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Purchasing Power Parity and Nonlinear Real Exchange Rate Adjustment: Evidence From High-Growth Countries

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This study applies nonlinear KSS unit root test (Kapetanios, Shinb, and Snell, 2003) and an Asymmetric Exponential Smooth Transition Auto-Regressive (AESTAR) unit root test, proposed by Sollis (2009), to investigate the validity of long-run Purchasing Power Parity (PPP) for six high-growth countries. The empirical results indicate that PPP holds for five of the six high-growth countries studied, namely Brazil, China, Indonesia, Mexico and South Korea, using the KSS test. Furthermore, using Sollis (2009) AESTAR unit root test reveals that real appreciations in the value of the Indonesia Rupiah-U.S. and Korea Won-U.S. dollar exchange rate are slower to revert do the mean (nonlinearly) than depreciations of the same proportionate amount, and the adjustment toward PPP is found to be nonlinear and asymmetric. On the other hand, Brazil/USD, China/USD and Mexico/USD adjustments are found to be nonlinear and symmetric. The governments of these five countries can use PPP to determine whether a currency is overvalued or undervalued, as well as if the country is experiencing differences between domestic and foreign inflation rates. These results have important policy implications for the emerging high-growth economies in this study.

Keywords: purchasing power parity, nonlinear unit root test, asymmetric, AESTAR test, high-growth countries
JEL classification: C22, F31

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1 Introduction

Considerable studies have been put into testing the validity of purchasing power parity (hereafter, PPP) hypothesis, during the past several decades. PPP can be seen as a preliminary measure of exchange rate equilibrium. Moreover, the validity of PPP hypothesis implies prices co-movement and evidence of well-developed trade relations between two countries. Although PPP hypothesis has been thoroughly examined for developed as well as developing countries, the literature is not rich for the high-growth countries. Because of PPP is a cornerstone of many exchange rate determination models, and is one of the most researched theories in the economics profession, as such, testing PPP has continued to attract the attention of many researchers, but with mixed results.

It is not surprising, there are an adequate number of empirical literatures on PPP, both academic and policy related has evolved, but unfortunately, none has been conclusive. There has been a lot of controversy concerning the usefulness of the PPP doctrine as an exchange rate determination model. Some economists even questioned its validity, such as Frenkel (1981), Baillie and Selover (1987), Taylor (1988) and Cushman (2008) who have found no support for long-run PPP.[1] On the other hand, researches by Johansen and Juselius (1992), Rogoff (1996), Taylor, Peel, and Sarno (2001), Sarno and Taylor (2002), Cheung, Lai, and Bergman (2004), Taylor and Taylor (2004), Lothian and Taylor (2008), Arize, Malindretos, and Nam (2010), and Chang, Lee, and Chou (2012) suggested that PPP is valid in the long run. Additionally, Mohsen (1993) and Odedokun (2000) also found mixed results about the existence of PPP in their studied.

The results from validity of PPP have important implications to decision or policy-makers of central banks, multinational firms and exchange rate market participants (Taylor, 1988; Taylor, 2003; Taylor, 2006; Lothian and Taylor, 1996). In particular, PPP is indicative of long-run relationship among nominal exchange rate, domestic and foreign prices. When PPP exists, it can then be used to determine the equilibrium exchange rate; however, when PPP does not hold, it cannot be used to determine the equilibrium exchange rate, thereby, the use of any monetary approach to determine the

[1] Frenkel (1981, p. 145) observed that the aspect of PPP as follows: During the 1970s short-run changes in exchange rates bore little relationship to short-run differentials in national inflation rates and frequently, divergences from PPP have been cumulative.

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exchange rate is invalidated and invalid PPP also disqualifies the monetary approach to exchange rate determination, which requires PPP to hold true. Estimates of PPP exchange rates are important for practical purposes such as determining the degree of misalignment of the nominal exchange rate and the appropriate policy response, the setting of exchange rate parities, and the international comparison of national income levels.

In the international finance literature, the question is widely asked whether the deviation from PPP is persistent or not (Rogoff, 1996; Taylor and Taylor, 2004; Taylor, 2003; Taylor, 2006). Various univariate and multivariate econometrics techniques with improved power are applied to examine if real exchange rates are stationary (Taylor, 1988). The apparently high persistence of real exchange rates is most responsible for these various efforts to better understand their stochastic properties. It is well-known that most unit root tests have little power for highly persistent real exchange rates (Taylor, 1990; Lothian and Taylor, 1997). Also, Caporale, Nikitas, and Panayotis (2003) show that the Dickey-Fuller unit root tests display quite behavior when they are applied to real exchange rate data series. Lothian and Taylor (2008) pointed out that if the equilibrium exchange rate is moving gradually over time but statistical tests for real exchange rate stability assume that the equilibrium exchange rate is constant, then estimates of the speed of reversion towards the mean will be biased, and this bias may be at least partly responsible for Rogoff's PPP puzzle. Evidence suggestive of a bias arising from this source is provided by studies which have found that allowing for linear or nonlinear deterministic trends can make a material difference in resolving the puzzles about whether and how fast the exchange rate moves to its PPP level (Taylor, 2002; Lothian and Taylor, 2000).

For previous studies, one possible explanation for the inconsistencies in the existing empirical evidence on the PPP hypothesis is that the prior studies implicitly assume that exchange rate behavior is intrinsically linear in nature. However, Sarno (2000) and Baharumshah, Liew, and Chowdhury (2010) manifested that the adoption of linear stationarity test is unsuitable for the detection of mean reversion if the true process of the data generation of the exchange rate is in fact a stationary non-linear process. Taylor (2001) indicated that the power of the conventional ADF test is poor if the series follow a non-linear threshold process. To do that, the non-linear unit root test based on an exponential smooth transition auto-regressive (ESTAR) proposed by Kapetanios, Shinb, and Snell (2003) and it shows that the power of their test is higher than that of the ADF test. The omission of some

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structural breaks is a possible cause of the traditional unit root tests failing to reject the unit root null for real exchange rate. Perron (1989) argued that if there is a structural break, the power to reject a unit root decreases when the stationary alternative is true and the structural break is ignored. Perron (2006) indicated that no unit root is often supported by incorporating structural break in the unit root test. As we know that exchange rates might be affected by internal and external shocks generated by structural changes may be subject to great short-run variation. If real exchange rate is found stationary by using unit root test with structural breaks, the effects of shocks such as monetary shocks are only temporary. Narayan (2005) and Narayan (2006) provide evidence that, when structural breaks are included for individual countries, real exchange rate is stationary, implying support for purchasing power parity. Then, PPP is valid in the long run.

