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Economic growth and stationarity of real exchange rates: Evidence from some fast-growing Asian countries

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Abstract

The possible implications of economic growth for the trending behavior of real exchange rates are explored. Given that prior studies consider mainly industrial countries, an issue concerns whether parity reversion holds for countries with sharply different growth experiences. For countries undergoing dramatic income growth from a low level, substantial changes in the relative price structure between tradables and nontradables can often occur. An implication is that real exchange rates for these countries are likely afflicted by trend shifts, which should bear on empirical testing for unit-root nonstationarity. This study analyzes the trend stationarity property of dollar-based real exchange rates for several fast-growing Asian countries. Considerable evidence is found to support the trend-shift hypothesis and reject the unit-root hypothesis. © 1998 Elsevier Science B.V. All rights reserved.

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1. Introduction

The theory of purchasing power parity (PPP) suggests that the equilibrium exchange rate between two currencies equals the ratio of the countries' price

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levels. The empirical relevance of PPP has been examined extensively, witness the voluminous work reviewed by Breuer (1994), Froot and Rogoff (1995) and Rogoff (1996). Empirical findings reported in recent studies appear generally supportive of PPP reversion (e.g., Abuaf and Jorion, 1990; Cheung and Lai, 1993, Cheung and Lai, 1994; Culver and Papell, 1995; Diebold et al., 1991; Lothian and Taylor, 1996; Glen, 1992), although earlier studies often find less favorable results. Even with the recent evidence, nonetheless, the actual extent to which parity reversion prevails across different types of countries is still the subject of a contested issue.

Froot and Rogoff (1995) raise the question of whether sample-selection bias exists in the foregoing studies of PPP. These PPP studies consider primarily industrial countries; consequently, the exchange rates examined in the literature are mainly between pairs of countries that have shared similar growth experiences and enjoyed continually high incomes relative to the rest of the world. This leads to, as Rogoff (1996) has explicitly noted, the issue of whether PPP will hold between two countries with substantially different growth experiences. To date, not much work has been devoted to investigating countries whose incomes have undergone sharp changes. For such countries, the relative price of traded goods to nontraded goods can change drastically, inducing significant shifts in the real-exchange-rate process; it is this type of countries for which parity reversion may most likely fail to work. In an illustrative example, Froot and Rogoff (1995) analyze the real rates of the Argentine peso against both the dollar and the British pound and find that the hypothesis of a unit root cannot be rejected, even for long-horizon data. These authors caution, therefore, that those existing results from PPP studies of industrial countries may overstate the actual extent of empirical support for parity reversion across countries in general.

To explore and contribute evidence pertinent to the issue, this study examines the dynamic behavior of real exchange rates based on the recent experience of several Asian countries since the early 1970s. These countries are particularly interesting to study because they have commonly displayed income growth at a spectacular pace from a low level. The surge in income growth can generate considerable changes in the relative price structure between traded goods and nontraded goods in individual countries. For example, differences in the relative productivity growth between tradables and nontradables can alter the relative price structure and lead to structural shifts in the real-exchange-rate process, a result often labeled as the Balassa–Samuelson effect (Balassa, 1964; Samuelson, 1964). It follows that tests of stationarity in real exchange rates for these countries should properly account for possible structural breaks, which are known to bias unit-root tests toward finding nonstationarity too often (Perron, 1989; Rappoport and Reichlin, 1989).

This study illustrates the possibility of a trend shift in explaining the behavior of real exchange rates for several fast-growing Asian countries. In view of the potential instability in the PPP relationship associated with rapid income growth, unit-root tests that permit a trend break naturally lend themselves to analyzing the long-run behavior of real exchange rates for these countries. Empirical results show that real exchange rates may seemingly contain a unit root when in fact they do not and trend shifts may be responsible.

The allowance for structural breaks in testing for PPP is not entirely new. Studies by Culver and Papell (1995) and Perron and Vogelsang (1992) have reported that some historical series of real exchange rates can be captured by trend- or level-shift models with no unit root, but the relevance of these models in applying to other PPP analyses remains an open empirical question. The present study does not deal with long historical data. Indeed, it focuses the analysis on an interesting but relatively short sample period. Nevertheless, considerable evidence for a trend shift and against a unit root can still be uncovered. The evidence supports that trend shifts are likely to afflict the behavior of real exchange rates for rapidly growing countries.

