Nonlinear and Asymmetric Adjustment to Purchasing Power Parity in East-Asian Countries

Wen-Chi Liu

***Abstract*—**This study applies a simple and powerful nonlinear unit root test to test the validity of long-run purchasing power parity (PPP) in a sample of 10 East-Asian countries (i.e., China, Hong Kong, Indonesia, Japan, Korea, Malaysia, Philippines, Singapore, Taiwan and Thailand) over the period of March 1985 to September 2008. The empirical results indicate that PPP holds true for half of these 10 East-Asian countries under study, and the adjustment toward PPP is found to be nonlinear and in an asymmetric way.

***Keywords*—**Purchasing Power Parity, East-Asian Countries, Nonlinear Unit Root Test, Asymmetry.

I. Introduction

DURING the past several decades, empirical economic research has devoted to testing the validity of purchasing power parity (hereafter, PPP) hypothesis as it has important implications in the international macroeconomics. PPP states that the exchange rates between currencies are in equilibrium when their purchasing power is the same in each of the two countries. This means that the exchange rate between any two countries should equal to the ratio of two currencies' price level of a fixed basket of goods and services. The basic idea behind the PPP hypothesis is that since any international goods market arbitrage should be traded away over time, we should expect the real exchange rate to return to a constant equilibrium value in the long run. Studies on this issue are critical not only for empirical researchers but also for policymakers. In particular, a non-stationary real exchange rate indicates that there is no long-run relationship between nominal exchange rate and domestic and foreign prices, thereby invalidating the PPP. As such, PPP cannot be used to determine the equilibrium exchange rate, and an invalid PPP also disqualifies the monetary approach from exchange rate determination, which requires PPP to hold true.

For previous studies, one possible explanation for the inconsistencies in the existing empirical evidence on the purchasing power parity (PPP) hypothesis is that the prior studies implicitly assume that exchange rate behavior is inherently linear in nature. Taylor and Peel [1] demonstrate that the adoption of linear stationarity tests is inappropriate 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. The presence of nonlinear mean-reverting adjustment for real exchange rates has been advanced by recent theoretical developments that emphasize the role of transaction costs. Taylor et al. [2], Taylor and Peel [1], Juvenal and Taylor [3], and Lothian and Taylor [4] have argued that different speeds of adjustment at the disaggregated goods level average up to nonlinearity at the aggregate level. An alternative view is that nonlinearity at the aggregate level is caused by other influences, such as the effects of official foreign exchange intervention (Taylor [5]; Menkhof and Taylor [6]; Reitz and Taylor [7]) or heterogeneous agents (Kilian and Taylor [8]). For details on previous studies, please refer to the works of Taylor [9], Rogoff [10], MacDonald and Taylor [11]), Taylor and Sarno [12], Sarno and Taylor [13], Taylor and Taylor [14], and Lothian and Taylor ([4], [15]), who have provided in-depth information on the theoretical and empirical aspects of PPP and the real exchange rate. The majority of the models adopted in the prior empirical studies addressing the issue of equilibrium have generally failed to take into account the non-linear properties of the adjustment process; however, as noted by Laxton et al. [16], both bias and mistakes are increasingly likely when a linear and symmetrical methodology is adopted to test economic variables that are non-linear and asymmetric. This study analyzes PPP focus on the application of techniques that take into account the existence of nonlinearities. The first reason is related to the fact that the existence of trade barriers especially for East Asian countries and therefore, absence of arbitrage within exchange rate values, yields to a nonlinear behavior in the path of the variable (Kilian and Taylor [8]). Additionally, Taylor [5] claims that interventions in the foreign currency markets might generate a nonlinear behavior in the real exchange rate. Finally, the existence of structure changes in the real exchange rate might imply broken deterministic time trends and the result is a nonlinear pattern (Koop and Potter [17]).

The exponential smooth transition autoregressive (ESTAR) model has proven to be popular in economics for the analysis of time-series data, such as data on real exchange rates. The presence of transactions 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 capable of capturing this type of nonlinearity (see for example Baum et al. [18], Taylor et al. [2], and Sollis [19]). A number of tests on the unit root hypothesis against stationary ESTAR nonlinearity have recently been proposed (see for example Kapetanios et al. [20] and Park and Shintani [21]). However, the assumption of symmetric mean reversion (linear or nonlinear) in some empirical applications is too restrictive. One might expect asymmetry in the adjustment of the process toward its equilibrium. For example, in the case

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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 et al. [22] and Sollis [19] found evidence suggesting that asymmetric nonlinear mean reversion is an important feature of data on real exchange rates against the U.S. dollar.

