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Testing for non-stationary hypotheses against local and global non-linear stationary alternatives

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ABSTRACT

This paper develops a sequential procedure for testing non-stationary hypotheses against local and global non-linear stationary alternatives. This involves: (i) developing a test against a local non-linear ESTAR stationary alternative, and (ii) developing a test against a global ESTAR non-linear stationary alternative process. The null hypothesis of non-stationary is tested against the local ESTAR non-linear stationary alternative; if the null is rejected, then testing against the global ESTAR non-linear stationary alternative process follows. These tests were applied to Indonesian real exchange rates, against its major trading partners - U.S, Japan and Singapore. The Indonesian government occasionally intervened in the foreign exchange markets, and consequently, there were notable structural breaks in real exchanges rates to be studied in this paper. This paper integrates three ideas - local stationarity, global stationarity and structural breaks - in developing tests to establish the stationary properties of real exchange rates. Application of these tests provides concrete evidence in establishing the global non-linear stationarity of all three real exchange rates series. These tests will have much wider application than those provided in this paper.

Keywords: ESTAR model, local stationarity, nuisance parameters, global stationarity, structural breaks, real exchange rates

JEL classification: F32, F31, C22, C5

1. Introduction

The standard linear autoregressive moving average (ARMA) framework has long been used to establish whether or not time series are stationary. However, there is a growing dissatisfaction among applied researchers regarding the use of this framework to investigate the stationary behaviour of economic and financial time series. Much of this discontent is due to theoretical predictions of stationarity of some variables - in several areas of economics such as international monetary economics and macro-finance - not being supported by ADF-type tests. For example, the finding that real exchange rates have unit roots caused apprehension among economists, wanting to use the long run purchasing power parity (PPP) equilibrium relationship in exchange rate determination. Another notable example is the finding of unit roots in real interest rates, and thus several economists argued that theories related to real business cycles and modern economic growth are hard to retain. Recently, in order to resolve this puzzle, numerous researchers developed tests that are more powerful than the ADF test in rejecting the unit root hypothesis. One strand of literature focused on the use of panel data and rejected the hypothesis of joint unit roots in a group of real exchange rates against the stationary alternative hypothesis. See, for example, Im, Pesaran and Shin (2002). Another strand of literature focused on fractional integration (Cheung and Lai, 1993) to establish the stationarity of time series.

Recently, in testing for the stationarity of economic and financial variables, several empirical studies exploited the fact that many time series are largely non-linear and that the ADF-type unit roots do not have power against these models. See, for example, Killian and Taylor (2001) and Taylor et al. (2002). Further, economies generally go through various phases of business cycles, and hence many economic and financial variables exhibit deepness and steepness asymmetries. Using threshold autoregressive (TAR) models to capture these asymmetries, Enders and Granger (1998) developed procedures to test against alternatives of asymmetric stationary roots. In these models, the regime switching occurs abruptly, and the studies that used these models to test the stationarity of real exchange rates have had limited success in supporting the stationary hypothesis. See, for example, Enders and Dibooglu (2001), and Lestrari, Silvapulle and

Kim (2003). Moreover, several theoretical and empirical economists have argued that there are cases in which the transformation or switching occurs smoothly rather than abruptly as occurs in TAR models, and thus the application of smooth transition autoregressive (STAR) framework that can capture such smooth switching from one regime to other has become very popular. It has been argued in the recent literature that the lack of empirical evidence supporting PPP appears to be caused by the breakdown in the assumptions (under which the PPP is valid) due to factors such as transaction costs, taxation, imperfect competition, foreign exchange market interventions, and the differential composition of commodity baskets and hence the price indices across countries. Subsequently, an alternative framework for the empirical analysis of real exchange rates, which allows for friction in the commodity trade, has emerged. Dumas (1992) and Sercu, Uppal and Van Hulle (1995) developed equilibrium models of exchange rate determination in the presence of transaction costs, hence providing empirical evidence of real exchange rate adjustment being a non-linear ESTAR mean-reverting process. See section 3.1 for more details.

The objective of this paper is to develop a new sequential procedure for testing non-stationarity hypotheses against local and global ESTAR non-linear stationary alternatives. These tests will be applied to Indonesian real exchange rates against its major trading partners - U.S., Japan and Singapore. The Indonesian government aggressive intervention in the foreign exchange markets has caused five notable structural breaks in the real exchanges rates in the sample period we chose to study. By incorporating dummy variables to capture these breaks, the proposed tests will be extended and applied to all three real exchange rates, thus establishing their global non-linear stationary properties.

Testing for a non-stationary process against the ESTAR stationary alternative is not new. Kapetanios et al. (2003) developed a test for the null hypothesis of non-stationary series against stationary non-linear alternatives, ignoring the problem of the presence of nuisance parameters under the null hypothesis of non-stationarity. See section 3.2 for more details. Further, the critical values tabulated in this paper are inappropriate to use in the empirical investigation to be carried out in this paper, due to the five structural breaks in Indonesian real exchange rates.

This paper proposes a test statistic for the null of non-stationarity against the local ESTAR stationary alternatives. In this testing, the smoothing and threshold parameters governing the ESTAR process become unidentified under the null. Earlier studies have linearised this process to overcome the problem of nuisance parameters, whereas our study takes the presence of such parameters into account in developing the testing procedure. Further, we also develop a testing procedure against the global ESTAR non-stationarity. Kapetanios et al. (2003) estimated the power of their proposed test against a global process, and this paper, on the other hand, will propose a test statistic against global stationary alternative hypothesis. Moreover, to date, tests developed to establish the stationarity properties of real exchange rates have been applied mainly to OCED countries, with exchange rates largely being market-determined. We found these tests to be inapplicable directly, for example, to some Asian real exchange rates, as aggressive government interventions caused significant impacts on these rates. In this study, we extend the proposed tests to accommodate the presence of structural breaks caused by such interventions.

This paper is organised as follows: Section 2 briefly outlines the ESTAR model. Section 3 outlines the model with linear and non-linear ESTAR components, justifying the use of such a model for capturing real exchange rate behaviour. Section 4 specifies the hypotheses related to local and global non-linear stationary processes and proposes test statistics for sequential testing of local and global ESTAR stationary alternatives. Section 4 discusses a simulation experiment conducted to estimate the critical values of the distributions, and assess the small sample properties of the proposed tests, followed by a discussion of their results. Section 5 describes the data series used in this study. Section 6 applies the tests proposed in this paper to Indonesian real exchange rates. Some concluding remarks are made in Section 7.