2. Purchasing power parity

Central to all debates concerning PPP reversion is the basic question: Are real exchange rates governed by permanent disturbances? The PPP theory suggests the presence of a long-run equilibrium relationship between national price levels of two countries when expressed in a common currency unit. 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 the logarithms of the domestic and foreign price indexes, respectively; 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. The real exchange rate is measured by $q_t \equiv p_t - p_t^* - s_t$, which equals u_t. For PPP to hold in the long run, the real exchange rate should be stationary and not driven by permanent disturbances.

3. A structural model of trending real rates

A possible determinant of the behavior of the real exchange rate is productivity growth. According to the Balassa–Samuelson effect, productivity growth can lead to time trends in real exchange rates. Froot and Rogoff (1995) and Obstfeld (1993), for example, present a structural model, illustrating such an effect. Consider an open economy that uses capital and labor to produce tradables priced in the world markets and nontradables priced at home. The production functions for tradables and nontradables are

$$Y_\mathrm{T} = A_\mathrm{T} L_\mathrm{T}^{\alpha} K_\mathrm{T}^{1-\alpha}, \tag{2}$$

$$Y_\mathrm{N} = A_\mathrm{N} L_\mathrm{N}^{\beta} K_\mathrm{N}^{1-\beta}, \tag{3}$$

where Y_T and Y_N are the domestic output of tradables and nontradables, respectively; A_T and A_N are the corresponding productivity parameters for the two sectors; L_T and L_N are labor inputs and K_T and K_N are the capital inputs in the relevant sectors. While all factor inputs are mobile across sectors, only capital is considered internationally mobile. Assuming perfect competition, profit maximization implies that

$$r = (1-\alpha) A_T (K_T/L_T)^{-\alpha}, \tag{4}$$

$$w = \alpha A_T (K_T/L_T)^{1-\alpha}, \tag{5}$$

$$r = (1-\beta) Q A_N (K_N/L_N)^{-\beta}, \tag{6}$$

$$w = \beta Q A_N (K_N/L_N)^{1-\beta}, \tag{7}$$

where r is the rental rate on capital, w is the wage rate (measured in tradables) and Q is the price of nontradables in terms of tradables. As observed by Froot and Rogoff (1995) and Obstfeld (1993), Q can be identified with the real exchange rate: a rise in Q implies a real appreciation, and a fall in Q implies a real depreciation.

When Eqs. (4) and (5) are combined, they yield a wage equation in terms of the factor productivity in tradables:

$$w = \alpha [(1-\alpha)]/r]^{(1-\alpha)/\alpha} A_T^{1/\alpha}, \tag{8}$$

where r is given and determined in world markets. Substituting this into Eq. (7) and combining it with Eq. (6) will then determine both Q and K_N/L_N. In particular, the equilibrium dynamics of Q can be shown to have the following form:

$$\hat{q} = (\beta/\alpha)\hat{a}_T - \hat{a}_N, \tag{9}$$

where $\hat{q} = \mathrm{dlog}\ Q$, $\hat{a}_T = \mathrm{dlog}\ A_T$ and $\hat{a}_N = \mathrm{dlog}\ A_N$. Hence, differential productivity growth between sectors will lead to trending dynamics for the relative price of nontradables. Note that since the capital intensity in production normally differs across sectors (i.e., $\alpha \neq \beta$), the model implies a time trend in the real exchange rate even when $\hat{a}_T = \hat{a}_N$. It follows that whether productivity growth is balanced or not, it can cause time trends in real exchange rates.

According to Froot and Rogoff (1995), the above analysis should be particularly relevant to fast-growth countries, which have typically experienced significant changes in the relative price of nontradables. This study explores the time-trend implication for the real exchange rate. To the extent that productivity growth is often not steady and substantial changes in productivity growth can occur over time, the above model suggests that real exchange rates for fast-growing countries are prone to trend shifts. Interestingly, the results reported later in this paper do confirm the presence of trend shifts in real exchange rates in most cases.

4. Data and some preliminary data analysis

The dynamic behavior of real exchange rates for five fast-growing Asian countries – Hong Kong, Indonesia, Korea, Malaysia, and Singapore – against the United States (a major trading partner of these countries) is investigated. The data under study are monthly series taken from the International Monetary Fund's *International Financial Statistics* data CD-ROM. The data series include monthly averages of spot exchange rates as well as national price levels measured by consumer price indexes. All data series cover the sample period from April 1973 through December 1995.

