NONLINEAR ADJUSMENT OF THE PURCHASING POWER PARITY IN INDONESIA

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Abstract

This paper models the dynamics of adjustment process to Indonesian long run purchasing power parity (PPP) relative to US, Japan and Singapore by employing a non-linear framework, which is recently shown to be appropriate in the presence of transaction costs associated with international trade. Using monthly observations from January 1979 to June 2003 (post-Bretton Woods period), covering the managed- and free-floating regimes in Indonesia, the real exchange rates were tested for their mean-reverting properties. A large number of studies found the real exchange series to be mean-averting and persistent, creating PPP puzzles. Using the linear framework many attempted to resolve these puzzles unsuccessfully. Motivated by the success of recent studies on PPP, applying a non-linear ESTAR to model the adjustment process, we tested for mean-reverting properties of all three real exchange rates for small and large deviations from the long-run equilibrium. We find that the small deviations are non-stationary, persistent and they can even be explosive, while the large deviations are stationary with adjustment process being very fast, making the overall adjustment process mean-reverting.

Keywords: Purchasing Power Parity, ESTAR model, Mean-reversion

JEL classification: F31, F32, C5, C22

1. Introduction

It is well known that international linkages for foreign exchange, goods and capital markets play a key role in the process that determines the exchange rate. The nature of the equilibrium implied by these linkages and the speed at which this equilibrium is attained have important implications for the ability of governments to pursue independent domestic monetary policies. It is widely accepted that the more integrated the international markets around the world the greater is the prospective difficulty in pursuing independent domestic monetary policies. An imperative question we attempt to answer in this paper is how to test for market integration. Despite the availability of a number of international parity conditions that can be used in such testing, this paper deals only with the purchasing power parity (PPP) condition, which measures the extent of integration between goods and foreign exchange markets across countries. This theory postulates that the nominal exchange rate between two national currencies adjusts to offset the excess of domestic inflation over foreign inflation, keeping the real exchange rate unchanged. This conventional PPP theory is unlikely to be valid if uncertainty is allowed and an explicit role for expectation is introduced.

The main objective of this paper is to examine the validity of the PPP hypothesis in the long run in Indonesia by focusing on the cross-currencies: Indonesian Rupiah-US Dollar, Indonesian Rupiah-Japanese Yen and Indonesian Rupiah-Singapore Dollar using a nonlinear framework. The United States and Japan were chosen since both are Indonesia's major trading partners, while Singapore is chosen in order to examine whether regional integration helps to achieve the long-run PPP relationship. Since Indonesia has adopted trade protection policies and limited the openness of its domestic markets, the transaction cost of trades is expected to be large. Consequently, the Indonesian market may not be fully integrated with the rest of the world. According to an emerging literature on nonlinear cointegration, it may be possible to model the adjustment process to long-run PPP for Indonesia using the mean-reverting exponential smooth transition autoregression (ESTAR) nonlinear model, the details of which will be discussed in the next section. Further, the nonlinear framework adopted in this paper is expected to resolve two purchasing power parity puzzles observed by a large number of studies in the literature. The first is the non-stationarity property of the real exchange rate (Rogoff, 1996 and Taylor et. al, 2001) and the second PPP puzzle is the high degree of persistence in the real exchange rate (Rogoff, 1996).

Testing for the validity of various forms of PPP is much studied in the empirical literature on international finance. These include an absolute form - the most restricted - and two unrestricted forms, with one being partly restricted and the other fully unrestricted. Many studies used conventional unit root procedures for testing the validity of PPP and failed to find evidence supporting it. Subsequently, several studies attempted to test this hypothesis using panel cointegration and fractional integration methodologies and found evidence in favour of the PPP theory. See Frankel and Rose (1996) and Taylor and Sarno (1998) for the former and Diebold et al. (1991) and Cheung and Lai (1993) for the latter.