2. Exponential smooth transition autoregressive (ESTAR) models

We consider a STAR model of order p for variable y_t, given as:

$$y_t = \varphi_0 + \Phi' x_t + (\lambda_0 + \Theta' x_t) G(s_t;\gamma,c) + u_t . \qquad (1)$$

where $x_t = (y_{t-1}, y_{t-2}, ..., y_{t-p})'$, $\Phi = (\varphi_{1,}\varphi_2, ..., \varphi_p)'$ and $\Theta = (\theta_{1,}\theta_2, ..., \theta_p)'$ are unknown parameter vectors and $u_t \sim n.i.d(0,\sigma^2)$. $G(.)$ is a transition function which is continuous, and bounded by zero and one. s_t is a transition variable, c is the threshold parameter and γ is the smoothing parameter. The transition variable s_t may be a single stochastic variable, for example, an element of x_t, a linear combination of stochastic variables or an exogenous variable. In many empirical studies employing STAR-type models, the transition variable s_t is generally assumed to be a lagged endogenous variable, that is, $s_t = y_{t-d}$ for an integer d (delay parameter) > 0 (Teräsvirta, 1994 and Franses and Van Dijk, 2000). The STAR model can be interpreted as a regime switching model with two regimes, associated with two extreme values of the transition function, $G(s_t;\gamma,c)$=0 and $G(s_t;\gamma,c)$=1, with the transition from one regime to the other being gradual rather than abrupt as in TAR models.

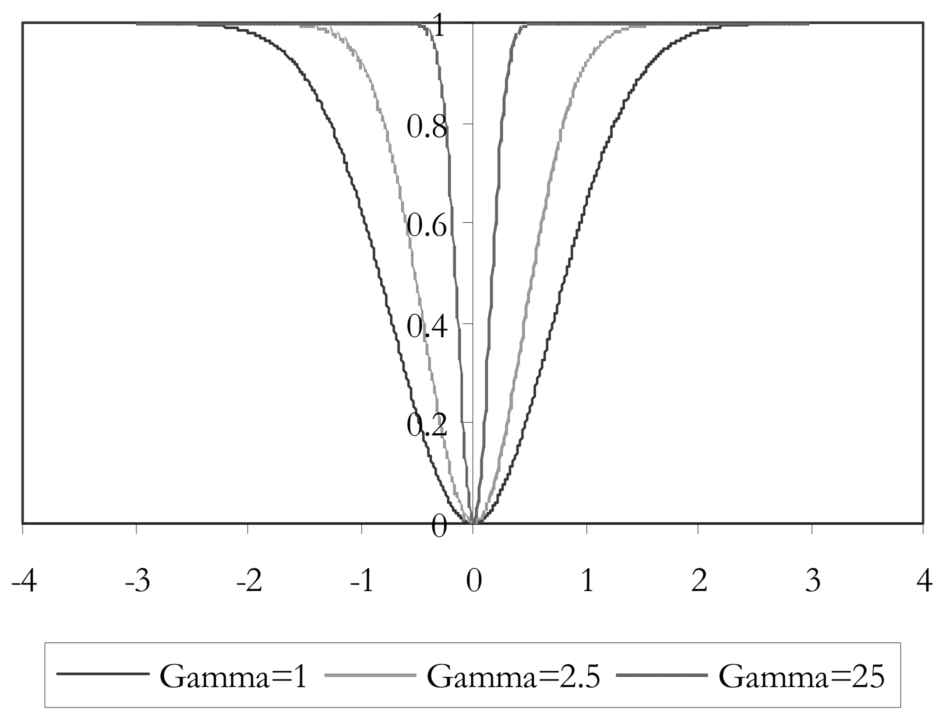
In this paper, the exponential smooth transition (EST) function, which is extensively being used in applied economic analysis of time series such as real exchange rates and real interest rates, will be studied. Its functional form is given as,

$$G(s_t;\gamma,c) = 1 - \exp\left[-\gamma(s_t - c)^2\right],\ \gamma > 0 \qquad (2)$$

where γ measures the speed of transition from one regime to another and c represents the location for threshold value for s_t. Equation (1) combined with (2) yields the Exponential STAR (ESTAR) model. The transition function of ESTAR is symmetric and U-shaped around c. Figure 1 shows the shapes of the transition function for some values of the smoothing parameter γ. It can be shown that as $\gamma \to 0$, the function

$G(.)$ goes to zero and (2) reduces to a linear model. On the other hand, as $\gamma \to \infty$, $G(.)$ converges to one and (2) reduces to a linear model with a different set of parameter values. It is noticeable in the graph presented in Figure 1 that the transition becomes faster as the values of γ increases, and for large γ, the transition from one regime to other is abrupt and the model becomes TAR.

Figure: Examples of the exponential function $G(s_t;\gamma,c)$ as given in (3) for various value of γ and threshold $c=0$



3. Modelling non-linear adjustments

We consider an ESTAR process of order *p* for y_t given as:

$$y_t = \sum_{j=1}^{p} \pi_j y_{t-j} + (\sum_{j=1}^{p} \pi_j^* y_{t-j}) \times \left\{1 - \exp\left[-\gamma\left(s_t - c\right)^2\right]\right\} + u_t \qquad (3)$$

where y_t is a stationary and ergodic process with zero mean, $u_t \sim n.i.d(0,\sigma^2)$, and $\gamma > 0$, which determines the speed of the transition process between two extreme regimes. The middle regime corresponds to $s_t = c$ and (3) becomes a linear $AR(p)$ model:

$$y_t = \pi y_{t-1} + \sum_{j=2}^{p} \pi_j y_{t-j} + u_t \tag{4}$$

The outer regime corresponds to $s_t = \pm\infty$, and (3) becomes the following linear $AR(p)$ model,

$$y_t = (\pi + \pi^*) y_{t-1} + \sum_{j=2}^{p} (\pi_j + \pi_j^*) y_{t-j} + u_t$$

with $\pi = \pi_1 \; and \; \pi^* = \pi_1^*$. Now, assuming that the higher order terms in AR(p) are zero, we re-write the model (3) as:

$$\Delta y_t = \delta y_{t-1} + \delta^* y_{t-1}(1 - \exp[(-\gamma (s_t - c)^2]) + u_t \tag{5}$$

where $\delta = 1 - \pi$ and $\delta^* = \pi^*$. The crucial parameters in (5) are δ and δ^* which determine whether or not the small and large deviations respectively are stationary processes. The presence of transaction costs suggests that there exists a region of no trading, and therefore, for small deviations from the equilibrium, the real exchange rate may be non-stationary, following either a random walk or an explosive process. Outside this region (that is, for large deviations), on the other hand, the possibility of international arbitrage can bring the series close to the equilibrium. See Killian and Taylor (2001) for an alternative explanation based on, among others, heterogenous traders and noise traders. For this reason, the small deviations are modelled as linear, whereas the large deviations as non-linear ESTAR, these being captured precisely by (5). While $\delta \geq 0$ is possible, the conditions that $\delta^* < 0$ and $\delta + \delta^* < 0$ must be satisfied for the process y_t to be globally stationary. See Michael, Nobay and Peel (1997) for more discussion on this. Using simulation studies, it has been shown that the ADF-type tests have no power against the non-linear ESTAR model. See, for example, Michael, Nobay and Peel (1997) and Taylor, Peel and Sarno (2001).

