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Hyeongwoo Kim[†] and Young-Kyu Moh[*]

[†]*Auburn University*, [*]*Sookmyung Women's University*

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Hyeongwoo Kim† and Young-Kyu Moh‡

Auburn University and Sookmyung Women's University

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Abstract

Researchers have encountered difficulties in finding empirical evidence of Purchasing Power Parity (PPP) especially when conventional linear unit root tests are employed for the Japanese yen real exchange rate. Chortareas and Kapetanios (2004), however, report strong evidence in favor of a Balassa-Samuelson type model of PPP by applying a nonlinear unit root test by Kapetanios *et al.* (2003) for the other G7 and Asian currencies relative to the Japanese yen, claiming that the yen real exchange rate may be (trend) stationary. We question the validity of this remark. First, we note that their claim is upset when we extend the data until 2008 even when the same nonlinear unit root test is used. Second, we apply the inf-t test by Park and Shintani (2005, 2010) which does not require the Taylor approximation, and find strong evidence against nonstationarity for most yen real exchange rates. Our results also corroborate the findings of Kim and Moh (2010) who report a possibility of misspecification problems with the use of Taylor-approximation based tests.

Keywords: Purchasing Power Parity; Transition Autoregressive Process; Nonlinear Adjustments; inf-t Unit Root Test

JEL Classification: C22; F31

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†Department of Economics, Auburn University, Auburn, AL 36849. Tel: 1-334-844-2928. Fax: 1-334-844-4615. Email: gmmkim@gmail.com.

‡Corresponding author: Young-Kyu Moh, Department of Economics, Sookmyung Women's University, Seoul, 140-742, Korea. Tel: 82-2-2077-7685. Fax: 82-303-0799-0381. Email: ykmoh@sookmyung.ac.kr.

1 Introduction

Purchasing power parity (PPP) serves as a key building block for many open macroeconomy models. Empirically testing PPP is typically carried out by implementing unit root tests for real exchange rates. Studies that employ conventional augmented Dickey-Fuller (ADF) type unit root tests often find weak evidence of PPP when the current float (post Bretton Woods system) real exchange rates are used. For example, Papell and Theodoridis (1998) show that the evidence of PPP is overall weak when the US dollar is used as a base currency, whereas they find stronger evidence of PPP with the German Deutschmark real exchange rate. Similarly, Papell and Theodoridis (2001) report very weak evidence of PPP when the Japanese yen serves as a base currency. Even though empirical evidence on PPP still remains elusive, the profession seems to find it less ambiguous that the yen real exchange rate is better approximated as nonstationary. See among others, Kim (1990), Cheung and Lai (1998), and Koedijk, Schotman, and van Dijk (1998).

Recognizing the difficulties in finding evidence in favor of (Casselian view of) PPP, Chortareas and Kapetanios (2004) investigate a weaker version of PPP for the yen real exchange rate as described below.[1]

First, Chortareas and Kapetanios (2004) notice a presence of a trend/drift in most yen real exchange rates they consider. With such observations, they test the null hypothesis of nonstationarity against the trend stationary *alternative* hypothesis for the real exchange rate, which may be consistent with a Balassa-Samuelson type model of PPP. Put differently, deviations of the real exchange rate off the *deterministic trend* are short-lived under the alternative hypothesis. This can happen if productivity factors grow deterministically in a Balassa-Samuelson type model of exchange rates.[2] However, it should be noted that the Balassa-Samuelson model can imply nonstationarity of the real exchange rate if productivity factors grow stochastically, which is against PPP.

Second, they investigate possibility of non-linear mean reversion process, employing a non-linear ADF test based on an exponential smooth transition autoregressive (ESTAR) model

[1] Casselian PPP implies that the real exchange rate hovers around the long-run equilibrium rate (with no deterministic time trend) and deviations from the equilibrium exchange rate die out eventually.

[2] See Mark (2001) for details.

(Kapetanios *et al.,* 2003). Recent theoretical and empirical studies on the real exchange rate have demonstrated the importance of non-linear adjustment dynamics in the real exchange rate. See, among others, Dumas (1992), Sercu *et al.* (1995), Michael *et al.* (1997), Obstfeld and Taylor (1997), and Kilian and Taylor (2003). Taylor (2001) also show that a failure to account for such nonlinearity may result in puzzles that underlie the difficulties in understanding real exchange rates dynamics.[3] It should be also noted that conventional linear unit root test tend to have low power problem when the true data generating process is nonlinear mean-reverting process.[4] That is, nonlinear models may improve the performance of conventional linear unit root tests for PPP and may provide explanations on why deviations from the long-run real exchange rate appear to be nonstationary. See, among others, Crucini and Shintani (2007).