The present study differs from those earlier examples by providing non-linear evidence whether the real exchange rate adjustment process toward its equilibrium is nonlinear in a symmetric or asymmetric way from a sample of 10 East Asian countries (i.e., China, Hong Kong, Indonesia, Japan, Korea, Malaysia, Philippines, Singapore, Taiwan and Thailand) using a simple and powerful nonlinear unit root test of Sollis [19]. Economic integration in Asia seems to be rising and Asia is also playing an important part of the world. The Asian nations are increasingly becoming key players in global markets based on their high levels of economic growth and of exports and imports. Increasing financial liberalization in East Asian countries since the mid-1980s has fuelled a lively debate regarding the optimum exchange rate regime for the region. Massive inflows of capital into these countries following their economic liberalization and financial deregulation in the early 1990s following played a key role in this respect and these inflows are not likely to diminish as these countries continue to deregulate and liberalize their financial markets. These countries which possess similar characteristics after undergoing various stages of financial liberalization provide a good platform for the study of financial integration (Baharumshah et al. [23]). With the liberalization of interest rates due to the open market policy and deregulation of financial markets, interest rates in the East Asian countries are expected to rise in the long term and are expected to be closely connected with the global markets. Moreover, the late 1990s economic turmoil that engulfed these countries has focused worldwide attention on several issues, including exchange rate dynamics of the Asian region.

With this, the current research hopes to fill the existing gap in the literature. To our knowledge, this study is the first, to date, that utilizes the asymmetric exponential smooth transition autoregressive (hereafter, AESTAR) unit root test in 10 East-Asian real exchange rates. With the exception of Sollis [19], he applies the same technique on the Nordic countries. We find that the AESTAR unit test strongly rejects the unit root process for half of the East-Asia countries (i.e., Indonesia, Japan, Malaysia, Taiwan, and Thailand), while the traditional unit root tests such as the ADF, PP (unit root null hypothesis), and KPSS (stationary null hypothesis) did not lead to rejection. Furthermore, the adjustment process toward its equilibrium for these five East-Asian countries is nonlinear and in an asymmetric way.

This paper is organized as follows. Section II presents the data used in our study. Section IIII briefly describes the AESTAR test and our empirical results. Section IV concludes the paper.

II. Data

Our empirical analysis covers a sample of 10 East-Asian countries: China, Hong Kong, Indonesia, Japan, Korea, Malaysia, Philippines, Singapore, Taiwan and Thailand. Monthly data are employed in this study, and the time span is from March 1985 to September 2008. All consumer price indexes, CPI (based on 2000 = 100), and nominal exchange rates relative to the U.S. dollar data are taken from the International Monetary Fund's International Financial Statistics CD-ROM. Each of the consumer price index and real exchange rate series was put into natural logarithms before the econometric analysis. Fig. 1 plots the real exchange rates series for these 10 country pairs. We find some significant upward and downward trends in the real exchange rate series. From these figures, for most of the series, some nonlinear adjustment patterns seem to be evident.

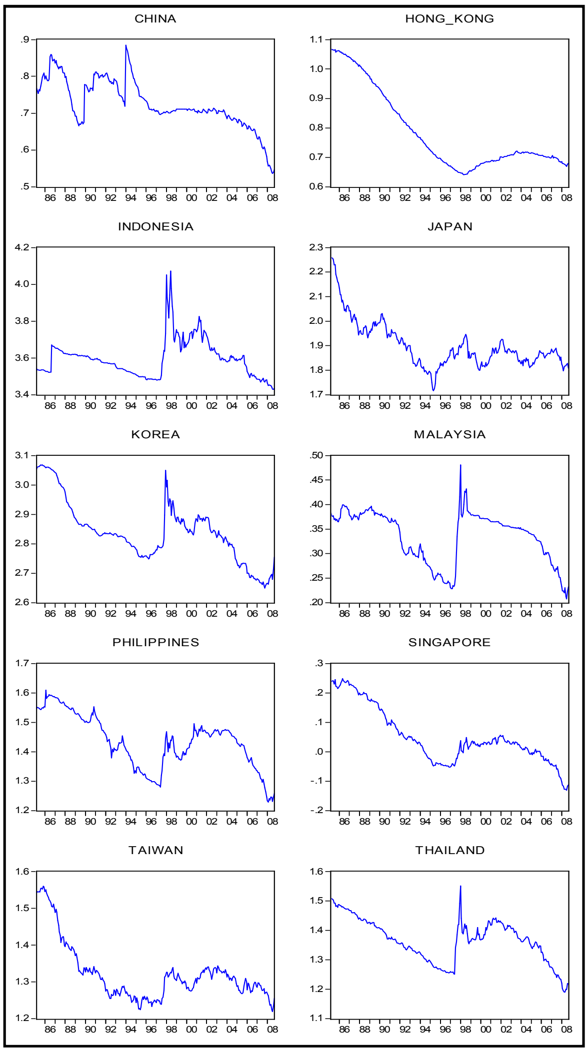


Fig. 1 The plots for real exchange rates of ten Asian countries

III. Methodology and Empirical Results

A. The AESTAR Unit Root Test Proposed by Sollis [19]

Sollis [19] proposed testing the unit root hypothesis using an extended version of the ESTAR model that allows for symmetric or asymmetric nonlinear adjustment under the alternative hypothesis to a unit root. The extended ESTAR model developed here, the AESTAR model, employs both an exponential function and a logistic function as follows:

$$\Delta y_t = G_t(\gamma_1, y_{t-1})\{S_t(\gamma_2, y_{t-1})\rho_1 + (1 - S_t(\gamma_2, y_{t-1}))\rho_2\}y_{t-1} + \varepsilon_t \quad (1)$$