The ESTAR model has proved to be popular in economics for the analysis of time series data, such as data on real exchange rates. The presence of trading costs suggests that while large deviations of real exchange rates from their equilibrium values will be corrected by arbitrage, small deviations may not be corrected, and the globally stationary ESTAR model with a unit root central regime is able to capture this type of nonlinearity (Baum, Barkoulas, and Caglayan, 2001; Taylor, Peel, and Sarno, 2001; Sollis, 2009; Chang, Lee, and Chou, 2012). A number of tests of the unit root hypothesis against stationary ESTAR nonlinearity have recently been proposed (Kapetanios, Shinb, and Snell, 2003; Park and Shintani, 2005; Chang, Lee, and Chou, 2012). In some empirical applications the assumption of symmetric mean reversion (linear or nonlinear) is not overly restrictive, however in other empirical applications one might expect asymmetry in the adjustment of the process towards its equilibrium in other empirical applications. For example, in the case of real exchange rates, one might expect asymmetric adjustment if domestic or foreign policymakers behave asymmetrically in response to appreciations and depreciations of the same proportionate amount. Sollis, Leybourne, and Newbold (2002) and Sollis (2009) find evidence suggesting that asymmetric nonlinear mean reversion is an important feature of data on real exchange rates against the U.S. dollar.

The aim of this empirical study is to investigate one of the most controversial theories in international economics –PPP– whether or not long-run holds in emerging market countries, especially, high-growth countries (Brazil, China, India, Indonesia, Korea, and Mexico), and the adjustment process toward its equilibrium is nonlinear in a symmetric or an asym-

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metric way. We test the stationarity of real exchange rate for six high-growth countries using a simple nonlinear unit root test of Sollis (2009), that is Asymmetric Exponential Smooth Transition Auto-Regressive (hereafter, AESTAR) unit root test. The major advantage of this approach is that it allows us to simultaneously investigate nonstationarity and nonlinearity in symmetric or asymmetric. In this study, we find that AESTAR unit root test strongly rejects the unit root hypothesis for the majority of six high-growth countries examined, while the traditional unit root tests such as the ADF, PP, and KPSS, lead to no rejection at all. In addition, the adjustment process toward its equilibrium is nonlinear and asymmetric for Indonesia/USD and Korea/USD real exchange rate.

The period under this study was a tumultuous one in the world economy that includes the Asia financial crisis and global financial crisis, and these events most certainly impacted Asian emerging countries such as China, Korea, India, Indonesia and Latin American countries as Mexico, Brazil. As we know, these financial events affected the economy seriously such that these countries share some characteristics as depreciation of the real exchange rate, high inflation, increase in consumer price index and trade openness which might have led to quicker adjustment in relative prices and contributed for PPP to hold.

The remainder of this paper is organized as follows: section 2 discusses the theoretical model of real exchange rate and the theory of PPP. Section 3 presents the data in our study and preliminary investigation. Section 4 introduces the econometric methodologies employed in this study and reports the empirical results. Finally, section 5 reviews the conclusions we draw.

2 The Foundation in Theoretical Model of Real Exchange rate and PPP

The validity of long-run PPP can be evaluated by testing for stationarity in real exchange rate, which is defined as follows:

$$R_t = \frac{E_t P_t^*}{P_t}, \tag{1}$$

where R_t is the real exchange rate, E_t is the nominal exchange rate in terms of domestic currency per unit of U.S. dollar, P_t and P_t^* are domestic and foreign price levels, respectively. The bilateral real exchange rate is the relative price of foreign good in terms of domestic good. It measures the deviation

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from purchasing power parity. We use Consumer Price Index (CPI) in our study. Taking natural logarithm of both sides of Eq. (1) and rearranging the terms yields

$$r_t = e_t + p_t^* - p_t, \tag{2}$$

where r_t is log-transformed the real exchange rate, e_t is log-transformed the nominal exchange rate, p_t^* and p_t are log-transformed the U.S. and domestic consumer price index, respectively.

One approach to testing for PPP is to test r_t for stationarity. Evidence of a unit root in the real exchange rate indicates nonstationarity, thus PPP does not hold in the long run. If PPP holds, it implies that nominal exchange rate is corrected for inflation differentials (Taylor and Taylor, 2004). If r_t has a unit root, then it implies that deviations from the parity are cumulative and not ultimately self-reverting. Nonstationarity in real exchange rates has many macroeconomic implications. For example, Dornbusch (1987) has argued that if real exchange rate depreciates, it could bring a gain in international competitiveness, which in turn could shift the employment towards the depreciating country. Therefore, it is important to establish the empirical validity of the PPP theory.

3 Data

In this research we analyze the logarithms of monthly observations on real exchange rates of six high-growth countries against the U.S. dollar (USD). These time series constructed from consumer price indices (CPI, based on 2005 = 100) and nominal exchange rates recorded in the Taiwan Economic Journal (TEJ) database. O'Neill, Stupnytska, and Wrisdale (2011) proposed eight high-growth countries: China, India, Brazil, South Korea, Turkey, Mexico, Indonesia and Russia. Due to insufficient data for Russia and Turkey, our empirical analysis only covers six countries. Table 1 presents the time span for each empirical country and monthly data are employed in this study.

A summary of the statistics of bilateral real exchange rate is given in Table 2. The Jarque-Bera test results indicate that for all six high-growth countries, the bilateral real exchange rate data sets are approximately non-normal. Figure 1 plots the real exchange rate series for six high-growth countries. We find significant upward or downward trend in some real exchange rate series. Most of the series seem to exhibit some nonlinear adjustment patterns, espe-

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Table 1: The Study Periods for six High-Growth Countries

Country	Study Periods	Obs.
China	1990M1–2012M7	271
India	1998M1–2012M3	171
Brazil	1998M1–2012M3	171
South Korea	1990M1–2012M7	271
Mexico	1998M1–2012M3	171
Indonesia	1990M1–2012M7	271

Table 2: Summary Statistics: Ln(real Exchange Rate)

Country	Brazil	China	India	Indonesia	Korea	Mexico
Mean	1.093	1.838	3.759	9.028	7.118	2.411
Max.	2.407	2.111	3.945	10.136	7.629	2.637
Min.	0.300	1.087	3.423	8.713	6.893	2.259
Std. Dev.	0.602	0.289	0.140	0.274	0.149	0.072
Skewness	0.485	−1.309	−0.758	1.140	0.514	0.609
Kurtosis	1.977	3.252	2.409	4.384	2.470	3.282
Jarque-Bera	14.157	78.085	18.850	80.315	15.112	11.143
Prob.	0.001	0.000	0.000	0.000	0.001	0.004

Notes: ln (real exchange rate) = ln(nominal exchange rate) + ln (foreign price level) – ln(domestic price level). the base country is U.S.

cially Indonesia and Korea have structure breaks at 1998 that can be clearly seen by the Asia financial crisis and Mexico has structure break at 2008 was associated with the global financial crisis.