Motivated by the work of Enders and Grangers (1998), Lestari, Kim and Silvapulle (2003) examined the validity of the Purchasing Power Parity (PPP) using non-linear threshold autoregressive (TAR) models. The results supported the fully unrestricted and partially unrestricted forms of PPP for Indonesia-U.S and Indonesia-Singapore exchange rates, while no evidence was found in supporting the restricted form. The assumption under the restricted form is that the nominal exchange rate fully adjusts for excess domestic inflation over foreign inflation. According to the theory, this adjusted series - defined as the real exchange rates - is expected to be stationary. Contrary to the expectation, all three real exchange rates were found to be non-stationary even in the nonlinear TAR framework. However, it has been recently argued that the lack of empirical evidence supporting Purchasing Power Parity is due to factors such as transaction costs, taxation, subsidies, actual or threatened trade restrictions, the existence of nontraded goods, imperfect competition, foreign exchange market interventions, and the differential composition of market baskets and hence the price indices across countries. Subsequently, an alternative framework for the empirical analysis of the PPP that allows for frictions in commodity trades has emerged. Dumas (1992) and Sercu, Uppal, and Van Hulle (1995) put forward theoretical arguments and developed equilibrium models of exchange rate determination in the presence of transaction costs and showed that the adjustment of real exchange rates towards the PPP is a non-linear mean-reverting process.

A few empirical studies have supported the long-run Purchasing Power Parity (Taylor, 1988 and Lothian and Taylor, 1996) when the sample period covered a long time span including the pre- and the post-Bretton Woods era. The results were somewhat mixed when the recent floating period was examined. Using the standard unit root tests, Corbae and Ouliaris (1988) cannot reject the presence of a unit root in the real exchange rate in the managed-float regime, providing the evidence against the PPP hypothesis in the long run. In contrast, Hakkio (1984) and Papell (1997) have found strong support for PPP hypothesis using panel data. However, in a simulation study, Taylor, Peel and Sarno (2001) found that the panel unit root null hypothesis is over-rejected, casting doubts on the panel cointegration results, while the conventional unit root test was found to have low power in small samples. The overall findings of these studies prompted researchers to come up with an alternative nonlinear framework for the empirical analysis of the PPP that allows for market frictions in the commodity trade and also to resolve the puzzles. Some of these studies are briefly discussed below.

Michael, Nobay and Peel (1997) investigated nonlinearities in the long-run Purchasing Power Parity relationship for the US, UK, France, Germany and Japan. They employed the exponential smooth transition autoregression (ESTAR) to model the adjustment process to long-run PPP and test for the mean-reverting property of real exchange rates. Then, they applied impulse response analysis to examine the dynamic adjustments of the long-run PPP. The results showed that four major real bilateral dollar exchange rates could be characterised by nonlinear mean-reverting behaviour during the interwar period.

Chen and Wu (2000), on the other hand, re-examined the long run PPP for US-Japan and US-Taiwan using the nonlinear framework. Japan and Taiwan were chosen because they have instituted a continuing policy of financial market liberalisation and experienced rapid growth, which has lead to increasingly strong ties to the US. The empirical analysis is based on monthly data of spot exchange rates and consumer price indices for the US, Japan, and Taiwan. The sample period spans January 1974 to December 1997 for Japan, and January 1980 to December 1997 for Taiwan. They employed the ESTAR and found that the parameter estimates of the ESTAR model revealed atypical behaviour of adjustment process for PPP deviations, this being random walk behaviour for small deviations and fast adjustment (mean-reverting) for large deviations from the PPP, which will be investigated in our study.

Baum, Barkoulas and Caglayan (2001) also studied the nonlinear adjustment of the deviations from the long-run PPP during the post-Bretton Woods period. They studied 17 countries of US trading partners and found the evidence of a mean reverting dynamic process for sizable deviations from PPP in several countries. Using generalised impulse response functions they also found evidence supporting nonlinear dynamic structure, but convergence to long-run PPP in the post-Bretton Woods era was found to be very slow. There is a parallel study by Taylor, Peel and Sarno (2001) on testing nonlinear mean reverting real exchange rates over the post-Bretton Woods period for the UK, Germany, France, and Japan. They tested the univariate model of the PPP, in the nonlinear framework and argued that the ESTAR model is more appropriate for modelling the real exchange rate movements since it captures the symmetric behaviour of its deviations well. Their results showed that these countries bilateral real exchange rates were characterised by a nonlinear mean reverting process during the floating rate period since 1973.