All series of real exchange rates are first tested for a unit root using the standard augmented Dickey–Fuller (ADF) test. The ADF test results can serve as a benchmark for comparison with other results. For a time series, $\{y_t\}$, the ADF test involves estimating the following regression:

$$(1-L)y_t = \mu_0 + \mu_1 t + \beta_0 y_{t-1} + \sum_{j=1}^{p} \beta_j (1-L) y_{t-j} + \epsilon_t, \tag{10}$$

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

In testing for unit-root nonstationarity in real exchange rates, the power of the statistical test used is of critical importance. Deviations from PPP can be slow to reverse, and traditional unit-root tests are known to have low power to identify stationary but persistent dynamics. As a result, empirical failure to uncover parity reversion can simply reflect the low power of the unit-root test employed. To minimize the problem, a modified Dickey–Fuller test with good power properties is applied in this study, in addition to the ADF test.

Elliott et al. (1996) establish the asymptotic power envelope for unit-root tests by analyzing 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 autoregressive root and $\bar{c}$ is some negative constant. Based on asymptotic power calculation, Elliott et al. (1996) show that a modified Dickey–Fuller test, called the DF-GLS test, can achieve significant power gains over traditional unit-root tests. The superior performance of the DF-GLS test is also supported by the Monte Carlo evidence presented by Stock (1994). This test has been shown by Cheung and Lai (1994), Cheung and Lai (1997) to yield more favorable evidence of PPP reversion among industrial countries than the ADF test.

The DF-GLS test is conducted based on the following regression:

$$(1-L)y_t^{\tau} = \phi_0 y_{t-1}^{\tau} + \sum_{j=1}^{p} \phi_j (1-L) y_{t-j}^{\tau} + e_t, \tag{11}$$

Table 1
Results of the ADF and DF-GLS unit-root tests

Country	p	ADF statistic	p	DF-GLS statistic
Hong Kong	1	-0.296	1	-0.437
Indonesia	1	-3.095	1	-1.243
Korea	1	-1.490	1	-1.381
Malaysia	1	-2.008	1	-1.746
Singapore	1	-1.121	1	-1.388

Both ADF and DF-GLS tests for a unit root are performed on real exchange rates for the period from April 1973 through December 1995. The lag parameter p for either test is selected using a data-dependent method based on the Schwarz information criterion, with the maximum lag order $= 8$. Finite-sample critical values for the ADF test are based on Cheung and Lai (1995a); the 10 and 5% critical values are given by -3.129 and -3.418, respectively, for $p = 1$. Finite-sample critical values for the DF-GLS test are provided by Cheung and Lai (1995b); the 10 and 5% critical values are given by -2.618 and -2.906, respectively, for $p = 1$. All the test statistics reported in this table are not significant at the 10% level.

where e_t is the random error term; and y_t^{τ}, the locally detrended data process under the local alternative of $\bar{\rho} = 1 + \bar{c}/T$, is given by

$$y_t^{\tau} = y_t - z_t \bar{\beta}, \tag{12}$$

with $z_t = (1, t)$ and $\bar{\beta}$ being the regression coefficient of $\tilde{y}_t$ on $\tilde{z}_t$, for which $\tilde{y}_t = (y_1,(1 - \bar{\rho}L)y_2, \dots ,(1 - \bar{\rho}L)y_T)'$ and $\tilde{z}_t = (z_1,(1 - \bar{\rho}L)z_2, \dots , (1 - \bar{\rho}L)z_T)'$. The DF-GLS test statistic is given by the t-statistic testing $\phi_0 = 0$ against the alternative of $\phi_0 < 0$ in regression (11). Elliott et al. (1996) recommend that the parameter $\bar{c}$, which defines the local alternative through $\bar{\rho} = 1 + \bar{c}/T$, be set equal to -13.5.

The results of both the ADF test and the DF-GLS test are summarized in Table 1. The lag parameter p for either test is chosen using an information-based method (e.g., Hall, 1994). The test results are clearly not in favor of parity reversion. In no case can the null hypothesis of a unit root be rejected at any standard levels of significance for the ADF test. The DF-GLS test, though with better power than the ADF test, also fails to reveal any significant evidence of stationarity in real exchange rates. Without further analysis, these results would lead researchers to conclude that deviations from PPP for fast-growing countries are largely governed by permanent shocks.

5. Unit-root testing under the possibility of a trend break

The two unit-root tests considered earlier are all constructed under the maintained hypothesis of a single, linear deterministic trend. When a trend break occurs, the estimated models under the tests suffer misspecification, and neither test is consistent. In this situation, the unit-root tests will have low power against stationary/broken-trend alternatives. The broken-trend issue is particularly relevant to the present analysis, given that trend shifts in real exchange rates may likely occur in association with rapid income growth. As illustrated by the model discussed in Section 3, productivity growth can alter the trending behavior of real exchange rates.