With such motivations, Chortareas and Kapetanios (2004) applied Kapetanios *et al.*'s (2003) ESTAR unit root test to *detrended* yen-based real exchange rates for the other G7 and Asian/Pacific rim currencies, finding strong evidence of nonlinear mean reversion processes, that is, *trend* stationarity of the yen real exchange rate. Based on these findings, they conclude that the inability of rejecting the unit root null hypothesis for the yen real exchange rate may be due to the low power of linear unit root tests, thus previous findings do not reflect the failure of PPP.

The present paper casts doubt on the robustness of their findings. Our nonlinear unit root test for 13 G7 and Asian currencies relative to the yen hardly provide evidence in favor of stationarity when we extend the data until 2008. We check the evidence of Casselian PPP (with an intercept) as well as the Balassa-Samuelson type PPP (with deterministic trend). Kapetanios *et al.*'s (2003) test rejects the null of nonstationarity for a maximum 4 out of 13 countries in an array of tests, which hardly provides strong evidence of PPP.

Finding very weak evidence of stationarity for the yen real exchange rate, we employ a new nonlinear unit root test, the inf-t test, proposed by Park and Shintani (2005, 2010). The inf-t test is superior than previously proposed other nonlinear unit root tests in various aspects. For instance, the inf-t test does not need any Taylor approximation to deal with the so-called "Davies

[3] For example, half-life estimates based on linear models of real exchange rate tend to be biased *upward* when the true data generating process is a nonlinear stationary process.

[4] For instance, Pippenger and Goering (1993) report that linear unit root tests perform poorly when the true data generating process is the threshold autoregressive (TAR) model, and are sensitive to the speed of adjustment as well as location of the threshold parameter. Taylor *et al.* (2001) show with Monte Carlo simulations that the Dicky-Fuller test has low power against exponential smooth transition autoregressive (ESTAR) process.

problem," and requires much less stringent assumptions on the parameter space compared with other recently proposed tests. Considering three types of transition functions: ESTAR, band logistic smooth transition autoregression (BLSTAR), and band threshold autoregression (BTAR), the inf-t test did not reject the null of unit root for a maximum of 12 out of 13 currencies with standard lag selection procedures. That is, our results confirm empirical findings of previous researches. In what follows, we also show that our results confirm the findings of Kim and Moh (2010) that the use of ESTAR models may result in a misspecification problem that may not be detected when one uses Taylor approximation based tests such as the test by Kapetanios *et al.* (2003).

The remainder of the paper is organized as follows. Section 2 briefly describes Park and Shintani's (2005, 2010) inf-t test. In Section 3, we describe the three transition functions we employ in this paper. In Section 4, we provide a brief data description and report some pre-test results. Then, we report our main empirical results. Section 5 concludes.

2 The inf-t Test

The ESTAR unit root test by Kapetanios *et al.* (2003) has been popularly used in the current literature. So we provide a short description only on the inf-t test by Park and Shintani (2005, 2010). Consider a state-dependent autoregressive process model for a variable q_t, where the transition occurs between the following two regimes: the unit root regime,

$$\Delta q_t = u_t \tag{1}$$

and the stationary regime,

$$\Delta q_t = \lambda q_{t-1} + u_t, \tag{2}$$

where $\lambda < 0$ and u_t is the zero mean sequence of possibly serially correlated errors. Defining the transition function $\pi(q_{t-d}|\theta)$ as a weight on the stationary regime, the stochastic process of q_t can be represented by

$$\Delta q_t = \lambda q_{t-1}\pi(q_{t-d}|\theta) + u_t, \tag{3}$$

where q_{t-d} is the potentially *nonstationary* transition variable with delay lag $d \geq 1$.It should be noted that this is a very attractive property of the inf-t test. Many other nonlinear unit root tests such as the one by Caner and Hansen (2001) requires stationary transition variables, which can be a quite stringent requirement in practice. θ is an m-dimensional vector of parameters that can be identified only in the stationary regime and $\pi(\cdot)$ denotes a real-valued transition function on $(m+1)$-dimensional real space. Serial correlation in u_t can be accommodated as usual by adding lagged dependent variables in the right hand side of (3),

$$\Delta q_t = \lambda q_{t-1}\pi(q_{t-d}|\theta) + \sum_{j=1}^{k} \beta_j \Delta q_{t-j} + \varepsilon_t, \tag{4}$$

where ε_t is a martingale difference sequence that generates u_t.[5]