4 Methodologies and Empirical Results

Recently, there is a growing consensus that exchange rate might exhibit non-linearities and that conventional test such as the augmented Dickey-Fuller (ADF) unit root test has low power in detecting its mean reverting (stationary) tendency (Sarno, 2000; Sarno and Taylor, 2002). Based on linearity

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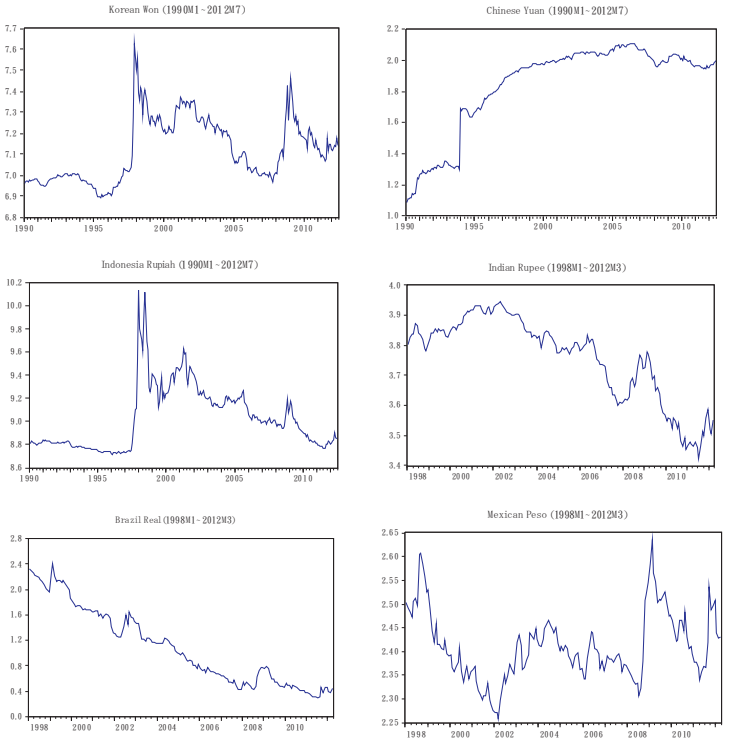


Figure 1: The Tendency of Real Exchange Rates Given Natural Logarithms for Six Growth Countries With US Base

tests (Teräsvirta, 1994), a number of studies have provided empirical evidence that exchange rates do exhibit nonlinear behavior in the developed countries (Baum, Barkoulas, and Caglayan, 2001; Elliott and Pesavento, 2006; Kilic, 2009), G-7 countries (Chortareas, Kapetanios, and Shin, 2002; Chang, Lee, and Chou, 2012), Euro-zone countries (Schnatz, 2007; Giannellis and Papadopoulos, 2010), the Middle East (Sarno, 2000) and Asian economies (Liew, Baharumshah, and Lau, 2002; Liew, Chong, and Lim, 2003; Sarno, 2000; Baum, Barkoulas, and Caglayan, 2001; Chortareas, Kapetanios, and Shin, 2002), amongst others, showed that the non-linear adjustment of exchange rate towards PPP can best be described by ESTAR model.[2] More recently, Liew, Chong, and Lim (2003) have demonstrated

[2]Teräsvirta (1994) provided theoretical details for ESTAR model, whereas Sarno (2000) and Baum, Barkoulas, and Caglayan (2001) demonstrated the usefulness of this model in

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that linear autoregressive model is inadequate in characterizing the Asian (including ASEAN-5) real exchange rates behaviors. Thus, it is reasonable for one to think that the puzzling results of PPP lie in the implicit linear assumption of exchange rate behavior. However, the finding of nonlinear adjustment does not necessarily imply nonlinear mean reversion (stationary). As such, formal stationary tests based on nonlinear framework must be applied.

4.1 Nonlinear ESTAR Unit Root Test

This paper aims to empirically investigate the validity of long run PPP for six high-growth countries. To fully capture the nonlinearities of variables investigated, our study employs a non-linear stationary test advanced by Kapetanios, Shinb, and Snell (2003) (henceforth, KSS) to determine the validity of PPP.

The KSS non-linear stationary test is alternatively based on detecting the presence of non-stationary against non-linear but globally stationary exponential smooth transition autoregressive (ESTAR) process:

$$\Delta y_t = G_t(\gamma, y_{t-1})\,\rho y_{t-1} + \nu_t, \tag{3}$$

$$G_t(\gamma, y_{t-1}) = 1 - \exp\left(-\gamma\left(y_{t-1}^2\right)\right), \quad \gamma \geq 0, \tag{4}$$

where y_t is the data series of variables considered, ν_t is an $i.i.d.$ error with zero mean and constant variance, and γ is the transition parameter of the ESTAR model that governs the speed of transition. If $\gamma = 0$ in Eq. (3), then y_t is non-stationary and contains a fixed unit root. Hence the relevant null hypothesis is $H_0 : \gamma = 0$. However, under this null hypothesis the parameter ρ is unidentified, and so testing this null using conventional methods is not feasible. One way to solve this identification problem is to replace $\{1 - \exp(-\gamma y_{t-1}^2)\}$ by its first-order Taylor expansion around $\gamma = 0$. After re-arranging, this gives the auxiliary model:

$$\Delta y_t = \delta y_{t-1}^3 + \varepsilon_t, \tag{5}$$

where $\varepsilon_t = \nu_t + \tau_t$, with τ_t denoting the remainder from the Taylor expansion and $\delta = \gamma\rho$. The null hypothesis is $H_0 : \delta = 0$ in Eq. (5). Kapetanios,

characterizing exchange rate adjustment.

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Shinb, and Snell (2003) deal with higher order dynamics by augmenting Eq. (5) with lags of Δy_t. The relevant auxiliary model can be written as:

$$\Delta y_t = \delta y_{t-1}^3 + \sum_{i=1}^{k} \eta_i \Delta y_{t-i} + \varepsilon_t. \tag{6}$$

The null hypothesis and alternative hypotheses are then expressed as $\delta = 0$ (non-stationarity) against $\delta \neq 0$ (non-linear ESTAR stationarity).

4.2 Advanced AESTAR Unit Root Test Proposed by Sollis (2009)

The ESTAR time series model has proved to be popular in economics for the analysis of time series data on the relative prices of equivalent goods or assets. But the assumption of symmetric mean reversion (linear or nonlinear) is not overly restrictive. Sollis (2009) proposed an extended version of the ESTAR model. The advantage of this method is that allows for symmetric or asymmetric nonlinear adjustment under the alternative hypothesis to a unit root. This model employs both an exponential function and a logistic function as follows:

$$\begin{aligned} \Delta y_t = {} & G_t(\gamma_1, y_{t-1}) \{S_t(\gamma_2, y_{t-1}) \rho_1 \\ & + (1 - S_t(\gamma_2, y_{t-1})) \rho_2\} y_{t-1} + \varepsilon_t, \end{aligned} \tag{7}$$

$$G_t(\gamma_1, y_{t-1}) = 1 - \exp\left(-\gamma_1 \left(y_{t-1}^2\right)\right), \quad \gamma_1 \geq 0, \tag{8}$$

$$S_t(\gamma_2, y_{t-1}) = \left[1 + \exp(-\gamma_2 y_{t-1})\right]^{-1}, \quad \gamma_2 \geq 0, \tag{9}$$

where $\varepsilon_t \sim iid(0, \sigma^2)$, and γ_i determining the speed of the transition. As with the original symmetric ESTAR model and the suggestion of Sollis (2009), the AESTAR model (Eq. (7)) can be extended to allow for higher-order dynamics:

$$\begin{aligned} \Delta y_t = {} & G_t(\gamma_1, y_{t-1}) \{S_t(\gamma_2, y_{t-1}) \rho_1 + (1 - S_t(\gamma_2, y_{t-1})) \rho_2\} y_{t-1} \\ & + \sum_{i=1}^{k} \kappa_i \Delta y_{t-i} + \varepsilon_t. \end{aligned} \tag{10}$$