The income growth rates are computed for the Asian countries under study for the period 1973–1995. Table 2 gives the average annual rates of growth in both real GDP and real per capita GDP for individual countries, along with their ratios to the corresponding growth rates for the United States (denoted by g_i/g_{US}) over the same sample period. Except for Hong Kong, the data on population and real GDP for all the other countries are from the *International Financial Statistics* data CD-ROM. For Hong Kong, the data are obtained from the Census and Statistics Department of Hong Kong, with some data being obtained from various issues of *Hong Kong Monthly Digest of Statistics* published by the Department. As shown in the table, all five Asian countries displayed high income growth at different average rates between 6.5 and 8.3% for real GDP and between 4.1 and 6.9% for real per capita GDP. During the same data period, these countries grew at about 2.8–3.5 times as fast as the United States in terms of real GDP and at about 3.0–5.0 times as much in terms of real per capita GDP. These data confirm that all the five Asian countries have had a much faster rate of income growth than the United States since the early 1970s.

To investigate whether the trend-shift hypothesis applies to the dynamics of real exchange rates for the fast-growing Asian countries, a number of sequential unit-root tests are carried out here. These sequential tests extend the ADF test by accounting for a possible jump or shift in trend in the underlying data process, with no prior knowledge of the break date. The treatment of an unknown break

Table 2
Average annual rates of growth in national income

	Real GDP		Real per capita GDP	
Country	g_i (%)	g_i/g_{US}	g_i (%)	g_i/g_{US}
Hong Kong	7.09	3.01	5.27	3.89
Indonesia	6.47	2.75	4.52	3.34
Korea	8.27	3.51	6.87	5.04
Malaysia	6.89	2.91	4.11	3.02
Singapore	7.58	3.22	6.07	4.48

Average income growth rates are computed for five Asian countries for the period 1973–1995. The column beneath 'g_i' gives the average growth rates in real GDP or real per capita GDP for individual countries (i = Hong Kong, Indonesia, Korea, Malaysia, and Singapore) over our sample period. The column beneath 'g_i/g_{US}' gives the corresponding ratios of the average growth rates for individual countries to that for the United States.

date is desirable since the timing of the break, if any, can vary across series of real exchange rates and since any arbitrarily fixed date can be subject to criticism of data mining. In fact, no theory seems able to offer information about exactly when a trend break will occur. Note that without perfect knowledge about the exact break date, no one-step, simultaneous testing method is available. This problem is handled, however, by estimating the likely break date directly from the data using some sequential testing procedures. The sequential tests applied in this study have been shown to perform well in identifying the true break date in simulation analysis.

Two basic approaches for modeling structural breaks in time series have been considered in the literature (Banerjee et al., 1992; Perron and Vogelsang, 1992; Vogelsang and Perron, 1994; Zivot and Andrews, 1992). One is the additive outlier (AO) approach that views the break as happening instantly, and the other is the innovational outlier (IO) approach that allows the change to take place gradually over time.

5.1. The additive outlier approach

There are several different AO models capturing different types of process shifts in either the mean or time trend or both. These models are represented by

$$y_t = c_0 + c_1 t + \theta d_t(k) + z_t, \tag{13a}$$

$$y_t = c_0 + c_1 t + \eta d_t^*(k) + z_t, \tag{13b}$$

$$y_t = c_0 + c_1 t + \theta d_t(k) + \eta d_t^*(k) + z_t, \tag{13c}$$

where $d_t(k) = I(t > k)$ and $d_t^*(k) = (t-k)I(t>k)$, with $I(\cdot)$ being the usual indicator function. The innovation process z_t is defined by $(1-\rho L)A(L) = B(L)\upsilon_t$, where $A(L)$ and $B(L)$ are polynomials in L with stable roots, and υ_t is a white noise. When $\rho = 1$, y_t has a unit root. Specification (13a) permits a level shift to occur at time k, which implies a shift in the process mean. Perron (1989) refers to this as the 'crash' model. In contrast, a trend shift (i.e., a change in the slope of the trend function) is allowed for at time k under specification (13b), as in Perron's 'changing growth' model. Specification (13c) admits shifts in both the mean and trend.