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

$$S_t(\gamma_2, y_{t-1}) = [1 + \exp(-\gamma_2(y_{t-1})]^{-1} \quad \gamma_2 \geq 0 \quad (3)$$

where y_t is the data of series interest and ε_t ~iid $(0, \sigma^2)$. Assuming for the purposes of exposition that $\gamma_1 > 0$ and $\gamma_2 \to \infty$, as y_{t-1} moves from zero towards $-\infty$ then since $S_t(\gamma_2, y_{t-1}) \to 0$, an ESTAR transition occurs between the central regime model,

$$\Delta y_t = \varepsilon_t \quad (4)$$

and the outer-regime model

$$\Delta y_t = \rho_2 y_{t-1} + \varepsilon_t \quad (5)$$

with γ_1 determining the speed of the transition. As y_{t-1} moves from zero towards ∞, then since $S_t(\gamma_2, y_{t-1}) \to 1$ an ESTAR transition occurs between the central regime model,

$$\Delta y_t = \varepsilon_t \quad (6)$$

and the outer-regime model,

$$\Delta y_t = \rho_1 y_{t-1} + \varepsilon_t \quad (7)$$

with γ_1 determining the speed of the transition. If ρ1≠ρ2, the autoregressive adjustment is asymmetric on either side of the attractor (in this case the attractor is zero). Global stationarity requires $\rho_1 < 0$, $\rho_2 < 0$, $\gamma_1 > 0$. Note that (1) nests the symmetric ESTAR specification of Kapetanios et al. [20], since if $\rho_1 = \rho_2 = \rho$, then (1) is equivalent to (1) of Kapetanios et al. ([20], see page 361 of Kapetanios et al.'s paper).

In the above example of asymmetry, for the purposes of exposition we assume that $\gamma_2 \to \infty$, in which case $S_t(\gamma_2, y_{t-1})$ reduces to a simple step function. It is then clear that the composite function

$$G_t(\gamma_1, y_{t-1})\{S_t(\gamma_2, y_{t-1})\rho_1 + (1 - S_t(\gamma_2, y_{t-1}))\rho_2\} \quad (8)$$

which here can be thought of as the first-order AR parameter (minus 1) at each t, is symmetric or asymmetric depending on the values of ρ_1 and ρ_2 . However, assuming $\rho_1 \neq \rho_2$, asymmetry can also occur for small and moderate values of γ_2, which generate a gradual transition of $S_t(\gamma_2, y_{t-1})$ between its limiting values. For $\gamma_2 \to 0$, it follows that $S_t(\gamma_2, y_{t-1}) \to 0.5$ $\forall$ t, and consequently the composite function (8) becomes symmetric irrespective of the values of ρ_1 and ρ_2 . Thus, for a particular value of ($\rho_2 - \rho_1$), γ_2 ultimately controls the degree of asymmetry. This turns out to be a useful feature of the model for deriving a test of symmetric ESTAR nonlinearity versus asymmetric ESTAR nonlinearity.

As with the original symmetric ESTAR model, the AESTAR model (1) can be extended to allow for higher-order dynamics:

$$\Delta y_t = G_t(\gamma_1, y_{t-1})\{S_t(\gamma_2, y_{t-1})\rho_1 + (1 - S_t(\gamma_2, y_{t-1}))\rho_2\}y_{t-1} + \sum_{i=1}^{k} k_i \Delta y_{t-i} + \varepsilon_t \quad (9)$$

We follow Sollis et al. [22], Kapetanios et al. [20], and Park and Shintani [21] and do not allow for transitions in the higher-order dynamic terms in (9).

B. Tests of the AESTAR Unit Root Hypothesis

The unit root hypothesis can be tested against the alternative hypothesis of globally stationary symmetric or asymmetric ESTAR nonlinearity with a unit root central regime by testing $H_0 : \gamma_1 = 0$ in (9). Unfortunately, γ_2, ρ_1 and ρ_2 are unidentified under this null, thus conventional methods cannot be used. As the treatment of Kapetanios et al. [20], using an auxiliary model by taking a first-order Taylor expansion of the exponential function in the original model around γ=0 (a first-order expansion is used) for testing. However, this treatment cannot fully avoid the problem of unidentified parameters. The suggestion of Sollis [19] is to simplify the model further by taking a Taylor expansion of the logistic function. The detailed discussion see page 121, Sollis [19]. For the purposes of testing, the resulting model is

$$\Delta y_t = a(\rho_2^* - \rho_1^*)\gamma_1\gamma_2 y_{t-1}^4 + \rho_2^*\gamma_1 y_{t-1}^3 + \eta_t \quad (10)$$

where $a = 1/4$, which can be written as

$$\Delta y_t = \phi_1 y_{t-1}^3 + \phi_2 y_{t-1}^4 + \eta_t \quad (11)$$

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

$$\Delta y_t = \phi_1 y_{t-1}^3 + \phi_2 y_{t-1}^4 + \sum_{i=1}^{k} k_i \Delta y_{t-i} + \eta_t \, . \quad (12)$$

The null hypothesis $H_0 : \gamma_1 = 0$ in (1) becomes

$$H_0 : \phi_1 = \phi_2 = 0 \quad (13)$$

in the auxiliary model (13).