As has been argued before, we will focus on another issue related to testing for non-stationarity of Indonesian real exchange rates. From our experience in empirical applications, we find that the tests and the models discussed in this paper so far and elsewhere are not directly applicable to analyse Asian real exchange rates, because the operation of these financial markets is different from those of OECD markets. In the former, exchange rate movements are managed by government interventions in the foreign exchange markets frequently, while in the latter, they are largely market determined. These interventions can have significant impacts on real exchange rates, resulting in structural breaks in the series. In the presence of such breaks in the series, the test statistics to be developed in the context of model (5) will be extended by integrating three ideas - local and global stationary hypotheses and the presence of structural breaks. See section 3.2 for more details.

3.1. Hypotheses for testing against local and global non-linear stationary processes and the testing procedures

Recall the model (5),

$$\Delta y_t = \delta y_{t-1} + \delta^* y_{t-1}(1-\exp[-\gamma\left(s_t - c\right)^2]) + u_t$$

We will develop procedures for testing two sets of hypotheses: One is, setting $\delta = 0$, the null hypothesis $H_{0L}:\delta^* = 0$ is tested against the stationary alternative hypothesis $H_{1L}:\delta^* < 0$, which lends a methodology for testing against the local nonlinear stationary process. Second is, the null hypothesis $H_{0G}:\delta + \delta^* \geq 0$ is tested against $H_{1G}:\delta + \delta^* < 0$, which lends a methodology for testing against the global stationary process. In what follows, we shall propose the test statistics.

(i) Under H_{OL}, the parameters γ and c become unidentified. In the presence of such nuisance parameters, a t-test for $H_{0L}:\delta^* = 0$ against $H_{1L}:\delta^* < 0$ based on $\hat{t}_{\delta^*=0}(\gamma, c)$ - which is the conventional t-test but a function of nuisance parameters γ and c. We overcome this problem by exploiting Davies (1987) idea and computing the largest possible value of t-statistics over the space spanned by γ and c. The Davies approach is to define the test statistic as:

$\sup \hat{t}_{L,\delta^*=0} = -\sup \left| \hat{t}_{L,\delta^*=0}(\gamma, c) \right| = -\sup \left| \frac{\hat{\delta}^*}{s.e(\hat{\delta}^*)} \right|$, where $\hat{\delta}^*$ is a function of the parameter space spanned by (γ, c).

In such cases, simulation methods are generally used to compute the critical values of the distribution of the test statistic. To improve the power properties of the tests, parameter space needs to be restricted, and the previous empirical studies and stylised facts related to ESTAR model can provide some guidance towards this end.

(ii) In order to test $H_{0G} : \delta + \delta^* \geq 0$ against $H_{1G} : \delta + \delta^* < 0$, we compute the $F_{G,(1,T-k)}$ statistic, and hence the statistic $t_{G,\delta+\delta^*=0}$, and its estimate has a non-standard *t*-distribution and its critical values will be generated using simulations.

Recall that y_t in (5) was assumed to have zero mean. In order to accommodate stochastic processes x_t with nonzero mean and/or linear deterministic trend, the processes should be demeaned as $x_t = \mu + y_t$ and/or detrended as $x_t = \mu + \beta t + y_t$, and the resultant residuals y_t be used in ESTAR model (5). Further, as has been argued before, there are five structural breaks present in the Indonesian real exchange rates, over the sample period to be studied in this paper. The dummy variables defined in (6) below will capture these breaks. Clearly, in the presence of structural breaks, the distribution of test statistics - $\sup \hat{t}_{L,\delta^*=0}$ and $t_{G,\delta+\delta^*=0}$ - will not be the same as for the demeaned and/or detrended series.

$$D_1 = \begin{cases} 1 & if \quad t < T_{B1} (T_{B1} = 1983:M4) \\ 0 & otherwise \end{cases}$$

$$D_2 = \begin{cases} 1 & if \quad T_{B1} \le t < T_{B2} (T_{B2} = 1986:M9) \\ 0 & otherwise \end{cases} \qquad (6)$$

$$D_3 = \begin{cases} 1 & if \quad T_{B2} \le t < T_{B3} (T_{B3} = 1997:M8) \\ 0 & otherwise \end{cases}$$

$$D_4 = \begin{cases} 1 & if \quad T_{B3} \le t < T_{B4} (T_{B4} = 1998:M9) \\ 0 & otherwise \end{cases}$$

$$D_5 = \begin{cases} 1 & if \quad t > T_{B4} \\ 0 & otherwise \end{cases}$$

In the presence of these breaks in the series, the appropriate critical values of the test statistics will be generated via simulations. The dummy-filtered series y_t is generated as follows:

$$x_t = \mu + \beta t + \sum_{i=1}^{5} \phi_i D_i + y_t$$

and used in model (5). In order to establish whether y_t is non-stationary or global non-linear ESTAR stationary, a sequential testing is done as follows:

(i) Test the null $\mathrm{H}_{0L}: \delta^* = 0$ against $\mathrm{H}_{1L}: \delta^* < 0$ using the $\sup t_{L,\delta^*=0}$ - statistic for which the critical values are tabulated in Table 1.

If the null is not rejected, conclude that the series under investigation is non-stationary. On the other hand, if the null is rejected, then conclude that the series is locally non-linear stationary, and proceed to test the following:

(ii) Test $H_{0G}: \delta + \delta^* \ge 0$ against $H_{1G}: \delta + \delta^* < 0$ using the one-sided $t_{G,\delta+\delta^*=0}$ statistic for which the critical values are tabulated in Table 2. According to the results of a limited simulation study carried out in this paper, this test appears to be "similar".

If the null is not rejected conclude that the series is non-stationary, and on the other hand, if the null is rejected conclude that the series is globally ESTAR non-linear stationary process.

On the selection of transition variable s_t: our experience showed that Δy_{t-d} is a more appropriate proxy for s_t, than y_{t-d} used by Kapetanios et al. (2003). However, to simplify the simulation study we assume that d = 1. In fact, d can be chosen from the set $\{1, 2, \ldots, d^{max}\}$ based on an information criterion such as AIC or a sequence of hypotheses tests. We do not also take the presence of serial correlation coefficient into account in developing these tests. We believe this issue can be handled as that in applying the ADF test.