Furthermore, Sollis, Leybourne, and Newbold (2002), Kapetanios, Shinb, and Snell (2003) and Park and Shintani (2005) suggest that do not allow for transitions in the higher-order dynamic terms in Eq. (10). The unit root hypothesis can be tested against the alternative hypothesis of globally

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stationary symmetric or asymmetric ESTAR nonlinearity with a unit root central regime by testing $H_0 : \gamma_1 = 0$ in Eq. (10). Unfortunately, γ_2, ρ_1 and ρ_2 are unidentified under this null, therefore conventional methods cannot be used. In terms of Kapetanios, Shinb, and Snell (2003), using an auxiliary model by taking a Taylor expansion of the exponential function in the original model around $\gamma = 0$ for testing. However, this treatment cannot fully avoid the problem of unidentified parameters. The suggestion of Sollis (2009) is to assume $\kappa = 0$ in Eq. (10) and replacing $G_t(\gamma_1, y_{t-1})$ with a first-order Taylor expansion around $\gamma_1 = 0$ gives

$$\Delta y_t = \rho_1\gamma_1 y_{t-1}^3 S_t(\gamma_2, y_{t-1}) + \rho_2\gamma_1 y_{t-1}^3 (1 - S_t(\gamma_2, y_{t-1})) + \eta_t. \quad (11)$$

Next, replacing $S_t(\gamma_2, y_{t-1})$ with $S_t^*(\gamma_2, y_{t-1})$ where $S_t^*(\gamma_2, y_{t-1}) = S_t(\gamma_2, y_{t-1}) - 0.5$, and substituting in Eq.(11) gives

$$\begin{aligned}\Delta y_t &= \rho_1\gamma_1 y_{t-1}^3 S_t^*(\gamma_2, y_{t-1}) + \rho_2\gamma_1 y_{t-1}^3 \left(1 - S_t^*(\gamma_2, y_{t-1})\right) \\ &\quad + 0.5\rho_1\gamma_1 y_{t-1}^3 - 0.5\rho_2\gamma_1 y_{t-1}^3 + \eta_t, \\ &= \rho_1^*\gamma_1 y_{t-1}^3 S_t^*(\gamma_2, y_{t-1}) + \rho_2^*\gamma_1 y_{t-1}^3 \left(1 - S_t^*(\gamma_2, y_{t-1})\right) + \eta_t, \quad (12)\end{aligned}$$

where ρ_1^* and ρ_2^* are linear functions of ρ_1 and ρ_2 with $\rho_1^* = -0.5\rho_1 + 1.5\rho_2$ and $\rho_2^* = 0.5\rho_1 + 0.5\rho_2$. Taking a Taylor expansion of $S_t^*(\gamma_2, y_{t-1})$ in Eq. (12) around $\gamma_2 = 0$, and after re-parameterization, the auxiliary regression can be written as follows:

$$\Delta y_t = a\left(\rho_2^* - \rho_1^*\right)\gamma_1\gamma_2 y_{t-1}^4 + \rho_2^*\gamma_1 y_{t-1}^3 + \eta_t, \quad (13)$$

where $a = 1/4$ and $\eta_t \sim iid(o, \sigma^2)$, which can be written as,

$$\Delta y_t = \phi_1 y_{t-1}^3 + \phi_2 y_{t-1}^4 + \eta_t, \quad (14)$$

where $\phi_1 = \rho_2^*\gamma_1$ and $\phi_2 = a(\rho_2^* - \rho_1^*)\gamma_1\gamma_2$. An augmented version is

$$\Delta y_t = \phi_1 y_{t-1}^3 + \phi_2 y_{t-1}^4 + \sum_{i=1}^{k} \kappa_i \Delta y_{t-i} + \eta_t. \quad (15)$$

The null hypothesis $H_0 : \gamma_1 = 0$ in Eq. (10) becomes $H_0 : \phi_1 = \phi_2 = 0$ in the auxiliary model (Eq. (15)).

A feature of the AESTAR model proposed is that if the unit root hypothesis has been rejected against the alternative of stationary symmetric or

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asymmetric ESTAR nonlinearity, the null hypothesis of symmetric ESTAR nonlinearity can then be tested against the alternative of asymmetric ESTAR nonlinearity using the auxiliary model (Eq. (15)) by testing $H_0 : \phi_2 = 0$ against $H_1 : \phi_2 \neq 0$ with a standard F-test (or t or LM test) (Teräsvirta, 1994). To clarify, it can be seen from Eq. (13) that if $\gamma_2 = 0$, the AESTAR auxiliary model (Eq. (15)) collapses to the ESTAR auxiliary model of Kapetanios, Shinb, and Snell (2003). For standard F critical values to be applicable for this test, $\phi_1 < 0$, so that under the null hypothesis being tested the series is stationary. Therefore in practice such a test using standard F critical values is only asymptotically valid if the consistent LS estimate of ϕ_1 is negative. Sollis (2009) indicated that for testing the unit root null hypothesis $H_0 : \phi_1 = \phi_2 = 0$ in Eq. (15), standard critical values cannot be used. Therefore, he derived the asymptotic distribution of an F-test of $H_0 : \phi_1 = \phi_2 = 0$ in Eq. (15), showing it to be a nonstandard function of Brownian motions. The F-test statistic can be written as follows:

$$F = \left(R\hat{\beta} - r\right)' \left[\hat{\sigma}^2 R \left\{\sum_t x_t x_t'\right\}^{-1} R'\right]^{-1} \left(R\hat{\beta} - r\right)/m, \qquad (16)$$

where $x_t = [y_{t-1}^3, y_{t-1}^4]'$, $m = 2$, R is a 2×2 identity matrix, $\hat{\beta} = [\hat{\phi}_1, \hat{\phi}_2]'$, $\hat{\phi}_1$ and $\hat{\phi}_2$ are the LS estimates of ϕ_1 and ϕ_2, $r = [0,\ 0]'$ and $\hat{\sigma}^2$ is the LS estimate of σ^2.

Finite-sample and asymptotic critical values for the F-test obtained by simulating under a random walk with $i.i.d.$ standard normal errors are given in Table 1 of Sollis (2009).