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 symmetric ESTAR nonlinearity can then be tested against the alternative of asymmetric ESTAR nonlinearity using the auxiliary model (13) by testing $H_0:\phi_2=0$ against $H_0:\phi_2\neq 0$ with a standard F-test (or t or LM test). To clarify, it can be seen from (12) that if γ_2=0, the AESTAR auxiliary model (12) collapses to the ESTAR auxiliary model of Kapetanios et al. [20]. For standard F critical values to be applicable for this test, $\phi_1<0$, so that under the null 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 [19] indicates that for testing the unit root null $H_0:\phi_1=\phi_2=0$ in (13), standard critical values cannot be used. Therefore, he derives the asymptotic distribution of an F-test of $H_0:\phi_1=\phi_2=0$ in (14), showing it to be a nonstandard function of Brownian motions. The test statistic can be written in the usual way,

$$F=(R\hat{\beta}-r)'\left[\hat{\sigma}^2R\left\{\sum_t x_t x_t'\right\}^{-1}R'\right]^{-1}(R\hat{\beta}-r)/m\cdot \quad (14)$$

Assuming k=0 in (12), it follows that $x_t=[y_{t-1}^3,y_{t-1}^4]'$, m=2, **R** is a 2×2 identity matrix, $\hat{\beta}=[\hat{\phi}_1,\hat{\phi}_2]'$, where $\hat{\phi}_1$ and $\hat{\phi}_2$ are LS estimates of ϕ_1 and ϕ_2, $r=[0,0]'$ and $\hat{\sigma}$ is the LS estimate of σ^2. Finite-sample and asymptotic critical values for the test obtained by simulation under a random walk with iid standard normal errors are given in Table I of Sollis [19].

C. Empirical Results

For the sake of comparison, we also incorporate the Augmented Dickey-Fuller (ADF), PP (Phillips and Perron [24]), and KPSS (Kwiatkowski et al. [25]) tests into our study. The three tests (i.e., ADF, PP (Phillips and Perron [24]), and the KPSS (Kwiatkowski et al. [25]) without a trend function are reported in Table I. In our study, we only consider a specification with a constant but without a time trend because time trend in real exchange rates is not consistent with the long-run PPP. Results from Table I clearly indicate that the ADF and the PP tests fail to reject the null hypothesis of non-stationary real exchange rates for all East-Asian countries, with the exception of Indonesia. The KPSS test also yields similar results indicating that the real exchange rates in East-Asian countries are non-stationary.

TABLE I
UNIVARIATE UNIT ROOT TESTS

	Level			1st difference		
	ADF	PP	KPSS	ADF	PP	KPSS
China	-2.5 (0)	-2.4 (6)	1.7 [14]***	-16.9 (0)***	-17.1(7)***	0.5 [7]**
Hong Kong	-2.3 (12)	-2.0(11)	0.6[14]**	-1.5 (11)	-15.1(11)***	1.2[11]***
Indonesia	-3.0(0)**	-2.8(6)*	1.0[14]***	-17.8(0)**	-17.9(2)***	0.1[2]
Japan	-2.3 (0)	-2.3 (8)	0.9[14]***	-15.8 (0)***	-15.8(14)***	0.1 [12]
Malaysia	-1.6 (0)	-1.7(2)	1.5[14]***	-15.6(0)***	-15.6(1)***	0.1 [1]
Philippine	-1.4 (1)	-1.6(6)	0.6[14]**	-16.1(0)***	-16.1(5)***	0.1 [6]
Singapore	-1.22(0)	-1.2 (5)	0.4[14]*	-15.6(0)***	-15.5(8)***	0.2[6]
South Korea	-2.2(2)	-2.3(3)	0.4 [14]*	-12.3 (1)***	-12.7(3)***	0.1[1]
Taiwan	-0.9(0)	-1.1(7)	1.4[14]***	-14.8(0)***	-14.8(6)***	0.5[6]**
Thailand	-1.5 (0)	-1.6 (5)	1.0[14]***	-15.8(0)***	-15.8(3)***	0.1[4]

Note: ***, ** and * indicate significance at the 0.01, 0.05 and 0.1 level, respectively. The number in parenthesis indicates the lag order selected based on the recursive t-statistic, as suggested by Perron [26]. The number in the brackets indicates the truncation for the Bartlett Kernel, as suggested by the Newey-West test [27].