4. A simulation study and the results

The simulation study is conducted in a number of stages:

(i) Calculate 1, 5 and 10 per cent critical values of the distribution of the test statistic $\sup \hat{t}_{L,\delta^*=0}$ for the models A, B and C. Since this statistic is a function of γ and c, it has been computed over the parameter space $\gamma = [0.1, 5.0]$, $c = [c^L, c^U]$, where c^L and c^U are the minimum and maximum values of sorted Δy_{t-1}, with 10 per cent of the top and the bottom of these values trimmed.

(ii) Calculate 1, 5 and 10 per cent critical values of the distribution of the test statistic $\sup \hat{t}_{L,\delta^*=0}$ over the parameter space defined in (i) above for the three models D, E and F, with five dummy variables present.

The stages (iii) and (iv) involve estimating the critical values of the distribution of $t_{G,\delta+\delta^*=0}$ statistic without and with the structural breaks present in the variable x_t. Further, in this section, we will outline the experimental design used in the simulation study and discuss the results.

4.1 Experimental Design

The critical values were computed using 20,000 simulations for the sample of sizes T=100, 300 and 1000. The sizes and the powers (the rejection frequencies under the null and alternative hypotheses respectively) were estimated using 2000 and 1000 simulations respectively, only for the sample of sizes T=100 and 300. Let x_t be the time series to be studied, this series is filtered for various deterministic variables using the following models:

$$x_t = \mu + \beta t + \sum_{i=1}^{5} \phi_i D_i + y_t \quad (7)$$

where μ is the mean, t is the time trend and D_i, $i=1, \ldots, 5$ are the dummy variables, capturing the structural breaks noted in Indonesian real exchange rate series.

Model A: Model (7) with $\mu = \beta = \phi_i (i = 1,..,5) = 0$

Model B: Model (7) with $\mu \neq 0$ & $\beta = \phi_i (i = 1,..,5) = 0$

Model C: Model (7) with $\mu \neq 0, \beta \neq 0$ & $\phi_i (i = 1,..,5) = 0$

The models D, E and F are the same as those A, B and C respectively, with $\phi_i \neq 0$ *for* $i = 1,..,5$.

The filtered series y_t is used in the model (5). The critical values for the stages (i) and (ii) above are reported in Table 1, while those for the stages (iii) and (iv) are reported in Table 2. These critical values were used to compute the rejection probabilities under the null (sizes) and under the alternatives (powers) of the tests proposed in this paper. The selected parameter sets for these calculations are defined as: δ = *{0.1 0.2 0.3}*, δ^* *={ -0.3 -0.5 –1.0}*, γ *={ 0.1 0.6 1.2}* and *c ={-0.3 0.0 0.3}*.

4.2 Simulation results

Table 1 reports the 1, 5 and 10 per cent critical values of the distribution of the $\sup \hat{t}_{L,\delta^*=0}$ statistic for models A, B and C for the sample sizes T = 100, 300 and 1000. The distribution of this statistic appears to be much wider for T = 100 than for the other two sample sizes. The tabulated critical values for T=300 do not notably differ from those for T=1000, and therefore, we can say that the sample size 300 is large enough for the asymptotic results to hold. Clearly, the distributions become wider for the models filtered with deterministic variables, in particular with five dummy variables, than those for the models filtered only with mean and/or deterministic trend. The critical values of the $t_{G,\delta+\delta^*=0}$-statistic were tabulated in Table 2. Although the asymptotic results are not yet derived theoretically, our experience with the simulation experiment conducted in this study shows that the distribution of $\hat{t}_{G,\delta+\delta^*=0}$- statistic, to some extent, depends on the parameter values of γ and c.

The rejection rates (sizes) of these tests under the null hypotheses are reasonably close to the selected nominal level of five per cent. The rejection rates (powers) of these tests under the corresponding alternative hypotheses are always higher than the sizes of the tests, and they tend to increase as the parameter values and/or the sample size increase as expected.

5. Data Series

The bilateral exchange rates between Indonesian Rupiah against U.S-Dollar, Japanese Yen, and Singaporean Dollar were collected from the IMF's International Financial Statistics CD-ROM for the period January 1979 to June 2003. The CPI series were also collected from IMF International Financial Statistics CD-ROM. The domestic price index is the Indonesian CPI and foreign price indices are the U.S, Japan and Singapore CPI series. The relative price is defined as the ratio of domestic price to foreign price. The CPI and nominal exchange rate series are used in natural logarithms.

Indonesia had a fixed nominal exchange rate regime until October 1978. In November 1978, with 48 percent rupiah devaluation, the Indonesian government changed the exchange rate regime to a managed-floating one with infrequent large devaluations, and concurrently pegged the rupiah to a basket of currencies of Indonesia's major trading partners. A 28 per cent exchange rate devaluation was experienced by Indonesian economy in March 1983, followed by a 31 per cent exchange rate devaluation in September 1986. The managed-floating system was sustained until a major financial crisis in July 1997. Following this crisis, the Indonesian rupiah was floated freely from August 1997 onwards and the government intervention ceased.

The three detrended real exchange rate series - Indonesia-U.S, Indonesia-Japan and Indonesia-Singapore - are plotted in Figures 1, 2 and 3 respectively. It can be seen that there are two jumps in the 80's due to devaluation in March 1983 and September 1986. The big jump in the 90's is due to the Asian financial crisis in 1997. The detrended series are further filtered for five structural breaks and plotted in Figures 4, 5 and 6 respectively. Employing standard ADF test, and threshold unit root tests of Enders and Granger (1998), Lesari, Kim and Silvapulle (2003) found real exchange rate series to be non-stationary. Closely analysing the behaviour of the series, the small deviations from the long-run equilibrium are found to be persistent, while the large ones to be reverting back to the mean very fast. Further, these deviations are generally symmetric around the mean.

6. Application of the testing procedures against local and global stationary alternatives

The non-linear ESTAR model (5) was fitted to all three real exchange rate series, and the tests developed for testing against local and global stationary process were applied. The test results are reported in Table 7 along with the model parameter estimates. We first examine the results for the real exchange rates detrended as in models B and C. An analysis of the local stationarity test results indicates that the $\sup\hat{t}_{L,\delta^*=0}$ test statistics (in absolute values) are larger than the corresponding five per cent critical values reported in Table 1. However, although statistically insignificant, the estimated value of the parameter δ is greater than zero in the lower regime in all cases. These results imply that all three series are locally stationary, while they are globally non-stationary as indicated by insignificant values of the $\hat{t}_{G,\delta+\delta^*=0}$ - statistic.