4.3 Empirical results

For comparison, first, we apply conventional ADF, PP (Phillips and Perron, 1988) and KPSS (Kwiatkowski et al., 1992) tests to examine the null of a unit root in the real exchange rate of each empirical country. Table 3 summarizes the results of the three unit root tests, ADF, PP, and KPSS, which show that the real exchange rates are non-stationary for six high-growth countries. As stated earlier, there is a growing consensus that the real exchange rate exhibits nonlinearities because of the existence of transaction costs, inflation and financial crisis, and consequently, conventional unit root tests such as the ADF test have low power in detecting the mean reversion of exchange rates. However, Perron (2006) indicates that no unit root is often supported

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Table 3: Univariate Unit Root Tests (Real Exchange Rates)

Country	ADF	PP	KPSS
Levels			
Brazil	−2.202(0)	−2.202(0)	0.314(10)***
China	1.961(0)	−1.115(5)	0.520(15)***
India	−2.489(0)	−2.552(4)	0.275(10)***
Indonesia	−0.036(2)	−2.226(3)	0.364(12)***
Korea	−0.196(2)	−0.211(14)	0.147(18)**
Mexico	−2.792(0)	−2.978(2)	0.142(10)*
1st Differences			
Brazil	−8.679(1)***	−12.577(4)***	0.181(4)
China	−18.145(0)***	−18.173(5)***	0.034(5)
India	−6.433(2)***	−11.774(3)***	0.182(4)
Indonesia	−13.022(1)***	−13.944(9)***	0.082(5)
Korea	−21.386(1)***	−34.559(9)***	0.051(14)
Mexico	−11.683(0)***	−11.665(2)***	0.069(0)

Notes: The critical values for the ADF t-statistics are from the MacKinnon (1996) table. The numbers in the parentheses of ADF is the appropriate lag lengths selected by MAIC, as suggested by Ng and Perron (2001), whereas the numbers in the parentheses of PP and KPSS are the optimal bandwidths decided by the Bartlett kernel of Newey and West (1994). ADF, Augmented Dickey–Fuller; PP, Phillips and Perron (1988); KPSS, Kwiatkowski et al. (1992). The nulls of ADF and PP are I(1), whereas the null of KPSS is I(0). ***, ** and * stand for 1%, 5% and 10% significant levels, respectively.

by incorporating structural break in the unit root test, thus we added the unit root test with structural break proposed by Perron (1997). The test results are reported in Table 4 and show that the majority of the high-growth countries exhibit no unit root after incorporating structural break point into the unit root test. Furthermore, a number of studies have also provided empirical evidence on the nonlinear adjustment of exchange rate. Therefore, we proceed to test the mean reversion behavior of exchange rate adjustment by applying the nonlinear unit root test of Kapetanios, Shinb, and Snell (2003). Table 5 presents the results of KSS's (2003) non-linear ESTAR unit root test, which indicates that bilateral real exchange rates are nonlinear sta-

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Table 4: Unit Root Test With Structural Break Base on Perron (1997)

Country	t_p	*Date*
Brazil	−4.641*	1999M03
China	−5.918***	1993M12
India	−2.512	2011M08
Indonesia	−6.342***	1998M02
Korea	−6.736***	1998M01
Mexico	−3.901	2011M08

Notes: t_p denotes the Perron (1997) unit root test and *Date* denotes the break date. The critical values of Perron (1997) test at the 1%, 5% and 10% significance levels of the t statistics are −5.57, −4.91 and −4.59, respectively. ***, ** and * stand significant at 1%, 5% and 10% level, respectively.

tionary with exception of India/USD real exchange rate.

Furthermore, we test the real exchange rate by using Sollis (2009) AESTAR nonlinear unit root test, the results are reported in Table 6 and substantiate that there is a unit root in real exchange rate for only one bilateral real exchange rate and it is for India/USD. The Sollis's AESTAR nonlinear unit root test yields the same results with the KSS's ESTAR test, indicating that real exchange rate are nonlinear processes that are not characterized by a unit root process consistent with the PPP for the majority of six high-growth countries, with the exception of India/USD.

Table 6 further indicates that the adjustment process for Indonesia/USD and Korea/USD real exchange rates are nonlinear and in an asymmetric way, since null hypothesis $H_0 : \phi_2 = 0$ against alternative hypothesis $H_1 : \phi_2 \neq 0$ was rejected for Indonesia and Korea. Moreover, Sollis (2009) unit root test confirms that there are nonlinear symmetric real exchange rates adjustment for Brazil/USD, China/USD and Mexico/USD. Sollis, Leybourne, and Newbold (2002) and Sollis (2009) suggest that asymmetric nonlinear mean reversion is an important feature of data on real exchange rates against the U.S. dollar. In fact one might expect asymmetry in the adjustment of the process toward its equilibrium for Indonesia and Korea, since they are more export-oriented (see Table 7). In the case of real

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Table 5: Nonlinear unit root test based on ESATR model of Kapetanios, Shinb, and Snell (2003)

Country	t-Statistics on $\hat{\delta}$
Brazil	$-2.754(1)^{**}$
China	$-3.332(1)^{***}$
India	$-1.161(2)$
Indonesia	$-3.301(2)^{***}$
Korea	$-3.434(2)^{***}$
Mexico	$-3.049(1)^{***}$

Notes: KSS, Kapetanios, Shinb, and Snell (2003). The critical values for t-statistic on $\hat{\delta}$ are tabulated in table 1 of Kapetanios, Shinb, and Snell (2003). The critical values for 10%, 5% and 1% are -1.92, -2.22 and -2.82, respectively. the number in parentheses indicates the lag order selected based on the recursive t-statistics, as suggested by Perron (1989). ** and *** significant at the 5% and 1% levels, respectively.

exchange rates, one might expect asymmetric adjustment if domestic or foreign policymakers behave asymmetrically in response to appreciations and depreciations of the same proportionate amount.

As indicated by Sollis (2009), when a rejection is obtained from the F statistics, it is interesting to estimate the AESTAR model in its raw form. We present the results for the case of Indonesia as an example. The real exchange rate series is plotted in Figure 2(a). The combination of fitted exponential and logistic functions multiplied by the nonlinear AR parameters obtained for the relevant AESTAR model:

$$G_t\left(0.604, y^*_{t-1}\right)\left\{S_t\left(4.344, y^*_{t-1}\right)(-0.222) + \left(1 - S_t\left(4.344, y^*_{t-1}\right)\right)(-0.898)\right\} \tag{17}$$

is plotted in Figure 2(b) against the threshold variable y^*_{t-1} with $y^*_{t-1} = y_{t-1} - \widehat{\mu}$. where $\widehat{\mu}$ is an estimate of the relevant population parameters obtained by LS regression prior to the ESTAR model being estimated. Clearly,

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Table 6: Nonlinear unit root tests based on AESTAR model of Sollis (2009) (with the US as the base country).

Country	ϕ_1	ϕ_2	$H_0: \phi_1 = \phi_2 = 0$	H_0: $\phi_2 = 0$	SIC	lags
Brazil	−0.016	−0.008	4.342*	−0.482	-7.762	0
China	−0.154	−0.153	5.901**	−1.012	-8.027	0
India	0.072	0.767	0.251	0.421	-9.597	13
Indonesia	−0.203	−0.589	44.357***	−3.752***	-6.861	14
Korea	−1.223	−2.490	29.614***	−2.100**	-6.359	14
Mexico	−4.147	−0.174	4.519*	−0.014	-1.408	0

Notes: The critical value of F-statistics for the null hypothesis of $\phi_1 = \phi_2 = 0$ are tabulated in Table 1 of Sollis (2009). ***, ** and * stand for 1%, 5% and 10% significant level. A feature of the proposed AESTAR model is that if the unit root hypothesis has been rejected against the alternative of stationary symmetric or asymmetric ESTAR nonlinearity, the null hypothesis of $\phi_2 = 0$ (symmetric ESTAR nonlinearity) can then be tested (following a standard F-distribution) against the alternative of AESTAR nonlinearity. Our choice of the appropriate lag length is based on Schwarz's information criteria (SIC; Schwarz (1978)). The auxiliary regression is:

$$\Delta y_t = \phi_1 y_{t-1}^3 + \phi_2 y_{t-1}^4 + \sum_{i=1}^{k} \kappa_i \Delta y_{t-i} + \eta_t.$$