As we know that there is a growing consensus that the real exchange rate exhibits nonlinearities, and consequently, conventional unit root tests such as the ADF test, have low power in detecting the mean reversion of exchange rate. A number of studies have also provided empirical evidence on the nonlinear adjustment of exchange rate. Therefore, we proceed to test the real exchange rate by using Sollis's [19] AESTAR nonlinear unit root tests. Sollis's [19] AESTAR nonlinear unit root test results substantiate that there is a unit root in the real exchange rate for only half of the bilateral real exchange rates. These are China/USD, Hong Kong/USD, Korea/USD, Philippines/USD and Singapore/USD, as shown in Table II. These results indicate that PPP holds true for half of these 10 East-Asian countries. Table II further indicates that the adjustment process for these five stationary real exchange rates is nonlinear and in an asymmetric way because $H_0:\phi_2=0$ against $H_0:\phi_2\neq 0$ was strongly rejected. Our results are consistent with those of Sollis et al. [22] and Sollis [19]. Both studies found evidence suggesting that asymmetric nonlinear mean reversion is an important feature of the data on real exchange rates against the U.S. dollar. In fact, one might expect the asymmetry in the adjustment of the process toward its equilibrium for these 10 East-Asian countries since these countries are more export-oriented. 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.

Open Science Index, Economics and Management Engineering Vol:8, No:2, 2014 publications.waset.org/9997317.pdf

TABLE II
AESTAR UNIT ROOT TEST OF SOLLIS [19]

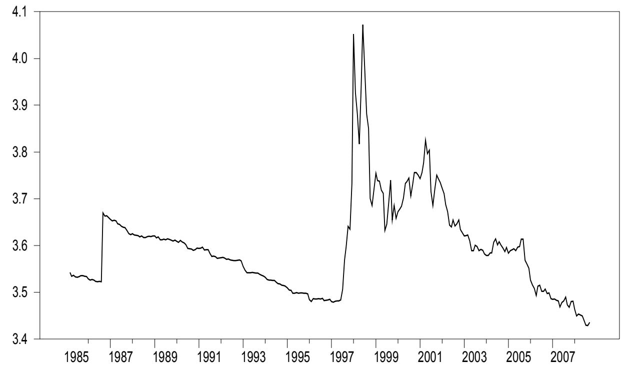
Country	ϕ_1	ϕ_2	$H_0:\phi_1=\phi_2=0$	$H_0:\phi_2=0$
China	-1.473	-11.787	2.162(1)	**NA**
Hong Kong	-0.213	0.651	3.875(8)	**NA**
Indonesia	0.878	-5.027	**22.255(1)*****	**4.881****
Japan	-1.911	4.553	**8.654(1)*****	**6.22****
Korea	-0.489	-0.268	2.168(1)	NA
Malaysia	-20.583	-157.94	**130.961(7)*****	**197.254*****
Philippines	-0.4325	-1.711	0.581(2)	NA
Singapore	-0.013	-1.687	1.379(1)	NA
Taiwan	-5.019	-30.484	**27.890(5)*****	**40.195*****
Thailand	-4.172	17.606	**6.147(4)****	**8.744*****

Notes:

1. The $F_{AE,\mu}$ statistic for the null hypothesis of $\phi_1 = \phi_2 = 0$ are tabulated at Table I of Sollis [19], and the critical values are 6.806, 4.971, and 4.173 at the 1%, 5% and 10% significant level, respectively. A feature of the AESTAR model proposed 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 asymmetric ESTAR nonlinearity.
2. ** and *** indicate significance at the 5% and 1% level, respectively.
3. Our choice of the appropriate lag length is based on the multivariate AIC.

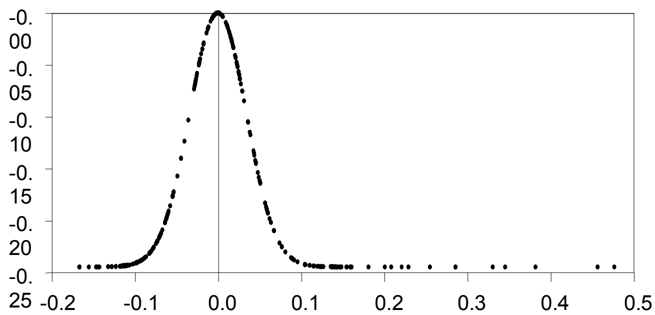
As indicated by Sollis [19], when a rejection is obtained from the F statistics, it is interesting to estimate the AESTAR model in its raw form and compare it graphically with the ESTAR model in its raw form. We present the results for the case of Indonesia as an example. The real exchange rate series is plotted in Fig. 2 (a). The fitted exponential function multiplied by the nonlinear AR parameter for the relevant ESTAR model, $G_t(320.2148, y_{t-1})(-0.2449)$, is plotted in Fig. 2 (b) against the threshold variable $y^*_{t-1}(y^*_{t-1} = y_{t-1} - \hat{u})$.