Different results emerge for the real exchange series filtered for structural breaks. The estimated value of the parameter δ is less than zero, although statistically insignificant, in the lower regime for all cases. However, the finding that $\hat{t}_{G,\delta+\delta^*=0}$ - statistic larger than the five per cent critical value for all three real exchange rates indicates that these three series are globally stationary. Further, Indonesia-U.S and Indonesia-Japan real exchange rates appear to have a slow transition from lower to upper regimes as indicated by small values of γ (0.70 and 0.37 respectively), while the Indonesia-Singapore exchange rate was governed by a large value of γ (2.30). As has been discussed before, previous studies found many Asian real exchange rates to be non-stationary because these studies largely ignored the government interventions in the financial markets and the subsequent structural breaks in exchange rates. Contrary to these findings, our study provides concrete evidence to establish the stationary properties of Indonesian exchange rates relative to three countries – U.S., Japan and Singapore.

7. Conclusion

This paper develops a new sequential testing procedure to investigate stationary properties of real exchange rates. This involves (i) developing a testing procedure, in the presence of nuisance parameters, against a local non-linear exponential smooth transition autoregressive (ESTAR) stationary process; (ii) developing a test against a global ESTAR non-linear stationary process; and (iii) extending these two tests by incorporating dummy variables – which capture structural breaks in the series - into the data generating process. In this sequential testing, the null hypothesis of non-stationary series against a local ESTAR non-linear stationary alternative process is tested. If the null is rejected, then, testing against a global ESTAR non-linear stationary alternative process follows.

The tests developed in this paper are applied to real exchange rates of Indonesia against its major trading partners - U.S., Japan and Singapore. There are five notable structural breaks in these real exchanges rates due to Indonesian government interventions in the foreign exchange markets. This paper integrates three ideas - local stationarity, global stationarity and structural breaks - in developing tests to establish global stationary properties of real exchange rates. Application of the sequential testing procedure proposed in this paper provides concrete evidence in establishing global non-linear stationarity of all three real exchange rates. These tests will have much wider applications than those provided in the paper.

Figure 1: Demeaned and detrended Indonesia-U.S. real exchange rate for the original data

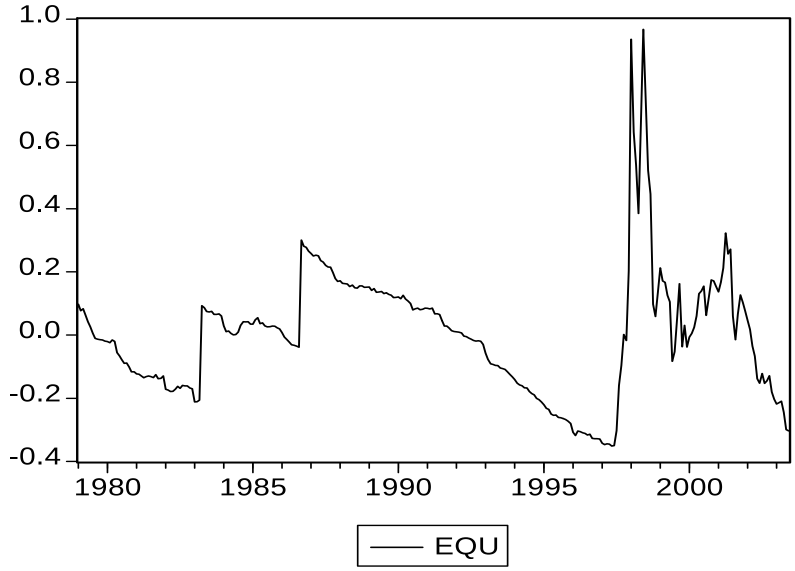


Figure 2: Demeaned and detrended Indonesia-Japan real exchange rate for the original data

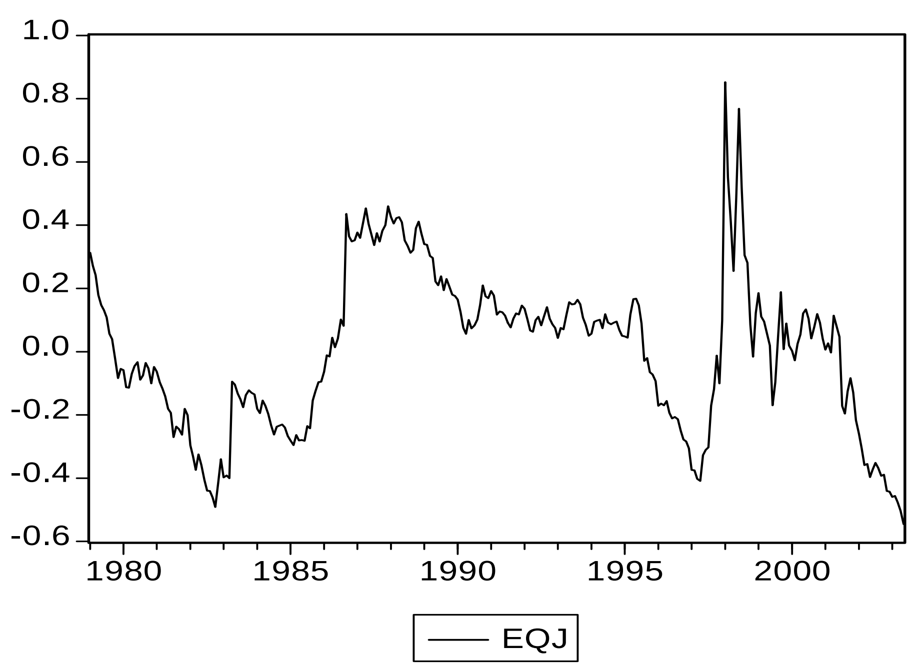


Figure 3: Demeaned and detrended Indonesia-Singapore real exchange rate for the original data

