle East and Turkey developing countries: Egypt, Islamic Republic of Iran, Syria, Turkey, and Malta [IMF (2002) classification].

Heavily Indebted Poor Countries (HIPCs): Comprises those countries (except Nigeria) considered by the IMF and World Bank for their HIPC debt initiative—Cameroon, Central African Republic, Côte d'Ivoire, Ethiopia, Guyana, Honduras, Kenya, Madagascar, Malawi, Mauritania, Myanmar, Niger, Senegal, Sudan, Tanzania, Togo, Democratic Republic of Congo, Zambia, Bolivia, Burkina Faso, Chad, Republic of Congo, Ghana, Rwanda, Sierra Leone, and Uganda [IMF classification, see Andrews et al. (1999)].

Developing Countries—Net Debtor Countries, Private External Financing: Argentina, Brazil, Chile, Egypt, Fiji, India, Indonesia, Islamic Republic of Iran, Kenya, Malaysia, Morocco, Myanmar, Paraguay, Seychelles, Thailand, Turkey, Colombia, Dominican Republic, Ecuador, Guatemala, Jamaica, Mexico, Malta, Peru, Sierra Leone, Trinidad and Tobago, Venezuela, Costa Rica, Lesotho, and Panama [IMF (2000) classification].

Developing Countries—Net Debtor Countries, Official External Financing: Cameroon, Central African Republic, Ethiopia, The Gambia, Guyana, Haiti, Madagascar, Malawi, Mauritania, Niger, Senegal, Tanzania, Togo, Samoa, Democratic Republic of Congo, Zambia, Burkina Faso, Chad, Republic of Congo, Gabon, Guatemala, Nepal, Rwanda, and Uganda [IMF (2000) classification].

Countries with Annual Income (real GDP per capita in U.S. dollars in 1995) of $102–$280: Ethiopia, Madagascar, Malawi, Niger, Sudan, Tanzania, Democratic Republic of Congo, Burkina Faso, Chad, Nepal, Nigeria, Rwanda, and Sierra Leone [World Bank (2002) classification].

Countries with Annual Income (real GDP per capita in U.S. dollars in 1995) of $281–$769: Cameroon, Central African Republic, Côte d'Ivoire, The Gambia, Haiti, Honduras, India, Kenya, Mauritania, Pakistan, Senegal, Sri Lanka, Togo, Zambia, Ghana, and Uganda [World Bank (2002) classification].

Countries with Annual Income (real GDP per capita in U.S. dollars in 1995) of $770–$2,111: Egypt, Indonesia, Islamic Republic of Iran, Morocco, Papua New Guinea, Paraguay, Philippines, Syria, Tunisia, Bolivia, Republic of Congo, Dominican Republic, Ecuador, El Salvador, Guatemala, and Jamaica [World Bank (2002) classification].

Countries with Annual Income (real GDP per capita in U.S. dollars in 1995) of $2,112–$5,792: Brazil, Chile, Hungary, Malaysia, Mauritius, Thailand, Turkey, Colombia, Gabon, Mexico, Peru, Trinidad and Tobago, Uruguay, Venezuela, and Costa Rica [World Bank (2002) classification].

Countries with Annual Income (real GDP per capita in U.S. dollars in 1995) of $5,793–$15,891: Portugal, Spain, Argentina, Korea, and Barbados [World Bank (2002) classification].

Countries with Annual Income (real GDP per capita in U.S. dollars in 1995) of $15,892–$43,600: Australia, Austria, Belgium, Canada, Finland, Germany, Iceland, Ireland, Italy, Japan, Netherlands, Norway, New Zealand, Sweden, Switzerland, United Kingdom, United States [World Bank (2002) classification].

IMF Pegged Exchange Rate Countries: Austria, Finland, Norway, Sweden, Argentina, Cameroon, Central African Republic, Côte d'Ivoire, Egypt, Ethiopia, Fiji, Guyana, Haiti, Honduras, Hungary, Islamic Republic of Iran, Kenya, Malawi, Malaysia, Mauritania, Mauritius, Morocco, Myanmar, Niger, Papua New Guinea, Paraguay, Senegal, Seychelles, Sudan, Syria, Tanzania, Thailand, Togo, Samoa, Zambia, Barbados, Burkina Faso, Chad, Republic of Congo, Dominican Republic, El Salvador, Gabon, Guatemala, Malta, Rwanda, Trinidad and Tobago, Venezuela, Lesotho, and Panama [IMF (AREAER classification), see Appendix I].

IMF Intermediate Exchange Rate Countries: Belgium, Germany, Ireland, Italy, Netherlands, Nepal, [IMF (AREAER classification), see Appendix I].