When $\lambda = 0$, the stochastic process of q_t is governed entirely by the unit root regime. Therefore, one may test the null of the unit root hypothesis,

$$H_0 : \lambda = 0$$

against the alternative hypothesis

$$H_1 : \lambda < 0,$$

which would imply that q_t obeys a nonlinear mean-reverting process.

One may implement the test as follows. Let Θ_n denote a random sequence of parameter spaces given for each n as functions of the sample $(q_1, ..., q_n)$. For each $\theta \in \Theta_n$, one obtains the t-statistic for λ in (4),

$$T_n(\theta) = \frac{\hat{\lambda}_n(\theta)}{s(\hat{\lambda}_n(\theta))}, \tag{5}$$

where $\hat{\lambda}_n(\theta)$ is the least squares estimate and $s(\hat{\lambda}_n(\theta))$ is the corresponding standard error. The inf-t test is then defined as

$$T_n = \inf_{\theta \in \boldsymbol{\Theta}_n} T_n(\theta), \tag{6}$$

which is the infimum of t-ratios in (5) taken over all possible values of $\theta \in \Theta_n$. The limit

[5] Park and Shintani (2005, 2010) assume that lagged differenced terms are not state-dependent, even though the test can be modified to allow that. See their papers for details.

distribution of inf-t statistic is free from any nuisance parameters and depends only on the transition function and the limit parameter space. The test can apply to a wide array of nonlinear partial adjustment AR models by employing a broad choice of the transition function $\pi(\cdot)$ as will be discussed in the next section.

3 The Nonlinear Models of the Real Exchange Rate

Let p_t be the natural logarithm of the price level in the home country, p_t^* be the log foreign price level, and e_t be the log nominal exchange rate as the unit price of the foreign currency in terms of the home currency. The log real exchange rate q_t is then defined as $e_t + p_t^* - p_t$. The present paper considers the following three nonlinear stationarity alternatives for the real exchange rate (q_t) including ESTAR, BLSTAR, and BTAR models described in (7) – (9), respectively.

$$\Delta q_t = \lambda(q_{t-1} - \mu)\left[1 - \exp\left\{-\kappa^2 \left(q_{t-1} - \mu\right)^2\right\}\right] + \sum_{i=1}^{k} \beta_i \Delta q_{t-i} + \varepsilon_t \tag{7}$$

$$\Delta q_t = \lambda \left[\frac{q_{t-1} - \tau_1}{1 + \exp\left\{\kappa\left(q_{t-1} - \tau_1\right)\right\}} + \frac{q_{t-1} - \tau_2}{1 + \exp\left\{-\kappa\left(q_{t-1} - \tau_2\right)\right\}}\right] + \sum_{i=1}^{k} \beta_i \Delta q_{t-i} + \varepsilon_t \tag{8}$$

$$\Delta q_t = \lambda\left[(q_{t-1} - \tau_1)\mathrm{I}\left\{q_{t-1} \leq \tau_1\right\} + (q_{t-1} - \tau_2)\mathrm{I}\left\{q_{t-1} \geq \tau_2\right\}\right] + \sum_{i=1}^{k} \beta_i \Delta q_{t-i} + \varepsilon_t, \tag{9}$$

where μ, τ_1, and τ_2 are the location and threshold parameters and κ is the scale parameter.

All regression equations include an intercept, which is appropriate for conventional Casselian view PPP. That is, these transition functions are considered to properly model the commodity arbitrage view of PPP with fixed transaction cost. Putting it differently, the real exchange rate may follow a unit root process *locally* around the long-run equilibrium PPP. Such a property may be well captured by ESTAR models. The BLSTAR and BTAR models can further allow an *inaction* band ($[\tau_1, \tau_2]$). In other words, when a deviation of the real exchange rate is not big enough, q_t follows a unit root process inside the inaction band. Note also that for a very high value for κ, the smooth transition function collapses to a discrete transition function. For instance, the BLSTAR model becomes the BTAR model in such a case because transition occurs abruptly when κ is sufficiently large.