Testing for a unit root in the AO models involves a sequential two-step procedure. Detrended series are first obtained as follows:

$$y_t = \mu_0 + \mu_1 t + \alpha d_t(k) + \tilde{y}_t^1, \tag{14a}$$

$$y_t = \mu_0 + \mu_1 t + \gamma d_t^*(k) + \tilde{y}_t^2, \tag{14b}$$

$$y_t = \mu_0 + \mu_1 t + \alpha d_t(k) + \gamma d_t^*(k) + \tilde{y}_t^3, \tag{14c}$$

where the residual series, $\tilde{y}^1$, $\tilde{y}^2$ and $\tilde{y}^3$, gives the detrended series of y_t. Next, tests for $\beta_0 = 0$ under the null hypothesis of a unit root are performed using the following regressions:

$$(1-L)\tilde{y}_t^1 = \sum_{j=0}^{p} \omega_j D_{t-j}(k) + \beta_0 \tilde{y}_{t-1}^1 + \sum_{j=1}^{p} \beta_j (1-L)\tilde{y}_{t-j}^1 + \xi_t, \tag{15a}$$

$$(1-L)\tilde{y}_t^2 = \beta_0 \tilde{y}_{t-1}^2 \sum_{j=1}^{p} \beta_j (1-L)\tilde{y}_{t-j}^2 + \xi_t, \tag{15b}$$

$$(1-L)\tilde{y}_t^3 = \sum_{j=0}^{p} \omega_j D_{t-j}(k) + \beta_0 \tilde{y}_{t-1}^3 + \sum_{j=1}^{p} \beta_j (1-L)\tilde{y}_{t-j}^3 + \xi_t, \tag{15c}$$

where ξ_t is the error term. The dummy variables, $D_{t-j}(k) = I(t = k + j + 1)$ for $j = 0, \ldots, p$, are included in cases involving a mean shift to ensure the test robustness with respect to the error correlation structure (see (Vogelsang and Perron, 1994) for further discussion).

In implementing the above two-step procedure, the breakpoint k is assumed to be unknown and needs to be estimated from the data. By varying k for each regression over the sample period, specifically, k will be chosen to maximize over a sequence of F-statistics testing the statistical significance of the break parameters: $\alpha = 0$ in regression (14a); $\gamma = 0$ in regression (14b); and $\alpha = 0 = \gamma$ in regression (14c). This method has been shown to perform well in picking the true breakpoint. The t-statistics for testing $\beta_0 = 0$ computed at where $F(\tilde{k}/T) = \max_{r \leqslant k \leqslant T-r} F(k/T)$ are denoted by τ_{DF} (AO, $\tilde{k}$, i), $i = a$, b and c, for regressions (15a) to (15c). The trimming parameter r is set equal to the integer part of $0.15T$, following Banerjee et al. (1992).

5.2. The innovational outlier approach

In contrast to the AO approach, the IO approach entertains situations in which the break can occur not abruptly but slowly over time. In allowing for gradual structural changes, this approach provides more flexibility in modeling trend breaks than the AO approach. In fact, the IO approach appears intuitively attractive and should be particularly relevant for our present analysis since it is reasonable to expect that changes in economic growth occur gradually rather than instantly.

Various possible IO model specifications are

$$y_t = c_0 + c_1 t + \psi(L)\big(\theta d_t(k) + \nu_t\big), \tag{16a}$$

$$y_t = c_0 + c_1 t + \psi(L)\big(\eta d_t^*(k) + \nu_t\big), \tag{16b}$$

$$y_t = c_0 + c_1 t + \psi(L)\big(\theta d_t(k) + \eta d_t^*(k) + \nu_t\big), \tag{16c}$$

where $\psi(L)$ is some lag polynomial. Notice that through the function $\psi(L)$ in lags, a break can operate and impact the process gradually over a period of time. Parallel to the AO models, (16a) represents the mean-shift model specification, whereas (16b) is the trend-shift one. Specification (16c) gives a combination of the two models.