The combination of fitted exponential and logistic functions multiplied by the nonlinear AR parameters obtained for the relevant AESTAR model (allowing for a non-zero mean), $G_t(355.402, y_{t-1})\{S_t(4253.421, y_{t-1})(-0.175) + (1 - S_t(4253.421, y_{t-1}))(-0.81)\}$, is plotted in Fig. 2 (c) against the threshold variable $y^*_{t-1}(y^*_{t-1} = y_{t-1} - \hat{u})$. Clearly, it can be seen in Fig. 2 (c) that a high degree of asymmetry is estimated for this series (this is also visible in Fig. 2 (a). In particular, for positive deviations from its attractor, the real exchange rate is much more persistent than for negative deviations of the same absolute magnitude. The combined function varies between approximately −0.81 and 0 when the real exchange rate is below its attractor, but only between −0.175 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 reported in Table II. Thus, the fitted AESTAR model reveals that real appreciations of the dollar against the Indonesian Rupiah are slower to mean revert (nonlinearly) than real depreciations of the same proportionate amount. Note that the conventional ESTAR model as employed by Kapetanios et al. [20] and Park and Shintani [21] do not explicitly take into account this type of asymmetric behavior. Since these 10 East-Asian countries are more export-oriented, for example, in the case of real exchange rates, one might expect asymmetric adjustment when domestic or foreign policymakers behave asymmetrically in response to appreciations and depreciations of the same proportionate amount.



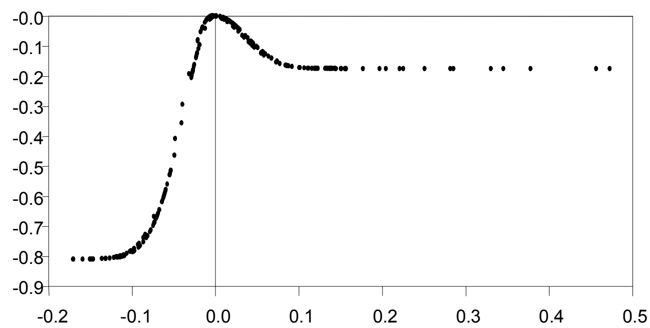
(a)

$$\Delta y_t = G_t\left(430.2148, y_{t-1}\right)\left(-0.2449\right)$$



(b)

$$\Delta y_t = G_t\left(355.402, y_{t-1}\right)\left\{S_t\left(4253.421, y_{t-1}\right)\left(-0.175\right) + \left(1 - S_t\left(4253.421, y_{t-1}\right)\right)\left(-0.81\right)\right\}$$



(c)

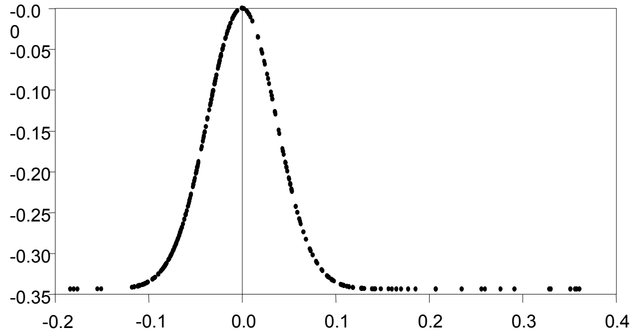
Fig. 2 (a) Real exchange rate of the Indonesia against the U.S. dollar 1985:3 - 2008:9, (b) Function plot for Indonesia: ESTAR model, (c) Function plot for Indonesia: AESTAR model

Figs. 3-6 plot the cases of Japan, Malaysia, Taiwan, and Thailand. The real exchange rate series is plotted in the top of each figure, and the fitted exponential function multiplied by the nonlinear AR parameter for the relevant ESTAR model is plotted in the middle of each figure against the threshold variable $y^*_{t-1}(y^*_{t-1} = y_{t-1} - \hat{u})$. The combination of fitted exponential and logistic functions multiplied by the nonlinear AR parameters obtained for the relevant AESTAR model (allowing for a non-zero mean) is plotted in the bottom of each figure against the threshold variable $y^*_{t-1}(y^*_{t-1} = y_{t-1} - \hat{u})$. Again, we found that the fitted AESTAR model for these four countries reveals that real appreciations of the dollar against these four currencies are slower to mean revert (nonlinearly) than real depreciations of the same proportionate amount.

Apparently, our empirical results from the Sollis' AESTAR test provide strong evidence favoring the long-run validity of PPP for the 10 East Asian countries under study. Therefore, it is possible to claim that deviations in the short-run form the PPP are not prolonged for most of the East Asian countries and there are some forces which are capable of bringing the exchange rate back to its PPP values in the long-run. The major policy implication that emerges from our study is that PPP can be used to determine the equilibrium exchange rates for all these 10 East Asian countries and the unbounded gains from arbitrage in traded good are not possible among these 10 East Asian countries. Our results also had important policy implication on cross-border agreement for international trade and investment with these countries. Given the goods and services markets appeared quite integrated, future liberalization will be likely pronounced in financial markets. If we envision this process of integration continuing, in particular in the Asian region, and to the extent that this process requires even more political engagement, we believe the prospects for cooperation along a variety of dimensions are good. Apparently, our empirical results have important policy implications for these 10 East-Asian countries under study.