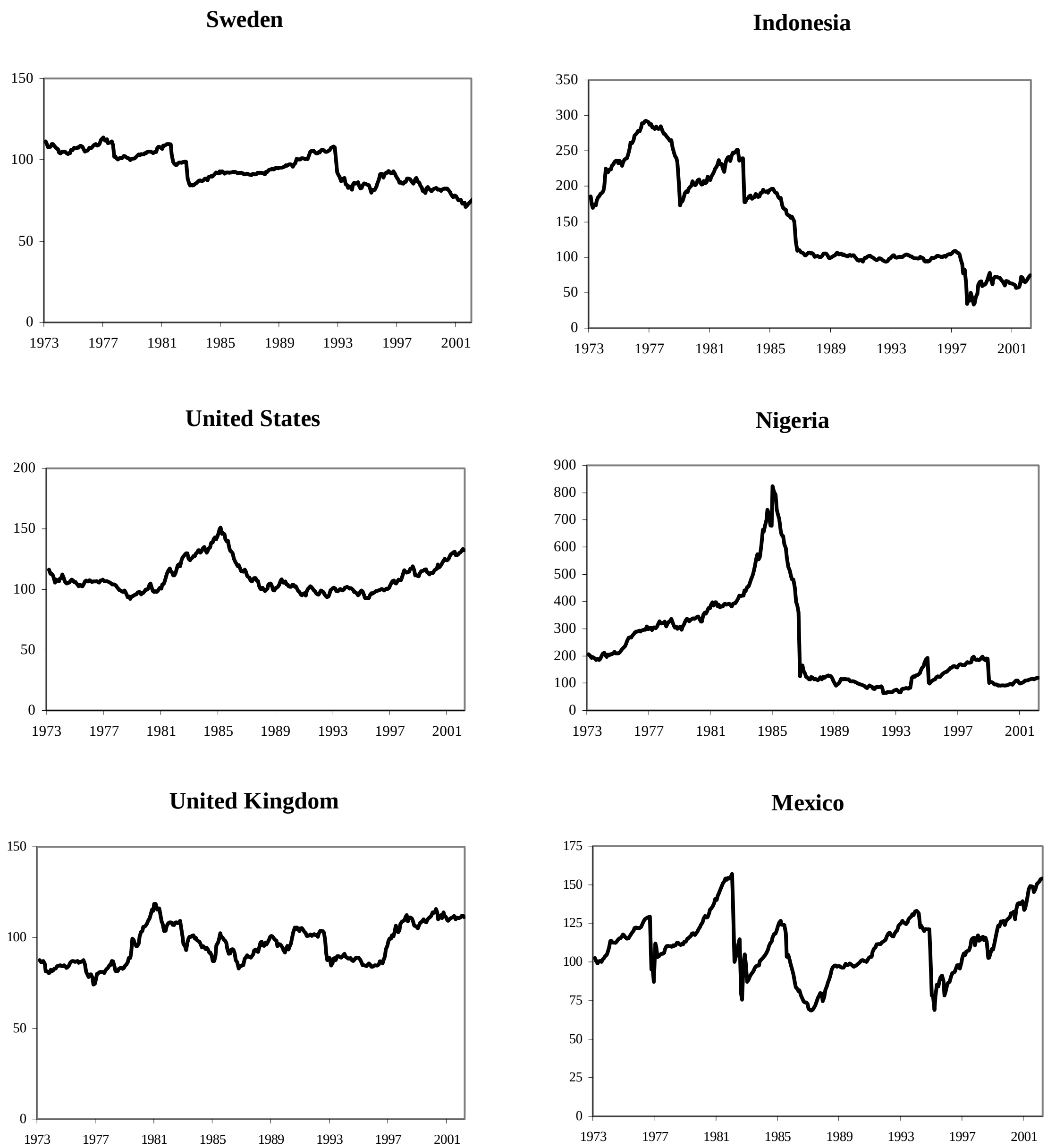
IMF Flexible Exchange Rate Countries: Australia, Canada, Iceland, Japan, New Zealand, Portugal, Spain, Switzerland, United Kingdom, United States, Brazil, Chile, The Gambia, India, Indonesia, Korea, Madagascar, Pakistan, Philippines, Sri Lanka, Tunisia, Turkey, Democratic Republic of Congo, Bolivia, Colombia, Ecuador, Ghana, Jamaica, Mexico, Nigeria, Peru, Sierra Leone, Uruguay, Costa Rica, and Uganda [IMF (AREAER classification), see Appendix I].

Reinhart-Rogoff Pegged Exchange Rate Countries: Austria, Belgium, Netherlands, Cameroon, Central African Republic, Côte d'Ivoire, Haiti, Kenya, Niger, Senegal, Thailand, Togo, Burkina Faso, Chad, Ecuador, El Salvador, Gabon, Guatemala, Mexico, Venezuela, Lesotho, and Panama [Reinhart and Rogoff (2002) classification, see Appendix I].

Reinhart-Rogoff Limited Flexibility and Managed Float Exchange Rate Countries: Canada, Finland, Iceland, Ireland, Italy, Norway, New Zealand, Portugal, Spain, Sweden, Switzerland, United Kingdom, Chile, Egypt, The Gambia, Guyana, Honduras, Hungary, India, Indonesia, Islamic Republic of Iran, Korea, Madagascar, Malawi, Malaysia, Mauritania, Mauritius, Morocco, Myanmar, Pakistan, Paraguay, Philippines, Sri Lanka, Syria, Tanzania, Tunisia, Bolivia, Colombia, Dominican Republic, Jamaica, Malta, Nepal, Nigeria, Costa Rica, and Uganda [Reinhart and Rogoff (2002) classification, see Appendix I].