This paper is organised as follows: Section 2 briefly outlines the Smooth Transition Autoregressive (STAR) model, particularly ESTAR model. Section 3 presents various hypotheses and the tests related to STAR framework. Section 4 discusses the estimation of ESTAR model and the difficulties arising from it. Section 5 outlines the tests for various diagnostic checks. Section 6 discusses the conditions for the nonlinear mean-reverting adjustment towards the long run PPP. Section 7 describes the data series used in this study and defines the variables and models to be used in subsequent empirical analysis. Section 8 reports and analyses the empirical results. Some concluding comments are made in section 9.

2. Smooth Transition Autoregressive (STAR) Models

Granger and Teräsvirta (1993) argued that the nonlinear adjustment process can be characterised my smooth transition autoregressive (STAR) models, the reasons for this are given below. The STAR model of order p, for a time series y_t, has the following specification

$$y_t = \varphi_0 + \varphi_1' x_t + (\theta_0 + \theta_1' x_t) G(s_t; \gamma, c) + u_t \tag{1}$$

where $x_t = (y_{t-1,} y_{t-2}, \ldots.., y_{t-p})'$, $\varphi_1 = (\varphi_{1,} \varphi_2, \ldots.., \varphi_p)'$, $\theta_1 = (\theta_{1,} \theta_2, \ldots.., \theta_p)'$ are unknown parameter vectors, $G(.)$ is a transition function which is continuous and bounded by zero and one, s_t is the transition variable and c is the threshold parameter. It is assumed that $u_t \sim n.i.d(0, \sigma^2)$. The s_t may be a single stochastic variable, for example, an element of x_t, or a linear combination of stochastic variables or a deterministic variable such as a linear time trend. In the STAR model, the transition variable s_t is generally assumed to be the lagged endogenous variable, that is, $s_t = y_{t-d}$ for a certain integer $d > 0$ (Teräsvirta, 1994). This model can be extended by allowing exogenous variables z_t as additional regressors, and indeed one of them can be the transition variable. In this case, the model is called the smooth transition regression (STR) model (see Teräsvirta, 1998 for details). The STAR model can be interpreted as a regime switching model with two regimes, associated with the two extreme values of the transition function, which are $G(s_t; \gamma, c)$ =0 and $G(s_t; \gamma, c)$ =1, with the transition from one regime to the other being gradual. The regime that occurs at time t can be determined by the observable variable s_t and the associated value of $G(s_t; \gamma, c)$.

In the STAR model, however, the adjustment towards the equilibrium takes place at every point in time, but the speed varies with the size of the deviations from the long run PPP. In contrast with non-linear TAR model developed by Tong (1990) in that the regime changes occur abruptly, as argued before, while in the STAR they occur gradually. Michael, Nobay and Peel (1997) stated that STAR model is more attractive than the TAR model in describing the nonlinear adjustment process for the following reasons: first, the adjustment process is generally expected to be smooth rather than being discrete. Second, even if economic agents make only dichotomous decisions, it is highly likely that these decisions are made at different points in time. Therefore, in the aggregated processes, the change in regime expected to be continuous and smooth rather than discrete and abrupt. Finally, the modelling and the statistical inference procedures are more fully developed for STAR models than for TAR models. The fact that the TAR model arises due to discontinuity at each of threshold-parameter values complicates testing for linearity null hypothesis against nonlinear TAR alternatives.