It can be seen in Figure 2(b) that a high degree of asymmetry is estimated for this series (this is also manifest in Figure 2(a). In particular, for positive deviations from its attractor the real exchange rate y_t is much more persistent than for negative deviations of the same absolute magnitude. The combined function (Eq. (17)) varies between approximately −0.898 and 0 when the real exchange rate is below its attractor, but only between −0.222 and 0 when the real exchange rate is above its attractor. This supports the strong rejection of symmetric ESTAR nonlinearity obtained by the second-stage test as reported in Table 5. Thus, the fitted AESTAR model reveals that real appreciations of the dollar against the Indonesia Rupiah are slower to mean revert (nonlinearly) than real depreciations of the same proportion ate amo-

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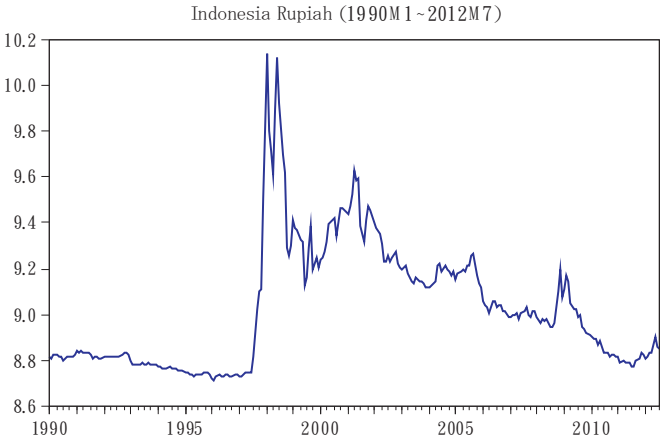
Table 7: Exports/GDP and imports/GDP ratios for 6 empirical countries (% of GDP)

	Korea		Indonesia		India		China		Mexico		Brazil	
	Export /GDP	Import /GDP	Export /GDP	Import /GDP	Export /GDP	Import /GDP	Export /GDP	Import /GDP	Export /GDP	Import /GDP	Export /GDP	Import /GDP
2000	35.0	32.9	41.0	30.5	12.8	13.7	23.3	20.9	26.3	26.9	10.0	11.7
2001	32.7	31.2	39.0	30.8	12.3	13.2	22.6	20.5	23.6	24.9	12.2	13.5
2002	30.8	29.3	32.7	26.4	14.0	15.0	25.1	22.6	23.5	24.9	14.1	12.6
2003	32.7	30.7	30.5	23.1	14.7	15.4	29.6	31.2	24.9	26.3	15.0	12.1
2004	38.3	34.5	32.2	27.5	17.6	19.3	34.1	30.4	26.2	28.0	16.4	12.5
2005	36.8	34.4	34.1	29.9	19.3	22.0	37.1	31.6	26.6	28.0	15.1	11.5
2006	37.2	36.4	31.0	25.6	21.1	24.2	39.1	29.5	27.6	28.8	14.4	11.5
2007	39.2	38.1	29.4	25.4	20.4	24.4	38.4	32.0	27.7	29.3	13.4	11.8
2008	50.0	50.0	29.8	28.8	23.6	28.7	35.0	27.3	27.9	30.2	13.7	13.5
2009	47.5	42.9	24.2	21.4	20.0	25.4	26.7	22.3	27.3	28.8	11.0	11.1
2010	49.4	46.2	24.6	22.9	22.0	26.3	29.4	25.6	29.9	31.1	10.9	11.9
2011	55.7	54.3	26.4	24.9	23.9	30.2	28.5	26.0	31.2	32.5	11.9	12.6
2012	56.3	53.5	24.3	25.9	24.0	30.7	27.3	24.5	32.6	33.7	12.6	14.0

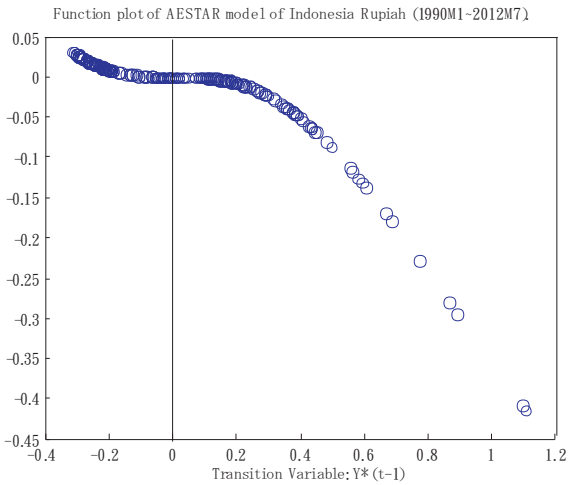
Notes: Data from World Development Indicators Database.

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2(a) Indonesia



2(b)



$$G_t(0.604, y^*_{t-1})\{S_t(4.344, y^*_{t-1})(-0.222)+(1-S_t(4.344, y^*_{t-1}))(-0.898)\}$$

Figure 2: (a) Real exchange rate of the Indonesia Rupiah against the USD, 1990M1–2012M7. (b) Function plot for AESTAR model of Indonesia Rupiah (strong asymmetry).

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unt. Note that the conventional ESTAR model as employed by Kapetanios, Shinb, and Snell (2003) does not explicitly take into account of this type of asymmetric behavior.

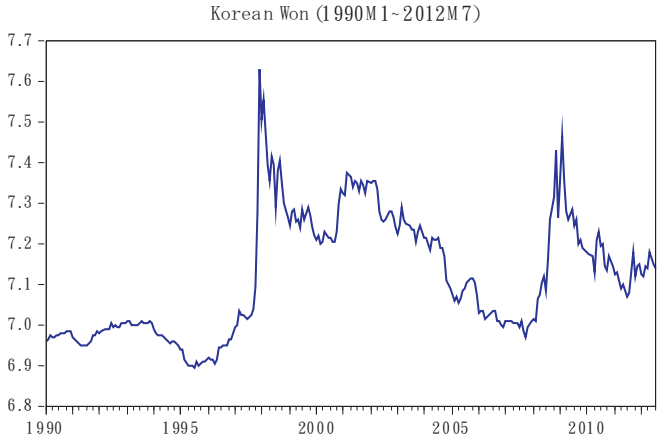
Figure 3 plots the case of South Korea. The real exchange rate series is plotted in Figure 3(a). The combination of fitted exponential and logistic functions multiplied by the nonlinear AR parameters obtained for the relevant AESTAR model is plotted in Figure 3(b) against the threshold variable y_{t-1}^* with $y_{t-1}^* = y_{t-1} - \widehat{\mu}$. Therefore, the fitted AESTAR model for South Korea reveals that real appreciations of the dollar against the Korea Won are slower to mean revert (nonlinearly) than real depreciations of the same proportionate amount. From the above results, we might assume asymmetry in the adjustment of the process toward its equilibrium for Indonesia and South Korea, since they are more export-oriented (see Table 7).