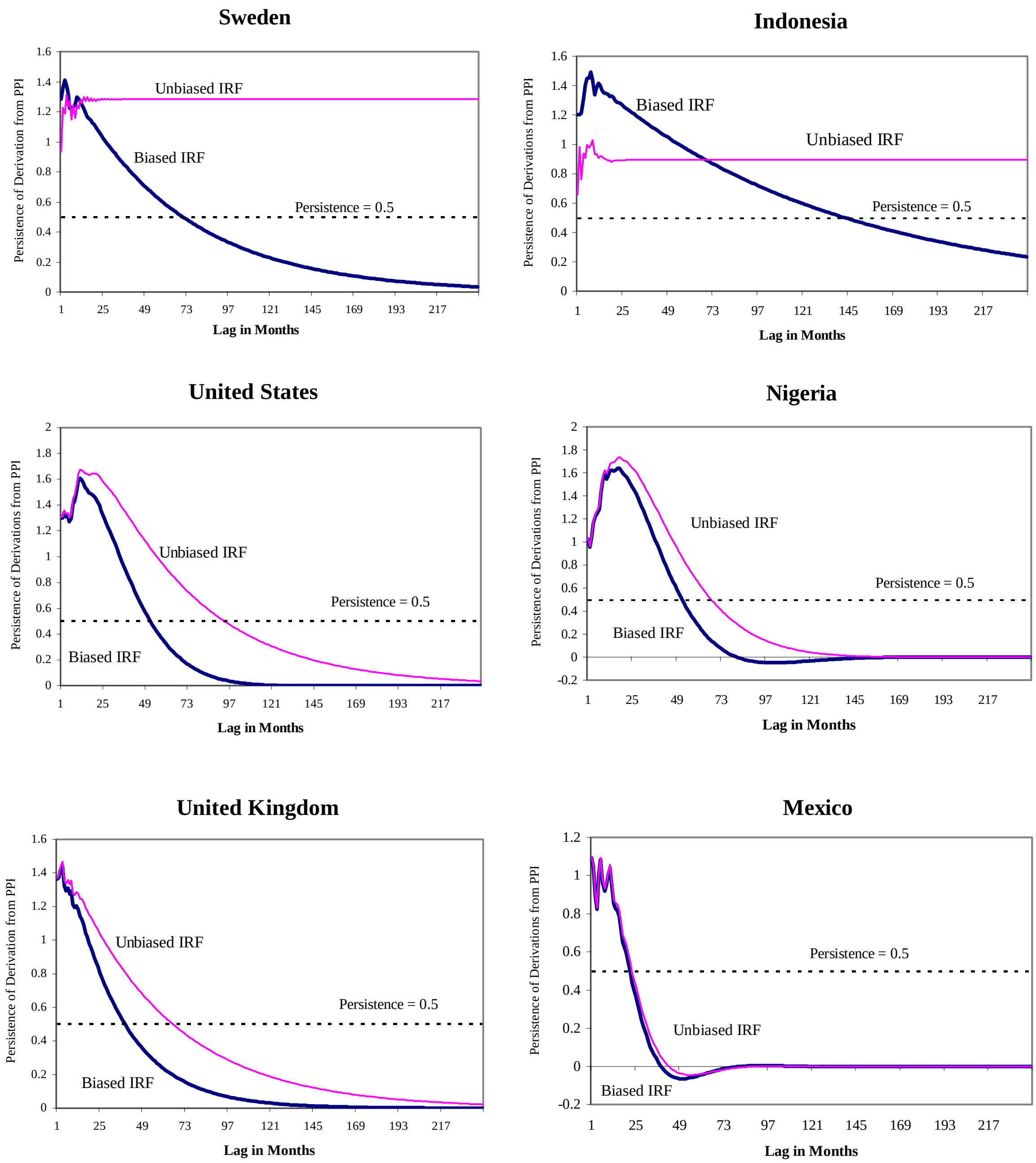
Reinhart-Rogoff Freely Floating/Falling Exchange Rate Countries: Australia, Germany, Japan, United States, Argentina, Brazil, Turkey, Zambia, Republic of Congo, Ghana, Peru, and Uruguay [Reinhart and Rogoff (2002) classification, see Appendix I].

Figure 1. Real Effective Exchange Rate, Selected Countries, 1973:3 - 2002:3, (1990 = 100)