For the scale parameter κ, we implement grid search for (6) over the parameter space given

$$[10^{-1}P_n, 10^3 P_n], \tag{10}$$

where $P_n = \left(\sum_{t=1}^{n} q_t^2/n\right)^{-1/2}$ as recommended by van Dijk *et al.* (2002). For the location parameter μ, we choose the interval

$$[\Psi_{n,15}, \Psi_{n,85}], \tag{11}$$

where $\Psi_{n,p}$ denotes the pth percentile of $(q_1, q_2, \cdots, q_n)$ as suggested by Caner and Hansen (2001). For the BLSTAR model, we grid search over the 2-dimensional parameter space of (κ, μ) spanned by (10) and (11).

4 Empirical Results

We consider 13 CPI based yen real exchange rates against other G7 and Asian/Pacific rim currencies from the *International Financial Statistics* CD-ROM. Other G7 countries include Canada, France, Germany, Italy, the UK, and the US. And 7 Asian and Pacific rim countries are Australia, Indonesia, Korea, Malaysia, the Philippines, Singapore, and Thailand. Observations are quarterly and span from 1973 through 1998 for the Euro-zone countries and through 2008 for the rest. We employ the General-to-Specific (GTS) rule for the linear model as recommended by Ng and Perron (2001) in selecting the number of lags (k). For nonlinear models (7) through (9), we employ the Partial Autocorrelation rule (PAR) following Granger and Teräsvirta's (1993) suggestion for the state-dependent autoregressive models. We choose a conventional value for the delay parameter, $d = 1$.

We start with the linear augmented Dickey-Fuller (ADF) test for the yen real exchange rates. Results are reported in Table 1. The test cannot reject the null of unit root for all 13 countries at the 5% significance level when an intercept is included. That is, we find no evidence of Casselian PPP when the linear ADF test is used. The ADF test with time trend rejects the null of nonstationarity for only one (Italy) out of 13 countries at the 5% significance level, which again implies extremely weak evidence in favor of a Balassa-Samuelson type PPP model.

Table 1

Next, we applied the ESTAR unit root test by Kapetanios *et al.* (2003), one of the most widely used nonlinear unit root tests, to the same 13 currencies relative to yens. We use three specifications for each test, one with no serial correlation ($k = 0$) and others that account for serial correlation ($k = 1$ and 2).[6] Results are reported in Table 2.

When demeaned series (Casselian view of PPP) are used, the test always rejects the null of a unit root for no other G7 countries at the 5% significance level. For Asian and Pacific rim countries, the test rejects the null only for Korea at the 1% significance level for all specifications and Indonesia at the 5% significance level when serial correlated errors are allowed. When detrended series (Balassa-Samuelson PPP) are used, the test rejects a minimum 1 and a maximum 4 out of 13 currencies at the 5% significance level. That is, though nonlinear adjustments for the yen real exchange rate are allowed, we still were not able to find reasonably strong evidence for PPP even with a weaker specification PPP.

Our results sharply contrast to the findings of Chortareas and Kapetanios (2004) who report strong evidence of the *trend* stationarity for the yen real exchange rate. This difference suggests that stationarity of yen real exchanges may be quite sensitive to the data points and one should interpret the test results carefully.

Furthermore, it should be noted that the ESTAR-ADF test of Kapetanios *et al.* (2003) requires the Taylor-approximation to avoid "Davies problem." Since it computes the test statistics without directly estimating key parameters, for instance, the error-correction coefficient, it is very difficult to detect potentially serious misspecification problems. To see whether this can be a serious problem, we apply Park and Shintani's (2005, 2010) inf-t test to our yen real exchange rates.[7]

Table 2

[6] $k = 1$ is selected by the PAR. We find very similar results when k is extended to 3.

[7] Park and Shintani's (2005, 2010) test does not consider the case in presence of time trend. That is, their test is able to test conventional Casselian PPP rather than the Balassa-Samuelson PPP.