There are two alternative forms for the transition function $G(.)$. The first is the logistic function, which can be written as

$$G(s_t;\gamma,c) = \{1+\exp[-\gamma(s_t-c)]\}^{-1}, \quad \gamma > 0 \qquad (2)$$

where γ measures the smoothness of transition from one regime to another and c is the threshold value for s, which indicates the halfway point between two regimes. Equation (1) combined with Equation (2) yields the Logistic STAR (LSTAR) model, in which there are two regimes, these being the appreciating and depreciating currencies in the foreign exchange market. They have different dynamics with the speed of adjustments varying with the extent of the deviation from the equilibrium. The transition function of LSTAR is of S-shape around c and monotonically increasing in s_t, yielding an asymmetric adjustment process towards the equilibrium, depending on whether or not these deviations are above or below the equilibrium.

The second is the exponential function, which can be written as

$$G(s_t;\gamma,c) = 1-\exp\left[-\gamma(s_t-c)^2\right], \qquad \gamma > 0 \qquad (3)$$

where, γ, as in LSTAR, measures the speed of transition from one regime to another and c represents the location for the threshold values for y. Equation (1) combined with Equation (3) yields the Exponential STAR (ESTAR) model. The transition function of ESTAR is symmetric and of U-shape around c. The ESTAR model suggests that the time series in the upper and lower regimes have rather similar dynamics. The ESTAR function in (3) defines a transition function about c where $G(.)$ is still bounded between 0 and 1. As $G(.)$ approaches either 0 or 1 the equation (3) reduces to a linear model.

The LSTAR and ESTAR models describe different dynamics of exchange rate behaviour. The main difference between these two STAR models is the discrepancies in the reaction of agents to shocks of the same size with opposite signs. The ESTAR models imply a symmetric U-shaped response of the exchange rate about the threshold parameter with respect to positive and negative shocks of the same magnitude. The asymmetries of S-shaped LSTAR responses, on the other hand, might be the result of differences in the reactions of the agents to these shocks.

3. Testing Hypotheses in the STAR Framework

Before proceeding to building STAR-type nonlinear models, an important step to carry out is to conduct various hypotheses in order to find statistically significance evidence supporting nolinearity hypothesis. This involves testing for linearity against STAR, misspecification testing and diagnostic checks, some of which are briefly outlined in this section.

Testing linearity against STAR

Testing for linearity against STAR is the first step towards building STAR models. Teräsvirta (1994, 1998) derived a linearity test against STAR. To explain this, first define $G^{*} = G - 1/2$, where G is the transition function defined in (3). Subtracting ½ from G is done only for the derivation of linearity test. Now, rewrite (1) as

$$y_t = \varphi_0 + \varphi_1' x_t + (\theta_0 + \theta_1' x_t) G^{*}(s_t; \gamma, c) + u_t \tag{4}$$

with previous notations being retained although φ_1 and φ_2 have changed. The assumption $u_t \sim n.i.d(0, \sigma^2)$ is made in order to derive the distribution of test statistics. Then, the conditional log-likelihood function of the model is given as

$$\sum_{t=1}^{T} \ell(\varphi, \theta, \gamma, c; y_t \mid x_t, s_t) = \alpha - \frac{T}{2} \log \sigma^2 - \frac{1}{2\sigma^2} \sum_{t=1}^{T} u_t^2 \tag{5}$$

The null hypothesis $H_0: \gamma = 0$ of linearity in (4) is tested against the alternative of nonlinearity hypothesis that $H_1: \gamma > 0$. However, there is an identification problem that arises in testing these hypotheses since the model is identified under the alternative but not under the null hypothesis. In order to resolve this problem, the transition function $G(.)$ is replaced by its third-order Taylor approximation, so the model becomes

$$y_t = \beta_0' x_t + \beta_1' x_t s_t + \beta_2' x_t s_t^2 + \beta_3' x_t s_t^3 + e_t . \tag{6}$$

Note that a higher order approximation can be used. The linearity hypothesis $H_0 : \gamma = 0$ in (4) is equivalent to the null $H_0' : \beta_1' = \beta_2' = \beta_3' = 0$ in (6). In small samples, the use of the F-test is recommended, because it has better size properties than the χ^2 version - LM test, which may heavily oversized in small samples (see Granger and Teräsvirta, 1993; chapter 7). Both, the χ^2 and F-test versions, can be computed by means of two auxiliary linear regressions. The F-test based on (6) can be computed as follows:

1. Estimate the model under the null hypothesis of linearity by regressing y_t on x_t. Compute the residuals $\hat{e}_t$ and the sum of squared residuals, say, $SSR_0 = \sum_{t=1}^{T} \hat{e}_t^2$.
2. Estimate the auxiliary regression of y_t on x_t and $x_t s_t^i$, $i = 1,2,3$. Compute the residuals $\hat{e}_t$ and the sum of squared residuals, say, $SSR_1 = \sum_{t=1}^{T} \hat{e}_t^2$.
3. Compute

$$F = \frac{(SSR_0 - SSR_1)/3p}{SSR_1/(T - 4p - 1)} \qquad (7)$$

where p is the number of explanatory variables.

Under the null hypothesis F-statistic has approximately an F distribution with $3p$ and $T - 4p - 1$ degrees of freedom.

Selecting the transition variable

The next stage is to select the appropriate transition variable to be used in the STAR model and the most suitable form of the transition function. The appropriate transition variable in the STAR model can be determined without specifying the form of the transition function. It is done by computing the F statistic for several candidate transition variables (say) $s_1, s_2, ..., s_n$, and selecting the one for which the p-value of the test statistic is the smallest. The rationale behind this procedure is that the test should have the maximum power, and, in this case, it means that the alternative model is correctly accepted. In other words, the correct transition variable is used (Van Dijk, 1999).

Selecting the transition function

If the STAR nonlinearity is accepted and the appropriate transition variable has been selected, the final decision to be made is to choose the most suitable form of the transition function $G(s_t;\gamma,c)$. The choice to be made is between the logistic function (2) and the exponential function (3). Teräsvirta (1994, 1998) suggests to use a decision rule based on a sequence of tests nested within the null hypothesis corresponding to F. He proposes to test the following hypotheses:

$$H_{04}: \beta_4' = 0,$$

$$H_{03}: \beta_3' = 0 / \beta_4' = 0, \tag{8}$$

$$H_{02}: \beta_2' = 0 / \beta_4' = \beta_3' = 0,$$

in

$$y_t = \beta_0' x_t + \beta_1' x_t s_t + \beta_2' x_t s_t^2 + \beta_3' x_t s_t^3 + \beta_4' x_t s_t^4 + e_t \tag{9}$$

using the LM -type tests. If $H_{04}: \beta_4' = 0$ is rejected, then LSTAR model is selected. Accepting $H_0: \beta_4' = 0$ and rejecting $H_{03}: \beta_3' = 0 / \beta_4' = 0$ imply that the ESTAR model in appropriate, while accepting both $H_0: \beta_4' = 0$ and $H_{03}: \beta_3' = 0 / \beta_4' = 0,$ but rejecting $H_{02}: \beta_2' = 0 / \beta_4' = \beta_3' = 0$ imply that the LSTAR model is appropriate. However, Granger and Teräsvirta (1993) and Teräsvirta (1994) argued that strict application of this sequence of tests can lead to the wrong conclusion. Then, they recommended that one should compute the p -values for all these F -tests and choose the STAR model on the basis of the lowest p -value. Therefore, if the rejection of $H_{04}: \beta_4' = 0$ or $H_{02}: \beta_2' = 0 / \beta_4' = \beta_3' = 0$ is accompanied by the lowest p -value, then the LSTAR model is chosen. On the other hand, if the rejection of $H_{03}: \beta_3' = 0 / \beta_4' = 0$ is accompanied by the lowest p -value, then the ESTAR model is chosen.

Furthermore, Van Dijk (1999) and Van Dijk, Teräsvirta and Francis (2001) stated that an alternative procedure for selecting the transition function proposed by Escribano and Jorda (1999) is superior to that developed by Teräsvirta (1994). Escribano and Jorda (1999) suggested testing the following hypotheses:

$$H_{0E} : \beta_2 = \beta_4 = 0$$

and (10)

$$H_{0L} : \beta_1 = \beta_3 = 0,$$

and their recommendation is to select the LSTAR(ESTAR) model if the minimum p-value is obtained for $H_{0L} (H_{0E})$.