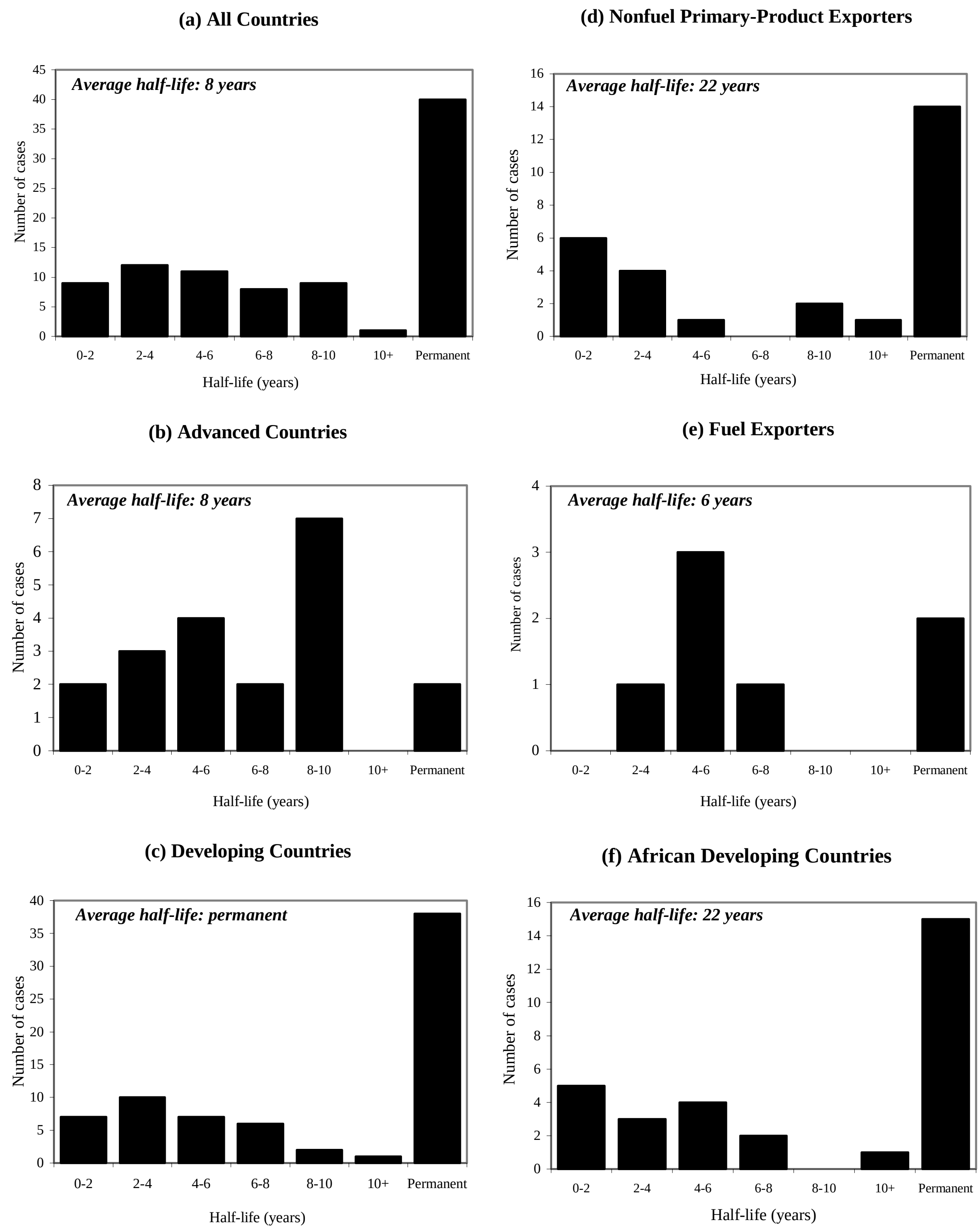
Source: International Monetary Fund.

Figure 2. Impulse Response Function (IRF) of a Shock to the Real Effective Exchange Rate, Selected Countries, 1973:3 - 2002:3



Source: Authors' calculations.

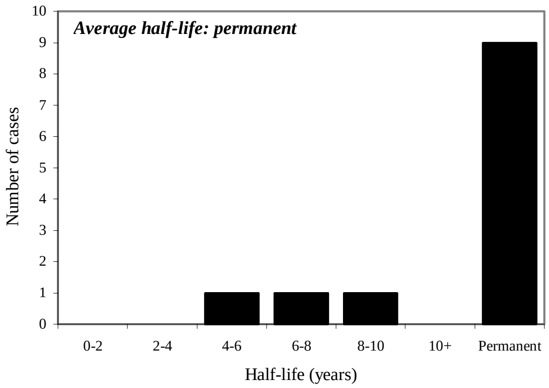
Figure 3A. Frequency Distribution of Median-Unbiased Half-lives of Deviations from Purchasing Power Parity, Country Groups, 1973-2002



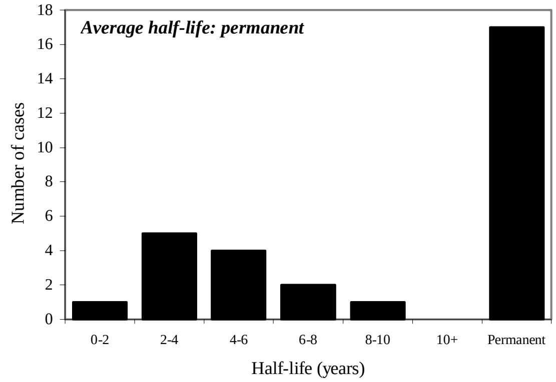
Source: Authors' calculations.