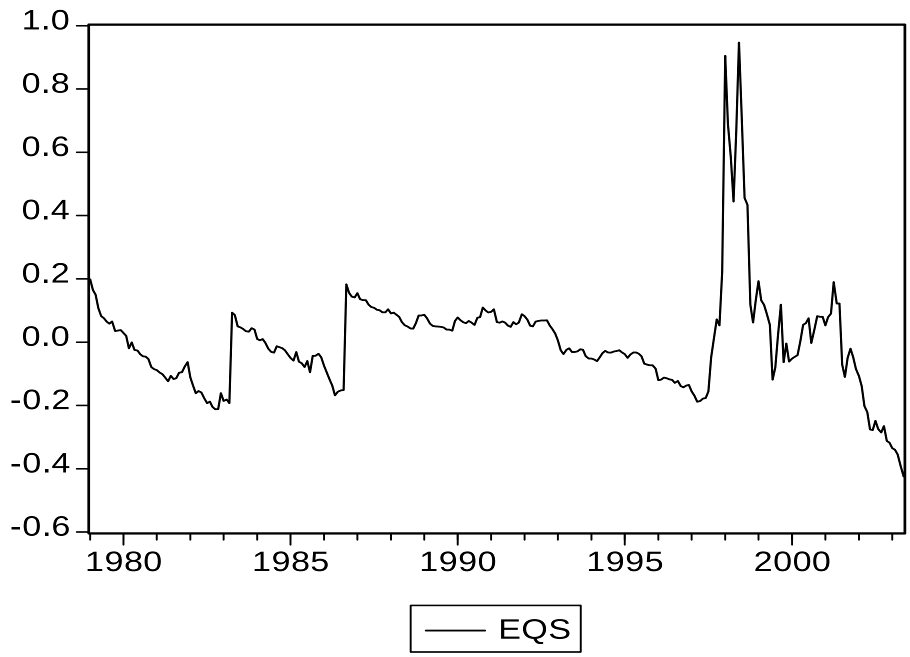


Figure 4: Indonesia-U.S real exchange rates for the dummy filtered data

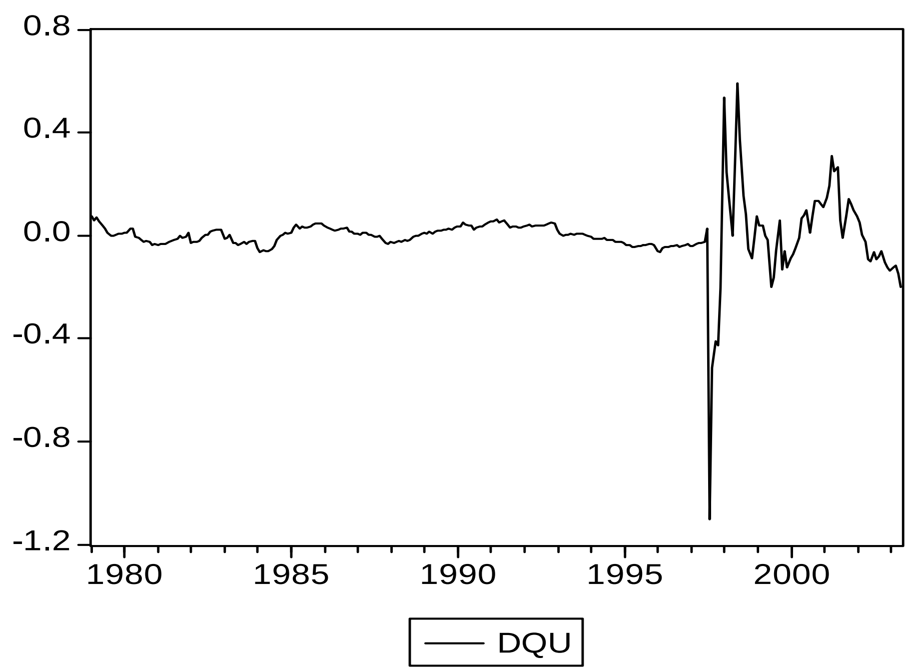


Figure 5: Indonesia-Japan real exchange rates for the dummy filtered data

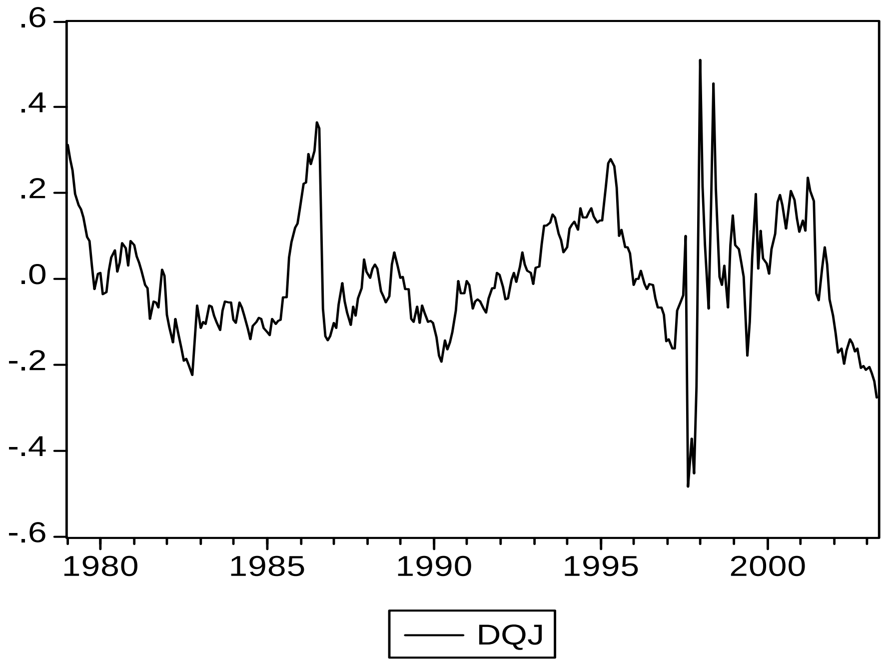


Figure 6: Indonesia-Singapore real exchange rates for the dummy filtered data

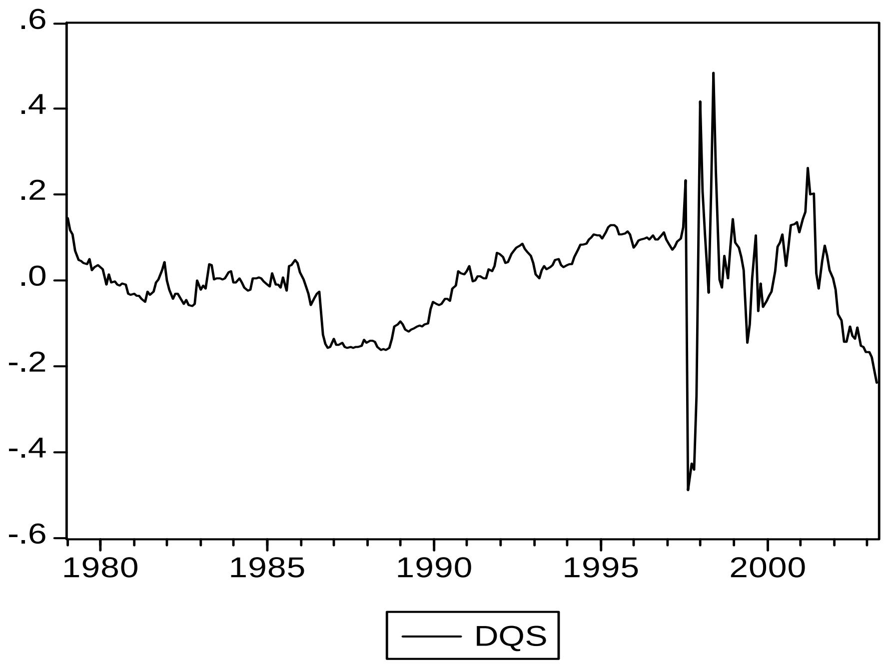


Table 1: Critical values of the estimated $\sup\hat{t}_{L,\delta^*=0}$ test statistic for a unit root against locally non-linear stationary process

Model: $\Delta y_t = \delta y_{t-1} + \delta^* y_{t-1}(1-\exp[(-\gamma(s_t - c)^2]) + u_t$ with $\delta = 0$.