We conduct the inf-t test for the three nonlinear AR models (7) – (9) and results are presented in Tables 3 through 5. As mentioned before, one clear advantage of using Park and Shintani's (2005, 2010) inf-t test over the Taylor-approximation based test is that it directly estimates all parameters in the model, thus can provide useful information on misspecification problems. Our inf-t test results with the ESTAR model clearly demonstrates that this may be the case (see Table 3). The inf-t test rejects the unit root null for Korea at the 1% significance level, which is roughly consistent with the results in Table 2. It should be noted, however, that the λ estimate is by far less than -2. Since $k = 0$ for Korea, this implies that the real exchange rate is cyclically explosive, inconsistent with stationarity. This implies that ESTAR models may not be appropriate for the data.

Table 3

Next, we implement the inf-t test with the BLSTAR specification and results are reported in Table 4. The test rejects the null of unit root only for Korea favoring the nonlinear stationarity alternative. One interesting finding is that the estimate for κ for Korea is still large (15.099), which implies that the yen/Korean won real exchange rate can be successfully approximated by the BTAR model. Our test with the BTAR specification (Table 5) reveals that this is indeed the case. We find quite similar values for λ and τ's as well as the inf-t statistics for Korea with the BTAR and BLSTAR specifications. This is not surprising, because the BLSTAR collapses to the BTAR process as κ increases to infinity. In a nutshell, we find very weak evidence of nonlinear stationarity for yen real exchange rates.

Table 4

Table 5

Recent studies on real exchange rate dynamics suggest that nonlinear models can provide an explanation for the poor performance of conventional linear unit-root tests and extremely slowly mean-reverting property (PPP puzzle) of the real exchange rates. Lothian (1990) reports empirical evidence in favor of PPP for yen real exchange rates using long-horizon data. And he also points out that the mean reversion process may not occur continuously, depending on the current state of the real exchange rate. Lothian's (1990) findings motivated Chortareas and Kapetanios (2004) to use a nonlinear model. However, as can be seen from our applications, one has to interpret empirical results if Taylor-approximation based nonlinear tests are used, because such tests may not be able to detect misspecification problems.

5 Concluding Remarks

Despite of popularity and volume of studies devoted, empirical evidence for PPP is mixed at best. However, the profession seems to find less ambiguity in the nonstationarity of the yen real exchange rate as they often fail to obtain the evidence of mean-reversion. Chortareas and Kapetanios (2004) report empirical findings that may suggest possible nonlinear adjustment of yen real exchange rates toward its *deterministic trend* in the long-run. We reexamine this issue by first testing the null of a unit root against the same alternative hypothesis as theirs but with extended set of observations. We find very weak evidence of nonlinear stationarity both with an intercept (Casselian view PPP) and with time trend (Balassa-Samuelson PPP). That is, Chortareas and Kapetanios' findings (2004) lack robustness and are easily upset when we extended the sample period.

We also investigate consequences of using a Taylor-approximation based nonlinear unit root test such as the one Chortareas and Kapetanios (2004) used. For this purpose, we employ a more rigorous nonlinear unit root test by Park and Shintani (2005, 2010) for an array of transition functions, the ESTAR, BLSTAR, and BTAR. We apply the inf-t test to 13 yen real exchange rates for G7 and Asian/Pacific rim countries. The test rejects the null of unit root only for yen/Korean won out of those 13 currencies. That is, allowing nonlinear adjustment to the yen real exchange rates fail to obtain reasonably strong evidence of PPP for yen real exchange rate, when a more rigorous nonlinear test is used.

Recently, Kim and Moh (2010) report some empirical evidence against the use of Taylor-approximation based ESTAR tests such as the one by Kapetanios *et al.* (2003), which may be unable to detect misspecification problems. Our results are consistent with the work of Kim and

Moh (2010).

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Table 1. Unit Root Test Results: Linear Model

$$\Delta q_t = \lambda q_{t-1} + \sum_{i=1}^{k} \beta_i \Delta q_{t-i} + \varepsilon_t$$

G7 Currencies		
Country	ADF$_c$	ADF$_t$
Canada	-2.071	-1.347
France	-1.313	-3.117
Germany	-1.065	-2.084
Italy	-1.812	-3.917*
UK	-2.044	-1.881
US	-2.431	-2.248

Asian and Pacific Rim Currencies		
Country	ADF$_c$	ADF$_t$
Australia	-1.723	-0.739
Indonesia	-1.419	-1.421
Korea	-2.191	-2.120
Malaysia	-1.860	-2.129
Philippines	-2.007	-1.651
Singapore	-2.905	-2.843
Thailand	-1.600	-1.845

Notes: i) The number of lags was chosen by the General-to-Specific rule (Hall, 1994) following Ng and Perron (2001). ii) ADF$_c$ and ADF$_t$ denotes the augmented Dickey-Fuller statistics with demeaned data and with demeaned and detrended data, respectively. iii) * and ** refer to the cases when the unit root null is rejected at the 5% and 1% significance levels, respectively. iv) The asymptotic critical values were obtained from Harris (1992).