Figure 3B. Frequency Distribution of Median-Unbiased Half-lives of Deviations from Purchasing Power Parity, Country Groups, 1973-2002

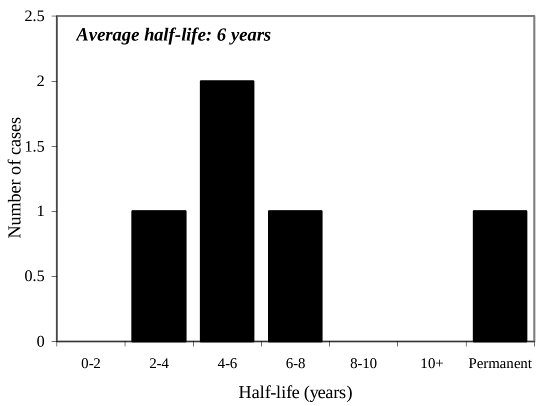
(g) Asian Developing Countries



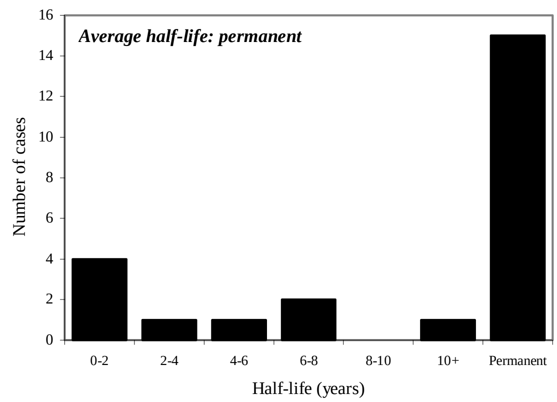
(j) Developing Countries: Net Debtor Countries, Private External Financing



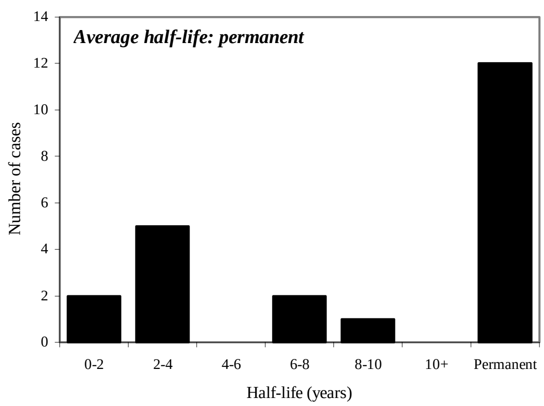
(h) Middle East and Turkey Developing Countries



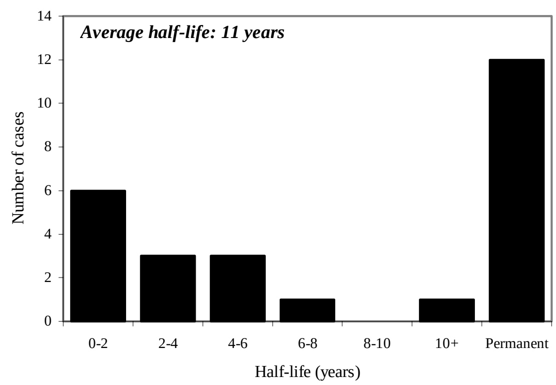
(k) Developing Countries: Net Debtor Countries, Official External Financing