Hypotheses: the null H_{0L}: $\delta^* = 0$ is tested against H_{1L}: $\delta^* < 0$

Test statistics: $\sup\hat{t}_{L,\delta^*=0} = -\sup\left|\hat{t}_{L,\delta^*=0}(\gamma,c)\right| = -\sup\left|\frac{\hat{\delta}^*}{s.e(\hat{\delta}^*)}\right|$, where $\hat{\delta}^*$ is a function of the parameter space spanned by (γ, c).

Sample size	*Models used to filter the series*	*Percentiles of the distribution of* $\sup\hat{t}_{L,\delta^*=0}$		
		1%	5%	10%
100				
	Model A	-3.211	-2.722	-2.364
	Model B	-3.799	-3.311	-2.996
	Model C	-4.304	-3.707	-3.501
300				
	Model A	-3.180	-2.672	-2.254
	Model B	-3.758	-3.200	-2.918
	Model C	-4.234	-3.686	-3.441
1000				
	Model A	-3.175	-2.575	-2.265
	Model B	-3.702	-3.175	-2.893
	Model C	-4.193	-3.663	-3.399
300				
	Model D	-3.700	-3.040	-2.710
	Model E	-4.210	-3.850	-3.520
	Model F	-4.980	-4.433	-4.164

Note: Model (*): $x_t = \mu + \beta t + \sum_{i=1}^{5}\phi_i D_i + y_t$

Model A: Model (*) with $\mu = \beta = \phi_i (i = 1,..,5) = 0$

Model B: Model (*) with $\mu \neq 0$ & $\beta = \phi_i (i = 1,..,5) = 0$

Model C: Model (*) with $\mu \neq 0$, $\beta \neq 0$ & $\phi_i (i = 1,..,5) = 0$

Models D, E and F are equivalent to (respectively) A, B and C with $\phi_i (i = 1,..,5) \neq 0$.

Table 2: Critical values of one-sided t-test for nonstationary against global nonlinear stationary

Model: $\Delta y_t = \delta y_{t-1} + \delta^* y_{t-1}(1 - \exp[(-\gamma(s_t - c)^2]) + u_t$ with $\delta \geq 0$.

Hypotheses: the null hypothesis $H_{0G}: \delta + \delta^* \geq 0$ is tested against $H_{1G}: \delta + \delta^* < 0$.

Test statistic: statistic $t_{G,\delta+\delta^*=0}$.

Sample size	*Deterministic variables*	*Percentiles of the distribution of* $t_{G,\delta+\delta^*=0}$.		
		1%	5%	10%
100				
	Model A	-2.801	-2.377	-2.101
	Model B	-3.301	-2.823	-2.389
	Model C	-3.724	-3.145	-2.985
300				
	Model A	-2.657	-1.991	-1.662
	Model B	-3.113	-2.479	-2.101
	Model C	-3.488	-3.002	-2.673
1000				
	Model A	-2.626	-1.990	-1.659
	Model B	-3.100	-2.458	-2.101
	Model C	-3.469	-3.000	-2.672
300				
	Model D	-3.102	-2.407	-1.999
	Model E	-3.603	-3.120	-2.573
	Model F	-4.218	-3.824	-3.564

See the notes for Table 1.
The percentiles of the distribution of the test statistic were calculated as functions of of the parameter space spanned by (γ, c).

Table 3: Rejection rates of $\sup \hat{t}_{L,\delta^*=0}$ test against local nonlinear stationary alternatives

Model: $\Delta y_t = \delta y_{t-1} + \delta^* y_{t-1}(1 - \exp[(-\gamma (s_t - c)^2]) + u_t$ with $\delta = 0$.

Hypotheses: the null H_{0L}: $\delta^* = 0$ is tested against H_{1L}: $\delta^* < 0$.

Test statistic: $\sup \hat{t}_{L,\delta^*=0}$

	δ^* =-0.1			δ^* = -0.5			δ^* = -1		
T=100 c=0	γ=0.2	0.6	1.2	γ=0.2	0.6	1.2	γ=0.2	0.6	1.2
Model A	0.128	0.330	0.446	0.358	0.991	0.997	0.945	1.000	1.000
Model B	0.114	0.288	0.435	0.251	0.978	0.978	1.000	1.000	1.000
Model C	0.109	0.249	0.418	0.197	0.986	0.886	1.000	1.000	1.000
T=100 c= 0.3									
Model A	0.112	0.230	0.305	0.420	0.867	0.957	0.709	0.995	1.000
Model B	0.108	0.128	0.260	0.338	0.779	0.880	0.652	0.971	0.991
Model C	0.075	0.103	0.234	0.283	0.638	0.786	0.570	0.950	1.000
T=300 c= 0									
Model A	0.339	0.795	0.951	0.996	1.000	1.000	1.000	1.000	1.000
Model B	0.290	0.687	0.898	0.956	1.000	1.000	1.000	1.000	1.000
Model C	0.245	0.594	0.700	0.931	1.000	1.000	1.000	1.000	1.000
T=300 c= 0.3									
Model A	0.237	0.611	0.788	0.800	1.000	1.000	1.000	1.000	1.000
Model B	0.145	0.579	0.605	0.699	0.988	1.000	1.000	1.000	1.000
Model C	0.138	0.481	0.520	0.631	0..800	1.000	1.000	1.000	1.000

See the footnote for Table 1.

Table 4: Rejection rates of $t_{G,\delta+\delta^*=0}$ for testing global nonlinear stationarity test)

Model: $\Delta y_t = \delta y_{t-1} + \delta^* y_{t-1}(1-\exp[(-\gamma\left(s_t - c\right)^2]) + u_t$ with $\delta \geq 0$.

Hypotheses: the null hypothesis $H_{0G} : \delta + \delta^* \geq 0$ is tested against $H_{1G} : \delta + \delta^* < 0$.

Test statistic: $t_{G,\delta+\delta^*=0}$.