Table 2. Unit Root Test Results: Taylor-Approximation Based Exponential Smooth Transition Autoregressive Model by Kapetanios *et al.* (2003)

$$\Delta q_t = \delta q_{t-1}^3 + \sum_{i=1}^k \beta_i \Delta q_{t-i} + \varepsilon_t$$

			G7 Currencies			
Country	$\text{NLADF}_c^{k=0}$	$\text{NLADF}_c^{k=1}$	$\text{NLADF}_c^{k=2}$	$\text{NLADF}_t^{k=0}$	$\text{NLADF}_t^{k=1}$	$\text{NLADF}_t^{k=2}$
Canada	-2.002	-2.122	-2.099	-0.715	-1.061	-0.977
France	-1.983	-2.625	-2.543	-3.439*	-4.646**	-5.379**
Germany	-1.865	-2.577	-2.449	-3.172	-4.091**	-4.160**
Italy	-1.987	-2.478	-2.702	-2.700	-3.692*	-4.114**
UK	-1.466	-2.149	-2.096	-1.225	-1.913	-1.881
US	-2.292	-2.480	-2.338	-2.292	-2.613	-2.427

			Asian and Pacific Rim Currencies			
Country	$\text{NLADF}_c^{k=0}$	$\text{NLADF}_c^{k=1}$	$\text{NLADF}_c^{k=2}$	$\text{NLADF}_t^{k=0}$	$\text{NLADF}_t^{k=1}$	$\text{NLADF}_t^{k=2}$
Australia	-2.384	-2.513	-2.710	-0.909	-1.163	-1.115
Indonesia	-2.439	-2.661	-3.036*	-3.148	-3.762*	-4.750**
Korea	-4.573**	-4.786**	-4.823**	-2.714	-2.410	-2.368
Malaysia	-1.524	-1.766	-1.932	-0.995	-1.404	-1.386
Philippines	-2.161	-2.289	-2.458	-1.485	-1.682	-1.883
Singapore	-1.280	-1.702	-1.934	-1.290	-1.900	-1.971
Thailand	-2.075	-2.317	-2.168	-1.287	-1.658	-1.255

Notes: i) NLADF denotes the t-statistic for δ as described in Kapetanios *et al.* (2003). ii) ADF_c and ADF_t denotes the augmented Dickey-Fuller statistics with demeaned data and with demeaned and detrended data, respectively. iii) * and ** refer to the cases when the unit root null is rejected at the 5% and 1% significance levels, respectively. iv) The asymptotic critical values were obtained from Kapetanios *et al.* (2003). Simulated critical values with actual sample sizes yielded same conclusions.

Table 3. Unit Root Test Results: Exponential Smooth Transition Autoregressive Model

$$\Delta q_t = \lambda(q_{t-1} - \mu)\left[1 - \exp\left\{-\kappa^2 (q_{t-1} - \mu)^2\right\}\right] + \sum_{i=1}^{k} \beta_i \Delta q_{t-i} + \varepsilon_t$$

G7 Countries					
Country	k	inf$-t$	λ	κ	μ
Canada	0	-2.016	-476.422	0.021	4.662
France	0	-2.026	-543.473	0.031	3.196
Germany	0	-2.024	-1075.274	0.023	4.413
Italy	0	-2.384	-390.247	0.039	-2.637
UK	1	-2.133	-1219.412	0.019	5.277
US	3	-2.512	-1337.517	0.021	4.823

Asian and Pacific Rim Currencies					
Country	k	inf$-t$	λ	κ	μ
Australia	0	-2.588	-502.025	0.022	4.631
Indonesia	0	-2.720	-55.552	0.028	-3.367
Korea	0	-4.602**	-426.418	0.048	-2.037
Malaysia	0	-1.980	-0.020	16.030	3.556
Philippines	0	-2.430	-57.596	0.078	1.302
Singapore	1	-2.327	-0.047	5.524	4.383
Thailand	0	-2.370	-47.418	0.068	1.497

Notes: i) The number of lags (k) was chosen by the Partial Autocorrelation rule following Granger and Teräsvirta (1993). ii) $\exp\{\cdot\}$ is an exponential function.iii) * and ** refer to the cases when the unit root null is rejected at the 5% and 1% significance levels, respectively. iv) The asymptotic critical values were obtained from Park and Shintani (2009).