4.4 Robust Checking

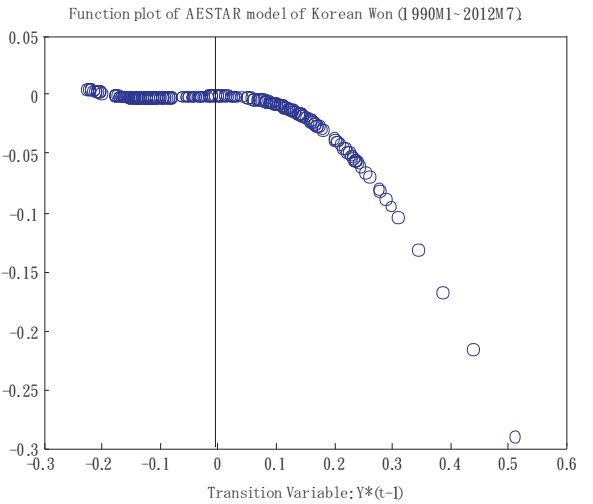
From Figure 2 and 3, it is clear that the real exchange rate adjustments in Indonesia and South Korea before and after the Asia financial crisis are quite different. Additionally, the sample in Table 7 only covers the period from 2000 to 2012. Therefore, we want to check whether the asymmetric adjustment is simply caused by period selection instead of whether the countries are more export-oriented or not. Does the Asia financial crisis influence the asymmetric adjustment behavior? To this end, we adopt the subsample period covered from the March 1998 to July 2012 for China, Indonesia and South Korea. Figure 4 and 5 plot the case of Indonesia and South Korea over the period from the March 1998 to July 2012, respectively. The results are reported in Table 8 and demonstrate rejection of the unit root hypothesis is only obtained for South Korea at the 1 percent level from the AESTAR test proposed here. Furthermore, the null hypothesis of symmetric ESTAR nonlinearity is rejected against the alternative of asymmetric ESTAR nonlinearity also at the 1 percent level of significance for Korean Won. In Indonesia, the rejections of the unit root hypothesis and symmetric ESTAR nonlinearity become insignificant when the estimated period is contracted from the January 1990 – July 2012 to the March 1998 – July 2012. Thus, these results may imply that the Asia financial crisis will influence the asymmetric adjustment behavior of the real exchange rate of Indonesia, and as the global financial crisis occurred, the real exchange rate of the economic con-

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3(a) South Korea



3(b)

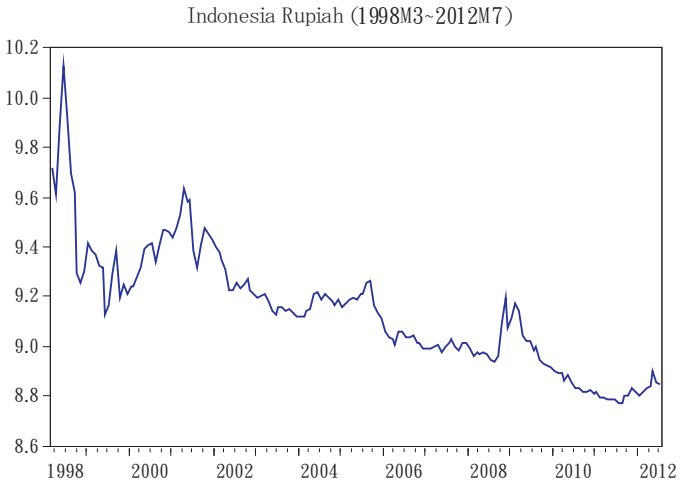


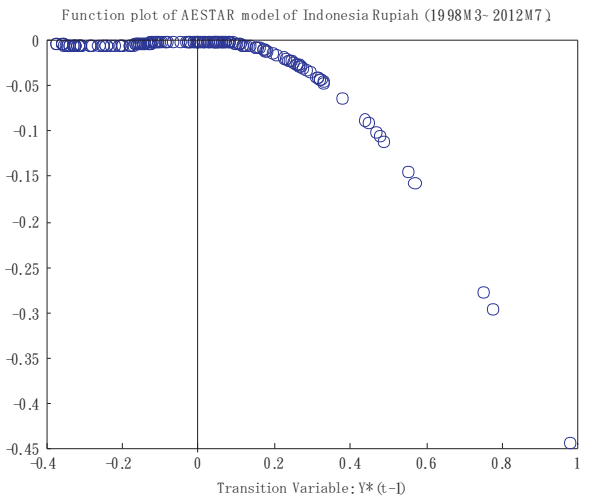
$$G_t(2.977,\ y^*_{t-1})\{S_t(3.405,\ y^*_{t-1})(-0.330)+(1-S_t(3.405,\ y^*_{t-1}))(-0.491)\}$$

Figure 3: (a) Real exchange rate of the Korean Won against the USD, 1990M1–2012M7. (b) Function plot for AESTAR model of Korea Won (strong asymmetry).

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4(a) Indonesia

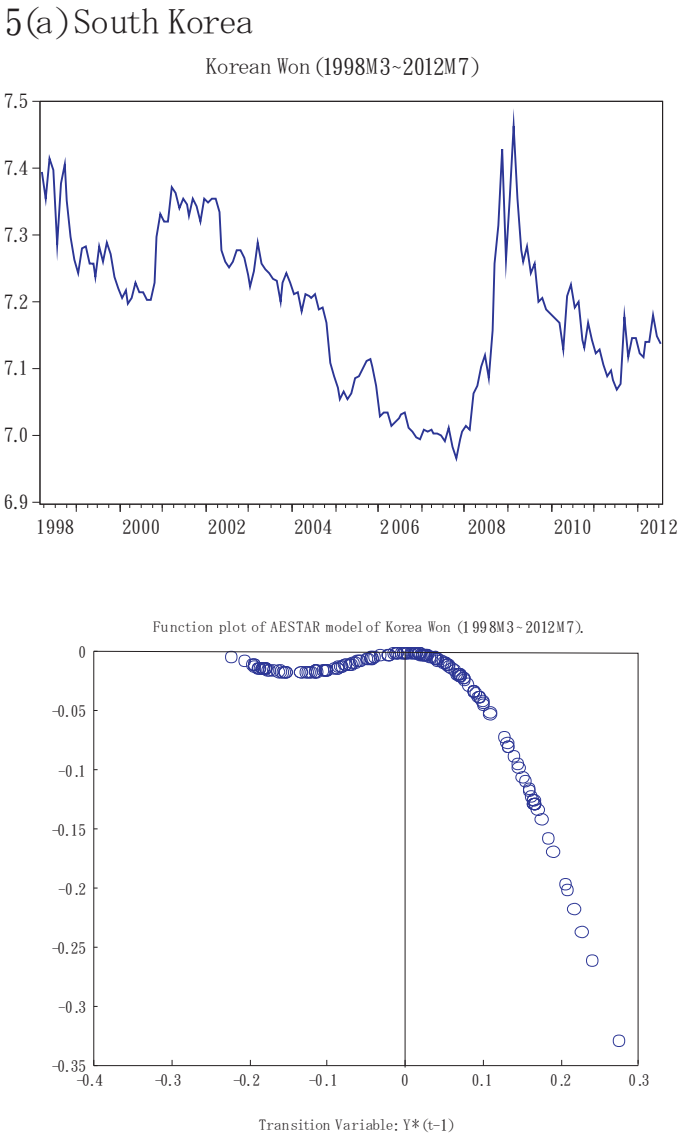




$$(0.809,\ y^*_{t-1})\{S_t(2.503,\ y^*_{t-1})(-0.923) + (1 - S_t(2.503,\ y^*_{t-1}))(-0.324)\}$$

Figure 4: (a) Real exchange rate of the Indonesia Rupiah against the USD, 1998M3–2012M7. (b) Function plot for AESTAR model of Indonesia Rupiah.