	δ^* =-0.3 & δ=0.1			δ^* = -0.5 & δ=0.2			δ^* = -0.7 & δ=0.3		
T=100 c=0	γ=0.2	0.6	1.2	γ=0.2	0.6	1.2	γ=0.2	0.6	1.2
Model A	0.126	0.433	0.706	0.249	0.580	0.748	0.369	0.688	0.896
Model B	0.119	0.412	0.648	0.245	0.568	0.686	0.352	0.648	0.852
Model C	0.112	0.391	0.629	0.230	0.539	0.599	0.343	0.589	0.850
T=100 c= 0.3									
Model A	0.340	0.593	0.806	0.419	0.680	0.848	0.599	0.788	0.984
Model B	0.324	0.582	0.748	0.395	0.653	0.786	0.532	0.748	0.952
Model C	0.300	0.489	0.698	0.328	0.539	0.699	0.483	0.727	0.915
T=300 c= 0									
Model A	0.580	0.730	0.992	0.749	0.867	1.000	0.897	1.000	1.000
Model B	0.570	0.697	0.932	0.699	0.839	1.000	0.760	1.000	1.000
Model C	0.550	0.660	0.904	0.642	0.761	0.888	0.742	0.898	1.000
T=300 c= 0.3									
Model A	0.808	0.900	1.000	0.749	0.967	1.000	0.940	1.000	1.000
Model B	0.762	0.857	0.972	0.702	0.989	1.000	0.858	1.009	1.000
Model C	0.645	0.760	0.934	0.669	0.916	0.938	0.808	0.992	1.000

Table 5: Powers of $\sup\hat{t}_{L,\delta^*=0}$ test against local nonlinear stationary alternatives in the presence of structural breaks

Model: $\Delta y_t = \delta y_{t-1} + \delta^* y_{t-1}(1-\exp[(-\gamma(s_t-c)^2]) + u_t$ with $\delta = 0$.

Hypotheses: the null $H_{0L}: \delta^* = 0$ is tested against $H_{0L}: \delta^* < 0$

	δ^* =-0.1			*δ^* = -0.5*			*δ^* = -1*		
	γ=0.2	0.6	1.2	γ=0.2	0.6	1.2	γ=0.2	0.6	1.2
T=300									
c= 0									
Model D	0.291	0.695	0.841	0.886	1.000	1.000	1.000	1.000	1.000
Model E	0.235	0.583	0.786	0.835	1.000	1.000	1.000	1.000	1.000
Model F	0.212	0.454	0.6310	0.821	1.000	1.000	1.000	1.000	1.000
c= 0.3									
Model D	0.221	0.568	0.698	0.711	0.943	1.000	0.936	1.000	1.000
Model E	0.155	0.470	0.615	0.610	0.901	1.000	0.923	1.000	1.000
Model F	0.145	0.447	0.556	0.586	0.873	1.000	0.900	1.000	1.000

Table 6: Powers of $t_{G,\delta+\delta^*=0}$ test against global nonlinear stationary alternatives in the presence of structural breaks

Model: $\Delta y_t = \delta y_{t-1} + \delta^* y_{t-1}(1-\exp[(-\gamma(s_t-c)^2]) + u_t$ with $\delta \geq 0$.

Hypotheses: the null hypothesis $H_{0G}: \delta + \delta^* \geq 0$ is tested against $H_{1G}: \delta + \delta^* < 0$.

	δ^* =-0.3 & δ=0.1			*δ = -0.5 & δ=0.2*			*δ = -0.7 & δ=0.3*		
	γ=0.2	0.6	1.2	γ=0.2	0.6	1.2	γ=0.2	0.6	1.2
T=300									
c= 0									
Model A	0.482	0.639	0.897	0.693	0.814	1.000	0.824	1.000	1.000
Model B	0.430	0.597	0.830	0.634	0.717	0.973	0.792	1.000	1.000
Model C	0.350	0.562	0.794	0.582	0.631	0.898	0.718	0.863	1.000
c= 0.3									
Model A	0.698	0.830	1.000	0.749	0.967	1.000	0.940	1.000	1.000
Model B	0.632	0.774	0.972	0.702	0.989	1.000	0.858	1.009	1.000
Model C	0.553	0.668	0.934	0.669	0.916	0.938	0.808	0.992	1.000

Dummy filtered data $x_t = \mu + \beta t + \sum_{i=1}^{5} \phi_i D_i + y_t$

Table 7: The results of unit roots Indonesian real exchange rates

Model: $\Delta y_t = \delta y_{t-1} + \delta^* y_{t-1}(1-\exp[(-\gamma(s_t - c)^2]) + \sum_{j=1}^{p}\alpha_j \Delta y_{t-j} + u_t$

Hypotheses: the null $H_{0L}: \delta^* = 0$ is tested against $H_{1L}: \delta^* < 0$; test statistic: $\sup_{\gamma,c} \hat{t}_{L,\delta^*=0}$

Hypotheses: the null hypothesis $H_{0G}: \delta + \delta^* \geq 0$ is tested against $H_{1G}: \delta + \delta^* < 0$; test statistic:. $\hat{t}_{G,\delta+\delta^*=0}$.

T=300 (approximately)	$\hat{\delta}$	$\hat{\delta}^*$	$\hat{\gamma}$	$\hat{c}$	$\hat{d}$	$\sup_{\gamma,c} \hat{t}_{L,\delta^*=0}$	$\hat{t}_{G,\delta+\delta^*=0}$	*p*
Model B:								
Indon-US	0.144	-0.867	0.702	0.115	2	-3.501	-2.298	3
	(0.102)	(0.232)	(0.311)					
Indon-Jap	0.239	-0.936	0.370	0.188	2	-3.272	-1.376	2
	(0.122)	(0.283)	(0.151)					
Indon-Sing	0.182	-1.290	2.005	-0.089	2	-3.404	-2.008	2
	(0.099)	(0.349)	(0.520)					
Model C:								
Indon-US	0.140	-0.579	0.612	0.118	2	-3.922	-2.551	3
	(0.112)	(0.215)	(0.254)					
Indon-Jap	0.218	-0.827	0.340	0.180	2	-3.830	-2.290	2
	(0.128)	(0.235)	(0.113)					
Indon-Sing	0.139	-0.998	2.215	0.105	2	-3.879	-2.442	2
	(0.123)	(0.320)	(0.488)					
Model E:								
Indon-US	-0.102	-1.213	0.628	0.006	1	-4.367	-3.898	1
	(0.070)	(0.288)	(0.255)					
Indon-Jap	0.106	-0.984	0.390	0.112	1	-3.938	-3.390	2
	(0.055)	(0.264)	(0.159)					
Indon-Sing	-0.232	-1.794	2.706	0.099	1	-4.104	-3.782	1
	(0.090)	(0.446)	(0.499)					
Model F:								
Indon-US	-0.344	-1.620	0.225	0.110	1	-4.976	-4.270	1
	(0.101)	(0.382)	(0.100)					
Indon-Jap	-0.239	-1.936	0.577	0.100	1	-4.764	-4.493	2
	(0.099)	(0.481)	(0.157)					
Indon-Sing	-0.100	-1.290	2.599	0.004	1	-4.509	-4.993	1
	(0.079)	(0.349)	(0.601)					

Note: The statistic $\sup \hat{t}_{L,\delta^*=0}$ was computed assuming $\delta = 0$. The transition variable $s_t = \Delta y_{t-d}$.

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