Corresponding to models (16a), (16b) and (16c), the following regressions for unit-root tests are estimated:

$$(1-L)y_t = \mu_0 + \mu_1 t + \omega D_t(k) + \alpha d_t(k) + \beta_0 y_{t-1} + \sum_{j=1}^{p} \beta_j (1-L) y_{t-j} + \zeta_t, \tag{17a}$$

$$(1-L)y_t = \mu_0 + \mu_1 t + \gamma d_t * (k) + \beta_0 y_{t-1} + \sum_{j=1}^{p} \beta_j (1-L) y_{t-j} + \zeta_t, \tag{17b}$$

$$(1-L)y_t = \mu_0 + \mu_1 t + \omega D_t(k) + \alpha d_t(k) + \gamma d_t^*(k) + \beta_0 y_{t-1} + \sum_{j=1}^{p} \beta_j (1-L) y_{t-j} + \zeta_t, \tag{17c}$$

where ζ_t is the error term. The regression models considered here follow Vogelsang and Perron (1994). Regression (17b) has been used by Banerjee et al. (1992) and Zivot and Andrews (1992). These two studies also consider regressions similar to (17a) or (17c), though without the one-time dummy variable $D_t(k)$. As for the AO models, the breakpoint k will be selected using the maxima of the F-statistics testing for $\alpha = 0$ in regression (17a), $\gamma = 0$ in regression (17b), and $\alpha = 0 = \gamma$ in regression (17c). In each regression, the t-statistic for testing $\beta_0 = 0$ is computed at the chosen breakpoint, and the resulted statistic is denoted by τ_{DF} (IO, $\tilde{k}$, i) with $i = a$, b or c.

5.3. Empirical results

Each series of real exchange rates is tested for a unit root using various sequential tests under both AO and IO settings, without restricting a priori the analysis to any specific type of structural breaks (instant or gradual; mean or trend shift). In performing each of these tests, the lag parameter p is chosen using a data-dependent method based on the Schwarz information criterion, as applied before to the ADF and DF-GLS tests.

Table 3 exhibits the results of sequential unit-root tests for the different AO models. In all cases, $p = 1$ is selected and used. The mean-shift model does not explain the data on real exchange rates well. In no case can the null hypothesis of

Table 3
Unit-root tests under additive outlier models of a broken trend

Country	The mean-shift model		The trend-shift model		The combined model	
	p	τ_{DF} (AO, $\tilde{k}$, a)	p	τ_{DF} (AO, $\tilde{k}$, b)	p	τ_{DF} (AO, $\tilde{k}$, c)
Hong Kong	1	−1.246	1	−4.261[a]	1	−3.020
Indonesia	1	−3.863	1	−1.999	1	−1.516
Korea	1	−2.464	1	−2.283	1	−2.492
Malaysia	1	−3.747	1	−4.531[b]	1	−4.597[a]
Singapore	1	−2.840	1	−5.923[b]	1	−5.934[b]

Statistical significance is indicated by [a] for the 10% level and by [b] for the 5% level.
This table reports unit-root test results under different additive outlier models of a trend break. The data on real exchange rates for five Asian countries over the period from April 1973 through December 1995 are examined. The lag parameter p for each test is selected using a data-dependent method based on the Schwarz information criterion, with the maximum lag order = 8. The unit-root test statistics given by τ_{DF} (AO, $\tilde{k}$, i) for $i = a$, b and c are obtained based on regressions (15a), (15b) and (15c), respectively. Asymptotic critical values for individual statistics are computed by Vogelsang and Perron (1994). For the mean-shift model, the critical values are given by −3.90 (the 10% significance level) and −4.17 (the 5% level). For the trend-shift model, the critical values are given by −4.04 (the 10% level) and −4.34 (the 5% level). For the combined model, the critical values are given by −4.31 (the 10% level) and −4.61 (the 5% level).

a unit root be rejected when a level shift is included under the alternative hypothesis. When the changing growth model permitting a trend shift is fitted, however, some evidence for stationarity in real exchange rates begins to unveil.

Table 4
Unit-root tests under innovational outlier models of a broken trend

Country	The mean-shift model		The trend-shift model		The combined model	
	p	τ_{DF} (IO, $\tilde{k}$, a)	p	τ_{DF} (IO, $\tilde{k}$, b)	p	τ_{DF} (IO, $\tilde{k}$, c)
Hong Kong	1	−2.113	1	−4.655[b]	1	−5.205[b]
Indonesia	1	−4.792[b]	1	−4.182[a]	1	−4.880[a]
Korea	1	−3.070	1	−2.442	1	−3.138
Malaysia	1	−4.353	1	−4.679[b]	1	−4.798
Singapore	1	−3.071	1	−6.059[b]	1	−5.994[b]