(i) Western Hemisphere Developing Countries

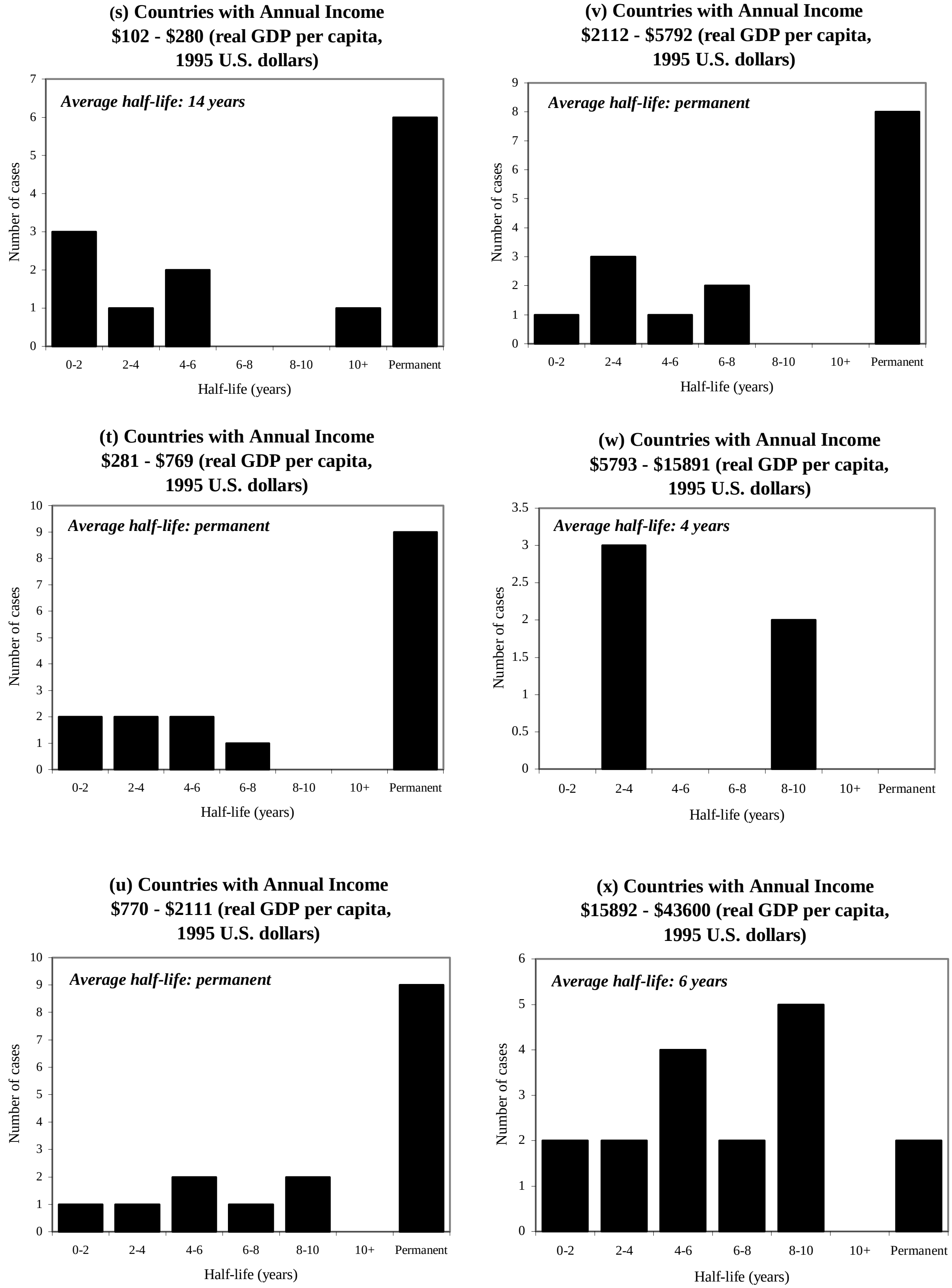


(l) HIPC Countries



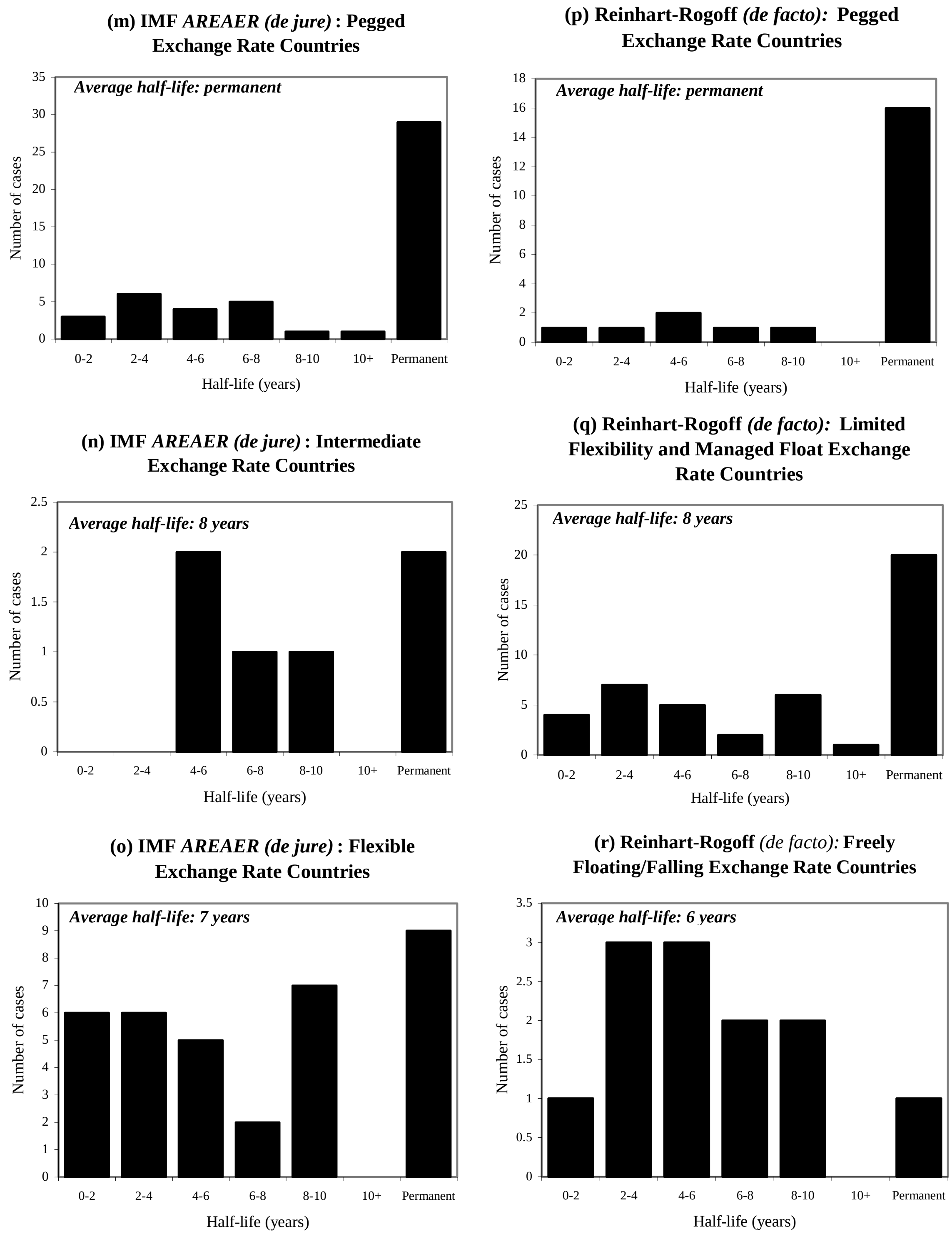
Source: Authors' calculations.