4. Estimation

After the transition variable s_t and the transition function $G(s_t; \gamma, c)$ have been selected, the next stage is estimating the unknown parameters in the STAR model. The estimation of the parameters in the STAR model is carried out by the nonlinear least squares (NLS) method. That is, the parameters $\theta = (\varphi_1', \varphi_2', \gamma, c)'$ can be estimated as

$$\hat{\theta} = \arg\min_{\theta} Q_T(\theta) = \arg\min_{\theta} \sum_{t=1}^{T} (y_t - F(x_t; \theta))^2 \tag{11}$$

where $F(x_t; \theta)$ is the skeleton of the model

$$F(x_t; \theta) = \varphi_0 + \varphi_1' x_t + (\theta_0 + \theta_1' x_t) G(s_t; \gamma, c) + u_t . \tag{12}$$

Under the assumption that the error u_t is normally distributed, the NLS is equivalent to maximum likelihood. If u_t does not follow a normal distribution, the NLS estimates are quasi-maximum likelihood estimates. Therefore, under certain regularity conditions, the NLS estimators are consistent and asymptotically normal, that is,

$$\sqrt{T}\left(\hat{\theta} - \theta_0\right) \rightarrow N(0, C) \tag{13}$$

where θ_0 denotes the set of true parameter values. The asymptotic covariance-matrix C of $\hat{\theta}$ can be estimated consistently as $\hat{A}_T^{-1} \hat{B}_T \hat{A}_T^{-1}$, where $\hat{A}_T$ is the Hessian evaluated at $\hat{\theta}$

$$\hat{A}_T = -\frac{1}{T} \sum_{t=1}^{T} \nabla^2 q_t(\hat{\theta}) = \frac{1}{T} \sum_{t=1}^{T} \left(\nabla F(x_t; \hat{\theta}) \nabla F(x_t; \hat{\theta})' - \nabla^2 F(x_t; \hat{\theta}) \hat{u}_t \right) \tag{14}$$

with $q_t(\hat{\theta}) = (y_t - F(x_t; \hat{\theta}))^2, \nabla F(x_t; \hat{\theta}) = \partial F(x_t; \hat{\theta}) / \partial \theta,$ and $\hat{B}_T$ is the outer-product of the gradient

$$\hat{B}_T = \frac{1}{T} \sum_{t=1}^{T} \nabla q_t(\hat{\theta}) \nabla q_t(\hat{\theta})' = \frac{1}{T} \sum_{t=1}^{T} \hat{u}_t^2 \nabla F(x_t; \hat{\theta}) \nabla F(x_t; \hat{\theta})' \tag{15}$$

The model can be estimated using any conventional nonlinear optimisation procedure (Van Dijk, 1999; Van Dijk et. al., 2001).

Concentrating on the sum of squares function

Van Dijk (1999) and Van Dijk, Teräsvirta and Francis (2001) argued that the problems arising from estimation of the model can be simplified by concentrating on the sum of squares function. The STAR model is linear in autoregressive parameters φ_1 and φ_2, when the parameters γ and c in the transition function are known and fixed. Therefore, conditional upon γ and c, estimates of $\varphi = (\varphi_1', \varphi_2')'$ can be obtained by ordinary least squares (OLS) as

$$\hat{\varphi}(\gamma, c) = \left(\sum_{t=1}^{T} x_t(\gamma, c) x_t(\gamma, c)' \right)^{-1} \left(\sum_{t=1}^{T} x_t(\gamma, c) y_t \right) \tag{16}$$

where $x_t(\gamma, c) = (x_t'(1 - G(s_t; \gamma, c)), x_t' G(s_t; \gamma, c))'$. The notation $\hat{\varphi}(\gamma, c)$ is used to indicate that the estimate of $\hat{\varphi}$ is conditional upon γ and c. Therefore, the sum of squares function $Q_T(\theta)$ can be concentrated with respect to φ_1 and φ_2 as

$$Q_T(\gamma, c) = \sum_{t=1}^{T} (y_t - \varphi(\gamma, c)' x_t(\gamma, c))^2 \tag{17}$$

This will reduce the dimensionality of the nonlinear least squares estimation, since $Q_T(\gamma, c)$ will be minimized with respect to only two parameters γ and c.