Statistical significance is indicated by [a] for the 10% level and by [b] for the 5% level.
This table displays unit-root test results under different innovational outlier models of a trend break. The data on real exchange rates for five Asian countries over the period from April 1973 through December 1995 are analyzed. The lag parameter p for each test is selected using a data-dependent method based on the Schwarz information criterion, with the maximum lag order = 8. The unit-root test statistics given by τ_{DF} (IO, $\tilde{k}$, i) for $i = a$, b and c are obtained from regressions (17a), (17b) and (17c), respectively. Critical values for the crash model and the changing growth models are both based on Banerjee et al. (1992); those for the mean-shift model are given by −4.50 (the 10% significance level) and −4.79 (the 5% level) for $T = 250$, and those for the trend-shift model are given by −4.12 (the 10% level) and −4.39 (the 5% level) for $T = 250$. For the combined model, asymptotic critical values are available from Vogelsang and Perron (1994); they are given by −4.86 (the 10% level) and −5.16 (the 5% level).

The unit-root hypothesis can be rejected in favor of stationary/trend-shift alternatives in two cases (Malaysia and Singapore) at the 5% level and one more case (Hong Kong) at the 10% level. The combined mean- and trend-shift model also yields evidence against a unit root in the cases of Malaysia and Singapore. The results presented next show that stronger and broader evidence rejecting a unit root can be further obtained when the IO models, which permit trend breaks to occur gradually, are fitted to the data.

Table 4 contains the results of sequential unit-root tests for the different IO models. Again, the mean-shift model performs unsatisfactory in capturing the data dynamics in general, producing little significant evidence against a unit root, except in the case of Indonesia only. For Indonesia, indeed, a mean shift and a trend shift may both be applicable. On the other hand, a trend shift appears more relevant to the other countries than a mean shift. When a trend shift is admitted under the alternative hypothesis, we can reject a unit root in real exchange rates in four out of the five cases under examination. The results from the hybrid model also mimic largely those from the trend-shift model. Combining all these results together, there is considerable evidence supporting that real exchange rates can be characterized as being stationary around a broken trend.

6. Further discussion of the results

The behavior of real exchange rates has been widely examined under both fixed-rate and flexible-rate systems in prior studies. Most of the countries considered in this study have had their currencies either pegged to the US dollar or a composite currency basket. It is known that the exchange rate arrangement may affect the behavior of real exchange rates. Under a fixed-rate system, the central banks of the countries concerned are expected to intervene in foreign exchange markets to maintain their currency values against any significant rate divergences. Such interventions, if they are successful, will smooth out the exchange rate fluctuations over time and may reduce the likelihood of substantial level shifts. This may be reflected in our findings that the mean-shift (or crash) model is generally not supported by the data (the authors thank an anonymous referee for this observation).

It should be noted that the type of dynamics that is of central interest to this analysis is not the mean-shift model but the trend-shift model. The latter model is consistent with the model of the Balassa–Samuelson effect under the presence of productivity growth changes. The results of the mean-shift model are, nonetheless, presented for the comparison purpose so that the empirical analysis avoids imposing a priori any specific form of break and allows entirely the data to determine the appropriate form. Empirical results confirm, indeed, that the dynamics of real exchange rates are mostly characterized by a trend-shift model.

Interestingly, the results further support that the IO model of a gradual trend shift generally fits the data better than the AO model of an instant, abrupt trend shift. This accords with our intuition that changes in productivity growth are more likely to be a gradual process instead of a one-time, abrupt event.

The gradualness of the trend shifts identified in our data is not incompatible with market efficiency. In efficient markets, asset prices will reflect news rapidly, thus creating possible jumps sporadically in both up and down directions. These jumps certainly contribute to the randomness and variability of the distribution of price changes; however, they differ in nature from the types of systematic structural changes considered in the trend-break models. The gradual shifts in time trends detected in the present analysis cannot be explained and generated by random jumps. Instead, such trend shifts can simply reflect systematic gradual changes in underlying economic fundamentals, like productivity growth.

Finally, it is instructive to investigate if the trend-shift dynamics, which appear common among fast-growing Asian countries, apply to industrial countries as well. Recent studies by, e.g., Liu and He (1991a,b), Pan et al. (1996) and Phylaktis and Kassimatis (1994) do not reveal any significant differences in the real-exchange-rate dynamics between Asian countries and industrial countries. The possibility of trend-shift dynamics has not been considered in these previous studies, however. To check whether the trend-shift models of real exchange rates hold for industrial countries, dollar-based real exchange rates for four major European countries (namely, France, Germany, Italy and the United Kingdom) over the same sample period are examined in the present study. Tables 5 and 6