Figure 3C. Frequency Distribution of Median-Unbiased Half-lives of Deviations from Purchasing Power Parity, Country Groups, 1973-2002



Source: Authors' calculations.

Figure 3D. Frequency Distribution of Median-Unbiased Half-lives of Deviations from Purchasing Power Parity, Country Groups, 1973-2002



Source: Authors' calculations.

Table 1. Half-Lives of Reversion to Parity (years), Real Effective Exchange Rates, March 1973–March 2002, Advanced and Developing Countries

Country (1)	Biased Half-life (2)	Unbiased Half-life (3)	Time to Peak (4)	Time to Unity (5)
		ADVANCED COUNTRIES		
Belgium	7.00	∞	3.00	∞
Sweden	5.83	∞	0.33	∞
Portugal	4.92	8.42	0.92	3.17
Austria	3.92	8.25	0.83	3.25
Italy	3.67	8.17	1.00	3.17
Spain	3.58	8.17	0.75	3.25
Australia	3.67	8.00	0.08	0.67
Canada	5.42	8.00	1.25	4.50
Japan	5.00	8.00	1.00	4.17
United States	4.25	7.83	1.00	4.58
Ireland	2.75	7.50	0.25	0.83
Finland	3.75	5.50	1.33	3.50
Germany	3.25	5.50	0.83	2.50
Netherlands	3.42	5.50	0.83	2.58
United Kingdom	3.25	5.42	0.33	2.25
Korea	2.08	3.17	0.17	0.83
Switzerland	1.75	2.67	0.08	0.92
Norway	1.83	2.50	0.08	0.33
New Zealand	1.75	1.83	0.08	0.92
Iceland	1.00	1.08	0.08	0.08
Advanced countries' median	3.63	7.67	0.79	
		AFRICA		
Burkina Faso	5.58	∞	0.08	NCU
Chad	10.92	∞	17.58	∞
Central African Republic	11.75	∞	0.08	NCU
Gabon	14.17	∞	3.67	∞
Lesotho	∞	∞	1.42	∞
Madagascar	8.33	∞	0.17	NCU
Mauritania	∞	∞	0.33	NCU
Mauritius	4.17	∞	0.17	NS
Morocco	9.67	∞	0.17	NS
Senegal	8.17	∞	0.50	NCU
Togo	6.42	∞	0.08	NCU
Tunisia	10.67	∞	2.08	∞
Niger	∞	∞	0.08	NCU
Gambia, The	1.17	∞	0.25	NCU

Table 1. Half-Lives of Reversion to Parity (years), Real Effective Exchange Rates, March 1973–March 2002, Advanced and Developing Countries (Continued)

Country (1)	Biased Half-life (2)	Unbiased Half-life (3)	Time to Peak (4)	Time to Unity (5)
Ethiopia	4.83	∞	0.83	∞
Tanzania	6.33	14.17	1.17	8.42
Cameroon	3.08	7.25	0.25	S
Seychelles	2.67	6.83	0.17	0.25
Nigeria	4.25	5.58	1.50	3.83
Congo, Republic of	1.42	4.83	0.08	S
Kenya	2.00	4.33	0.08	0.25
Sierra Leone	1.00	4.33	0.08	0.25
Côte d' Ivoire	2.17	3.25	0.25	S
Malawi	2.25	3.17	0.08	0.75
Ghana	1.75	2.75	0.08	0.17
Zambia	0.67	1.50	0.08	0.08
Uganda	1.00	1.25	0.08	S
Rwanda	0.50	1.17	0.08	0.17
Sudan	0.92	1.00	0.33	0.42
Democratic Republic of Congo	0.42	0.42	0.08	S
African median	4.21	22.08	0.17	
		WESTERN HEMISPHERE		
Colombia	8.00	∞	3.25	∞
Dominican Republic	5.25	∞	0.58	NS
Ecuador	6.33	∞	0.92	∞
El Salvador	∞	∞	0.75	∞
Guatemala	5.42	∞	0.50	∞
Haiti	19.33	∞	0.08	NS
Honduras	5.17	∞	0.67	∞
Paraguay	4.50	∞	0.17	NS
Venezuela	6.25	∞	0.17	∞
Panama	10.08	∞	0.08	NCU
Peru	1.00	∞	0.17	0.33
Guyana	12.50	∞	∞	∞
Jamaica	4.42	8.33	1.25	3.50
Trinidad and Tobago	3.92	7.75	1.00	4.17
Uruguay	3.25	7.25	0.08	0.33
Costa Rica	3.08	3.83	0.83	1.33
Argentina	2.75	3.58	0.75	2.08
Barbados	2.58	3.50	0.17	0.92
Chile	2.50	2.75	0.92	1.17

Table 1. Half-Lives of Reversion to Parity (years), Real Effective Exchange Rates, March 1973–March 2002, Advanced and Developing Countries (Concluded)