Table 4. Unit Root Test Results: Band Logistic Smooth Transition Autoregressive Model

$$\Delta q_t = \lambda \left[\frac{q_{t-1}-\tau_1}{1+\exp\{\kappa(q_{t-1}-\tau_1)\}} + \frac{q_{t-1}-\tau_2}{1+\exp\{-\kappa(q_{t-1}-\tau_2)\}}\right] + \sum_{i=1}^{k}\beta_i \Delta q_{t-i} + \varepsilon_t$$

G7 Countries						
Country	k	inf$-t$	λ	τ_1	τ_2	κ
Canada	0	-2.026	-0.080	4.442	4.824	1770.67
France	0	-2.015	-0.272	2.966	3.444	6.533
Germany	0	-2.159	-0.290	4.203	4.666	9.225
Italy	0	-2.274	-0.102	-2.727	-2.546	6.390
UK	1	-2.118	-0.112	5.107	5.466	4.070
US	3	-2.432	-0.187	4.633	5.056	4.722

Asian and Pacific Rim Currencies						
Country	k	inf$-t$	λ	τ_1	τ_2	κ
Australia	0	-2.544	-0.171	4.321	4.881	7.474
Indonesia	0	-2.704	-0.203	-4.067	-2.687	3.029
Korea	0	-5.339**	-0.718	-2.297	-1.792	15.099
Malaysia	0	-1.946	-0.021	3.476	3.688	0.026
Philippines	0	-2.731	-0.333	1.002	1.624	12.025
Singapore	1	-2.325	-0.071	4.293	4.546	82.745
Thailand	0	-3.000	-0.351	1.117	1.848	15.915

Notes: i) The number of lags (k) was chosen by the Partial Autocorrelation rule following Granger and Teräsvirta (1993). ii) I$\{\cdot\}$ is an indicator function. iii) * and ** refer to the cases when the unit root null is rejected at the 5% and 1% significance levels, respectively. iv) The asymptotic critical values were obtained from Park and Shintani (2009).

Table 5. Unit Root Test Results: Band Threshold Autoregressive Model

$$\Delta q_t = \lambda \left[(q_{t-1} - \tau_1) \mathrm{I}\left\{ q_{t-1} \leq \tau_1 \right\} + (q_{t-1} - \tau_2) \mathrm{I}\left\{ q_{t-1} \geq \tau_2 \right\} \right] + \sum_{i=1}^{k} \beta_i \Delta q_{t-i} + \varepsilon_t$$

G7 Currencies					
Country	k	inf$-t$	λ	τ_1	τ_2
Canada	0	-2.393	-0.061	4.422	4.632
France	0	-2.202	-0.181	3.096	3.354
Germany	0	-2.443	-0.231	4.283	4.586
Italy	0	-2.349	-0.081	-2.677	-2.466
UK	1	-2.145	-0.072	5.097	5.426
US	3	-2.621	-0.077	4.703	5.016

Asian and Pacific Rim Currencies					
Country	k	inf$-t$	λ	τ_1	τ_2
Australia	0	-2.976	-0.174	4.401	4.751
Indonesia	0	-2.598	-0.154	-3.937	-2.927
Korea	0	-5.505**	-0.722	-2.257	-1.082
Malaysia	0	-2.337	-0.070	3.496	4.278
Philippines	0	-2.844	-0.281	1.072	1.534
Singapore	1	-2.837	-0.095	4.173	4.466
Thailand	0	-3.441*	-0.367	1.167	1.758

Notes: i) The number of lags (k) was chosen by the Partial Autocorrelation rule following Granger and Teräsvirta (1993). ii) $\mathrm{I}\{\cdot\}$ is an indicator function. iii) * and ** refer to the cases when the unit root null is rejected at the 5% and 1% significance levels, respectively. iv) The asymptotic critical values were obtained from Park and Shintani (2009).