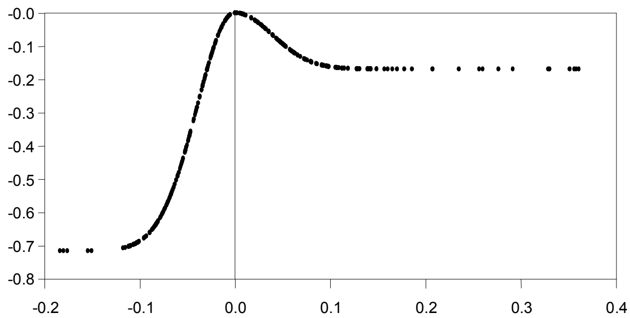
(a)

$$\Delta y_t = G_t(367.961, y_{t-1})(-0.344)$$



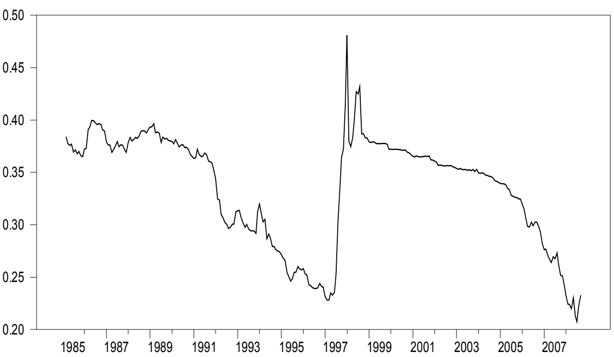
(b)

$$\Delta y_t = G_t(320.303, y_{t-1})\{S_t(3866.573, y_{t-1})(-0.169) + (1 - S_t(3866.573, y_{t-1}))(-0.716)\}$$



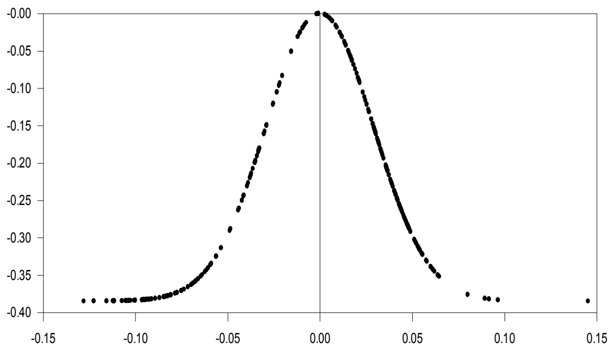
(c)

Fig. 3 (a) Real exchange rate of the Japan against the U.S. dollar 1985:3 - 2008:9, (b) Function plot for Japan: ESTAR model, (c) Function plot for Japan: AESTAR model



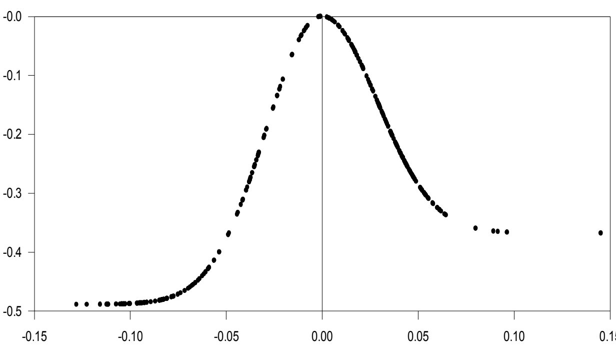
(a)

$$\Delta y_t = G_t\left(588.999, y_{t-1}\right)\left(\text{-}0.385\right)$$



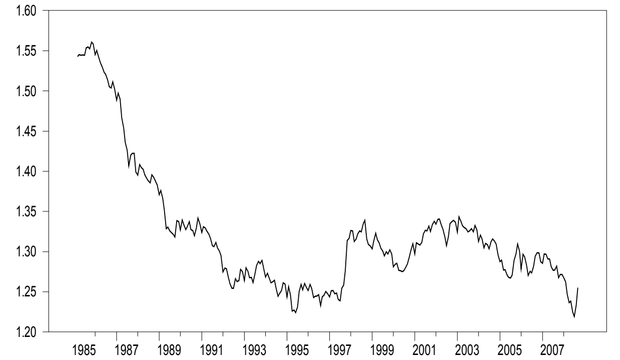
(b)

$$\Delta y_t = G_t\left(594.624, y_{t-1}\right)\left\{S_t\left(5548.703, y_{t-1}\right)\left(\text{-}0.368\right) + \left(1 - S_t\left(5548.703, y_{t-1}\right)\right)\left(\text{-}0.489\right)\right\}$$



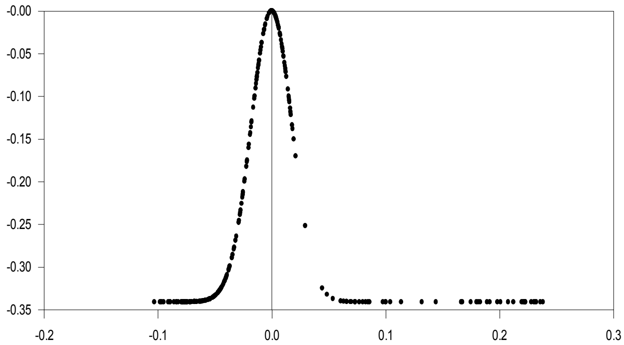
(c)