Country (1)	Biased Half-life (2)	Unbiased Half-life (3)	Time to Peak (4)	Time to Unity (5)
Brazil	1.75	2.50	0.08	0.42
Mexico	1.75	1.92	0.50	0.50
Bolivia	0.08	0.08	0.08	S
Western Hemisphere median	4.46	∞	0.54	
		ASIA, MIDDLE EAST, and EUROPE		
Fiji	4.75	∞	0.08	NCU
Hungary	∞	∞	1.83	∞
India	17.67	∞	16.75	∞
Indonesia	11.92	∞	0.75	NS
Malaysia	8.25	∞	0.33	∞
Malta	7.08	∞	20.00	∞
Myanmar	∞	∞	∞	∞
Pakistan	∞	∞	2.00	∞
Papua New Guinea	1.83	∞	0.08	NCU
Thailand	6.17	∞	0.17	NS
Nepal	9.33	∞	0.17	NCU
Philippines	3.50	8.08	0.42	2.17
Syria	4.42	7.92	0.75	3.25
Samoa	3.33	7.83	0.92	S
Turkey	1.92	5.67	0.08	0.17
Iran, Islamic Republic of	3.08	5.42	0.67	0.92
Sri Lanka	2.50	5.17	0.17	1.50
Egypt	2.00	2.42	0.42	0.25
Asia, Middle East and Europe median	5.46	∞	0.54	
All countries: Median	4.04	8.17	0.33	

Notes: Column (1): Country name. Column (2): Biased half-life of parity deviation, based on least squares estimation of the Augmented Dickey-Fuller regression of equation (1). The half-life for AR(p) models is calculated from the impulse response functions (equation (3)), and is defined as the time taken for as unit impulse to dissipate permanently by one-half from the occurrence of the initial shock. Column (3): Median-unbiased half-life of parity deviation, based on median-unbiased estimation of the Augmented Dickey-Fuller regression of equation (1), as given by Andrews and Chen (1994). The half-life is as described for column (2) above. Column (4): "Time to peak" is the number of months following the unit shock to the real exchange rate that the impulse response function (IRF) reaches its peak. Column (5): "Time to unity" is the number of months following the unit shock to the real exchange rate that the impulse response function (IRF) crosses unity. NCU denotes nonstationary series—the IRF never crosses unity. NS denotes nonstationary series—the IRF is above unity, then crosses unity but does not decay. The symbol ∞ denotes classic nonstationary series—the IRF rises above unity and stays there. S denotes stationary series—the IRF never rises above unity and has a monotonic decay. In calculating group and all-country medians, infinity (∞) is set to equal 30 years.

Source: Authors' calculations.

Table 2. Descriptive Statistics of Persistence of PPP Deviations, Country Groups

Country group	Average Persistence of Deviations from PPP (years)	Number of observations per country group	H_0: Equality of Median Deviation from PPP (*p*-value) WMW test	H_0: Equality of Variance of Deviation from PPP (*p*-value) BF test
(1)	(2)	(3)	(4)	(5)
All countries	8.17	90		
Advanced	7.67	20		
Developing	30.00	70	2.18 (0.029)	6.11 (0.015)
IMF (de jure) Classification:				
Peg	30.00	49		
Flexible	7.25	35	2.73 (0.006)	0.68 (0.413)
Reinhart-Rogoff (de facto) Classification:				
Peg	30.00	22		
Free floating/falling	5.58	12	3.22 (0.001)	0.67 (0.418)

Notes: Column (1): Country group; for definition and derivation see Appendix II. Column (2): Persistence of deviations from PPP is calculated as the group average (median) half-life (in years) of deviations of the real exchange rate from parity. Column (3): Number of observations is the number of countries in each country group. Column (4): WMW is the Wilcoxon-Mann-Whitney test statistic of the null hypothesis of equality of the median half-life of deviations from PPP for each country group; the *p*-value is for the asymptotic normal approximation to the Wilcoxon *t*-statistic (see Conover, 1999). Column (5): BF is the Brown-Forsythe test statistic of the null hypothesis of equality of the variance of the median half-life of deviations from PPP for each country group. The *F*-statistic for the BF test has an approximate *F*-distribution with $G = 1$ numerator degrees of freedom and $N - G$ denominator degrees of freedom, under the null of equal variances in each group, where G is the number of groups and N the number of observations; the approximate *p*-value is given in parentheses.

Table 3. Rank Correlation of Half-life of Deviation from Purchasing Power Parity with Country Characteristics, All Countries

Country characteristic	Rank correlation of characteristic with half-life of parity deviations	Approximate *p*-value (number of observations)
(1)	(2)	(3)
Rate of inflation	-0.281*	0.007 (90)
Volatility of the official nominal exchange rate	-0.379*	0.002 (70)
Volatility of the parallel-market nominal exchange rate	-0.282*	0.02 (70)
Productivity growth	-0.038	0.77 (90)
Government spending	-0.073	0.50 (88)
Trade openness	0.180	0.10 (89)

Notes: Column (1): Country characteristic; for definition and derivation see Appendix I. Column (2): Spearman rank correlation coefficient—the null hypothesis of the Spearman test is that there is no correlation between the (tie-adjusted) rank of each country's bias-corrected half-life of parity deviation and the rank of its period-average value of the country characteristic. Persistence of deviations from PPP is calculated as the median-unbiased half-life (in years) of real exchange rate deviations from parity (listed in column (3) of Table 1). Column (3): The approximate *p*-value is taken from Zar (1972); an asterisk (*) denotes that the null hypothesis of no rank correlation is rejected at the 5 percent level of significance. The number of country observations is given in parentheses.

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