Table 5
Results of additive outlier tests for European countries

	The mean-shift model		The trend-shift model		The combined model	
Country	p	τ_{DF} (AO, $\tilde{k}$, a)	p	τ_{DF} (AO, $\tilde{k}$, b)	p	τ_{DF} (AO, $\tilde{k}$, c)
France	1	−3.048	1	−2.508	1	−3.122
Germany	1	−3.579	1	−2.908	1	−3.093
Italy	1	−3.152	1	−2.290	1	−2.975
United Kingdom	1	−2.855	1	−2.493	1	−3.177

This table contains unit-root test results under different additive outlier models of a trend break. The data on real exchange rates for four major European countries over the period from April 1973 through December 1995 are examined. The lag parameter p for each test is selected using a data-dependent method based on the Schwarz information criterion, with the maximum lag order = 8. The unit-root test statistics given by τ_{DF} (AO, $\tilde{k}$, i) for $i = a$, b and c are obtained based on regressions (15a), (15b) and (15c), respectively. Asymptotic critical values for individual statistics are computed by Vogelsang and Perron (1994). For the mean-shift model, the critical values are given by −3.90 (the 10% significance level) and −4.17 (the 5% level). For the trend-shift model, the critical values are given by −4.04 (the 10% level) and −4.34 (the 5% level). For the combined model, the critical values are given by −4.31 (the 10% level) and −4.61 (the 5% level). All the test statistics reported in this table are not significant at the 10% level.

Table 6
Results of innovational outlier tests for European countries

	The mean-shift model		The trend-shift model		The combined model	
Country	p	τ_{DF} (IO, $\tilde{k}$, a)	p	τ_{DF} (IO, $\tilde{k}$, b)	p	τ_{DF} (IO, $\tilde{k}$, c)
France	1	−4.014	1	−2.702	1	−4.001
Germany	1	−4.182	1	−3.118	1	−4.094
Italy	1	−4.199	1	−2.389	1	−3.992
United Kingdom	1	−3.446	1	−2.555	1	−3.862

This table provides unit-root test results under different innovational outlier models of a trend break. The data on real exchange rates for four major European countries over the period from April 1973 through December 1995 are studied. The lag parameter p for each test is selected using a data-dependent method based on the Schwarz information criterion, with the maximum lag order = 8. The unit-root test statistics given by τ_{DF} (IO, $\tilde{k}$, i) for $i = a$, b and c are obtained from regressions (17a), (17b) and (17c), respectively. Critical values for the crash model and the changing growth models are both based on Banerjee et al. (1992); those for the mean-shift model are given by −4.50 (the 10% significance level) and −4.79 (the 5% level) for $T = 250$, and those for the trend-shift model are given by −4.12 (the 10% level) and −4.39 (the 5% level) for $T = 250$. For the combined model, asymptotic critical values are available from Vogelsang and Perron (1994); they are given by −4.86 (the 10% level) and −5.16 (the 5% level). All the test statistics reported in this table are not significant at the 10% level.

contain the results of the AO and IO trend-break tests, respectively. The mean-shift tests yield similar non-rejection results as those obtained for the Asian countries. On the other hand, the empirical results from the trend-shift tests contrast sharply with those reported for the Asian countries. In no case can the real-exchange-rate dynamics for the European countries be explained by a trend-shift model, regardless of whether the AO or IO model is considered.

7. Conclusion

The possible implications of economic growth for the trending behavior of real exchange rates have been explored. Prior analyses of parity reversion consider mainly exchange rates between industrial countries, which share similar growth experiences. An issue concerns the time-series behavior of real exchange rates for countries with sharply different growth experiences. Specifically, for countries undergoing dramatic income growth from a low level, substantial changes in the relative price structure between tradables and nontradables may likely occur. An implication is that real exchange rates for this type of countries are susceptible to trend shifts. If this is true, the likely occurrence of trend shifts should bear upon empirical testing for unit-root nonstationarity in real exchange rates.

The trend stationarity in dollar-based real exchange rates of five fast-growing Asian countries has been examined in this paper. Without allowing for any trend break, it is shown that unit-root tests consistently fail to reveal stationarity in real exchange rates for these countries, suggesting that deviations from PPP are driven mainly by permanent disturbances. On the other hand, when unit-root tests that permit a trend shift are applied, considerable evidence in favor of no unit root can be unveiled, rejecting the existence of a large permanent component in the dynamics of PPP deviations. It follows that real exchange rates may appear nonstationary when in fact they are not and trend shifts may be responsible. The results support the empirical relevance of the trend-shift hypothesis for the behavior of real exchange rates.

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