Fig. 4 (a) Real exchange rate of the Malaysia against the U.S. dollar 1985:3 - 2008:9, (b) Function plot for Malaysia: ESTAR model, (c) Function plot for Malaysia: AESTAR model



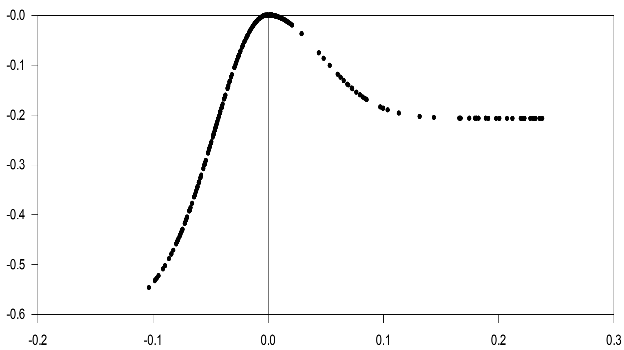
(a)

$$\Delta y_t = G_t\left(1547.028, y_{t-1}\right)\left(\text{-}0.341\right)$$



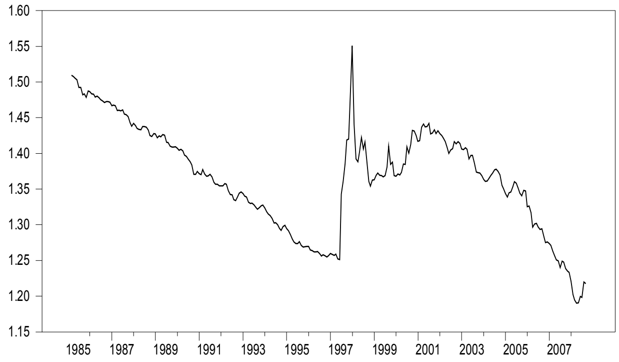
(b)

$$\Delta y_t = G_t\left(231.921, y_{t-1}\right)\left\{S_t\left(6885.041, y_{t-1}\right)\left(\text{-}0.207\right) + \left(1 - S_t\left(6885.041, y_{t-1}\right)\right)\left(\text{-}0.597\right)\right\}$$



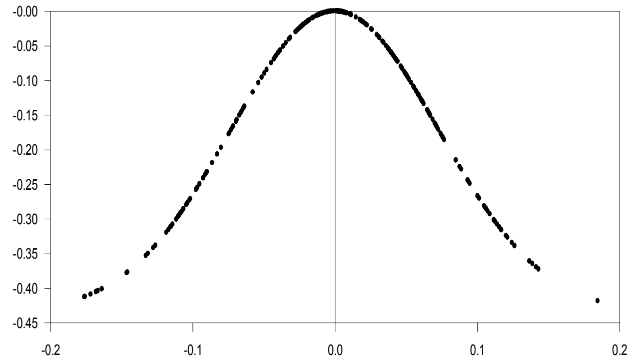
(c)

Fig. 5 (a) Real exchange rate of the Taiwan against the U.S. dollar 1985:3 - 2008:9, (b) Function plot for Taiwan: ESTAR model, (c) Function plot for Taiwan: AESTAR model



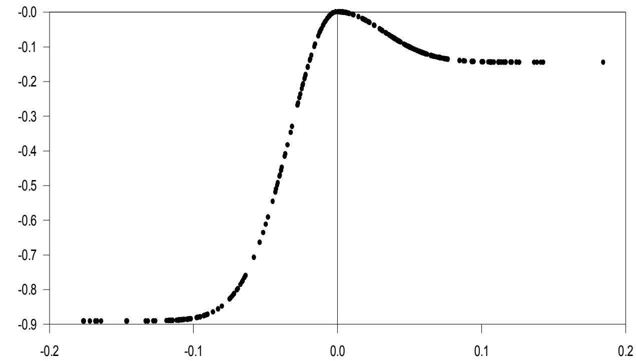
(a)

$$\Delta y_t = G_t\left(94.024, y_{t-1}\right)\left(\text{-}0.436\right)$$



(b)

$$\Delta y_t = G_t\left(474.669, y_{t-1}\right)\left\{S_t\left(4028.848, y_{t-1}\right)\left(\text{-}0.145\right) + \left(1 - S_t\left(4028.848, y_{t-1}\right)\right)\left(\text{-}0.892\right)\right\}$$



(c)

Fig. 6 (a) Real exchange rate of the Thailand against the U.S. dollar 1985:3 - 2008:9, (b) Function plot for Thailand: ESTAR model, (c) Function plot for Thailand: AESTAR model

IV. Conclusions

This study applies a simple and powerful nonlinear AESTAR unit root test proposed by Sollis [19] to test the validity of long-run PPP in a sample of 10 East-Asian countries (i.e., China, Hong Kong, Indonesia, Japan, Korea, Malaysia, Philippines, Singapore, Taiwan and Thailand) over the period of March 1985 to September 2008. The empirical results indicate that PPP holds for half of the East-Asian countries studied, and the adjustment toward PPP is nonlinear and in an asymmetric way.

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