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$$G_t(10.639, y^*_{t-1})\{S_t(7.397, y^*_{t-1})(-0.699) + (1 - S_t(7.397, y^*_{t-1}))(-0.124)\}$$

Figure 5: (a) Real exchange rate of the Korea Won against the USD, 1998M3–2012M7. (b) Function plot for AESTAR model of Korea Won (strong asymmetry).

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Table 8: Robustness check: estimated results of the AESTAR model for the period 1998M3–2012M7

Country	ϕ_1	ϕ_2	$H_0 : \phi_1 = \phi_2 = 0$	$H_0 : \phi_2 = 0$	SIC SIC	lags lags
China	−8.331	47.506	2.092	0.790	−6.694	4
Indonesia	−0.253	−0.552	2.285	1.803	−2.712	13
Korea	−3.819	−15.087	19.149***	18.158***	−3.680	5

Notes: as the same Table 6.

stitution vulnerable countries would be directly affected.[3] This result can be confirmed by our results of the structural break tests, which are developed by Perron (1997).[4]

Finally, the evidence shows that the real exchange rate of Korean Won exhibits asymmetric nonlinearity adjustment behavior over the empirical period from January 1990 to July 2012, and is not influenced by the Asia financial crisis. In fact one might expect asymmetry in the adjustment of the process toward its equilibrium for South Korea, since she is more export-oriented (see Table 7). In the case of real exchange rates, one might expect asymmetric adjustment if domestic or foreign policymakers behave asymmetrically. This result makes intuitive sense for an exporting country since depreciations stimulate net exports and therefore policymakers may allow them to persist for longer than appreciations of the same proportionate amount.

4.5 Policy Implications

As we know that these six high-growth countries started their liberalization programs in the late 1980s and early 1990s. In some of these countries, this period was characterized by dramatic improvements in budget deficits,

[3]Su, Chang, and Chang (2011) argued that exchange rates might be affected by internal and external shocks generated by structural changes may be subject to considerable short-run variation.

[4]Marcela, Gadea, and Serrano (2003), Narayan (2005) and Narayan (2006), Narayan and Prasad (2005) also provide evidence that, when structural breaks are included for individual countries, real exchange rate is stationary, implying support for purchasing power parity.

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debts and inflation. As these countries became increasingly open to trade (inflation and growth rate converge to developed countries). We expect to find more favorable evidence of parity condition using data in recent years. As the reform process (finance liberalization and trade opening) intensified, we could expect a reduction in persistent shocks to international parity.

One major policy implication of our study is that the validity of using PPP to determine the equilibrium exchange is unequivocal – for most of these six high-growth countries, with exception of India. Avoiding the Asia financial crisis, PPP does not hold for Indonesia/USD over the March 1998 – July 2012 period. The governments of these four countries (i.e., Brazil, China, South Korea, and Mexico) can use the PPP to predict the exchange rate to determine whether a currency is overvalued or undervalued, as well as if the country is experiencing difference between domestic and foreign inflation rates. Nevertheless, reaping infinite gains from arbitrage in traded goods and services are not possible for most of the high-growth countries we study here.

Non-stationarity in real exchange rates has also macroeconomic implications. For Indonesia and India in which the PPP seem not to be valid, the depreciation of exchange rates result in permanent deviation from the equilibrium exchange rates. The deviation from the equilibrium exchange rates also may lead demand for labor and hence wages to shift in favor of tradable goods sectors compared to non-tradable (service) goods sector, implying a change in the terms of trade in favor of tradable against non-tradable sector for the countries with higher rate of economic growth (the Balassa–Samuelson effect). Besides, non-stationary real exchange rates provide room to implement a deprecation policy to deal with external imbalances.

5 Conclusions

This study uses nonlinear KSS (Kapetanios, Shinb, and Snell, 2003) unit root test and AESTAR unit root test, proposed by Sollis (2009), to test the validity of long-run PPP for six high-growth countries (i.e. China, Indonesia, India, Brazil, South Korea and Mexico). Based on the results from the KSS unit root test, we find that PPP holds true for the majority of six high-growth countries examined, with exception the India/USD real exchange rates.

Furthermore, from the AESTAR unit root test, Indonesia/USD and Korea/USD real exchange rate adjustment toward PPP are found to be nonlin-

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ear and in an asymmetric way over the period from January 1990 to July 2012. Avoiding the Asia financial crisis, PPP does not hold for Indonesia/USD for the period March 1998 through July 2012. On the other hand, Brazil/USD, China/USD and Mexico/USD real exchange rate adjustment toward PPP are found to be nonlinear and symmetric, while the traditional linear unit root tests such as the ADF, PP and KPSS, lead to no reject the unit root (non-stationary) hypothesis at all. Moreover, the fitted AESTAR model for Indonesia and South Korea illustrates that real appreciations of the dollar against these two currencies are slower to mean revert (nonlinearly) than real depreciation of the same proportionate amount. Apparently, our empirical results have important policy implications for these emerging high-growth countries under our study. As concerns major policy, the PPP can be used to determine the equilibrium exchange rate for most of the high-growth emerging countries.

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購買力平價及實質匯率非線性調整: 高成長國家之實證

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本文利用非線性 KSS 單根檢定及 Sollis (2009) 提出的非對稱指數平滑移轉自我迴歸 (簡稱 AESTAR) 單根檢定, 檢測六個高成長國家 (巴西、中國、印尼、印度、墨西哥和南韓) 在1990年1月至2012年7月間長期購買力平價是否成立。實證結果顯示, 採用 KSS 單根檢定, 除了印度外, 巴西、中國、印尼、墨西哥和南韓長期 PPP 成立。進一步, 利用 Sollis (2009) 提出的 AESTAR 單根檢定, 結果發現, 印尼盾和韓圜對美元購買力平價具非線性和非對稱特性, 巴西、中國和墨西哥兌美元匯率則具非線性和對稱性質。印尼盾升值時修正至均衡值速度, 較貶值時均數復歸速度緩慢; 韓圜則是貶值時修正至均衡值速度, 較升值時均數復歸速度緩慢。當國內外通膨率有所差異時, 這五國政府可以利用購買力平價預測匯率是否過高或過低, 此結果對新興成長國家財政政策決策者具有重要意涵。

關鍵詞: 購買力平價, 非線性單根檢定, 非對稱, AESTAR 檢定, 高成長國家
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