Starting values

Starting values for the nonlinear optimisation can be obtained by two-dimensional grid search over γ and c. Replacing the transition function with

$$G(s_t; \gamma, c) = \left\{ 1 + \exp\left[-\frac{\gamma}{\hat{\sigma}_{s_t}^{n}} \prod_{i=1}^{n} (s_t - c) \right] \right\}^{-1} \tag{18}$$

where $\hat{\sigma}_{s_t}$ is the sample standard deviation of s_t, which makes γ to be approximately scale-free. The set of grid values for the location parameter c can be chosen from sample percentiles of the transition variable s_t. This guarantees that the values of the transition function contain enough sample variation for each choice of γ and c. If the transition function remains almost constant in the whole sample, the moment matrix of the regression in (16) is ill-conditioned, and the estimation procedure fails (Van Dijk et. al., 2001).

Estimating γ

As mentioned in the previous section, the smoothness of the transition between two regimes is characterised by γ. When the value of this parameter is large, it is difficult to obtain an accurate estimate of the smoothness of the transition between the two regimes. It is because, for such large values of γ**,** the STAR model becomes similar to a threshold model. To obtain an accurate estimate of γ, many observations in the immediate neighbourhood of c is needed. Because, even the large changes in γ have only a small effect on the shape of transition function. Therefore, the estimate of γ can be rather imprecise and often appear to be insignificant when it is judged by the t-statistic (Van Dijk, 1999 and Van Dijk et. al., 2001).

5. Diagnostic Checking

After estimating the parameters in the STAR model, the next step is to conduct specification testing to evaluate the fitted model. Various diagnostic checks need to be done to ensure that there is no residual autocorrelation, no remaining nonlinearity and parameter constancy.

Testing for residual autocorrelation

Eitrheim and Teräsvirta (1996) proposed the following test for serial independence in the residual. Consider the STAR model of order k with auto-correlated errors:

$$y_t = F(x_t;\theta) + \varepsilon_t \tag{19}$$

where $x_t = (1, \tilde{x}_t')'$, $\tilde{x}_t = (y_{t-1}, \ldots\ldots\ldots, y_{t-p})'$ as before and $F(x_t;\theta)$ is the skeleton of the model given in (12). An LM-test for q-th order serial dependence of ε_t can be obtained as nR^2, where R^2 is the coefficient of determination of the auxiliary regression of $\hat{\varepsilon}_t$ on $\nabla F(x_t;\hat{\theta}) = \partial F(x_t;\hat{\theta})/\partial\theta$ with $\theta = (\varphi_1', \varphi_2', \gamma, c)'$ and q lagged residuals $\hat{\varepsilon}_{t-1}, \ldots, \hat{\varepsilon}_{t-q}$. The symbol " $\wedge$ " indicates that the relevant quantities are the estimates under the null hypothesis of serial independence of ε_t. The resultant test statistic, denoted as $LM_{SI}(q)$, is asymptotically χ^2 distributed with q degrees of freedom.

Van Dijk (1999) and Van Dijk, Teräsvirta and Francis (2001) stated that this test statistic is a generalisation of the LM -test for serial correlation in an $AR(p)$ model of Godfrey-Breusch-Pagan (1979), which is based on the following auxiliary regression:

$$\hat{e}_t = \alpha_1 y_{t-1} + ... + \alpha_p y_{t-p} + \beta_1 \hat{\varepsilon}_{t-1} + ... + \beta_q \hat{\varepsilon}_{t-q} + \nu_t \tag{20}$$

where $\hat{\varepsilon}_t$ are the residuals of the $AR(p)$ model. Note that for a linear $AR(p)$ model, $F(x_t;\theta) = \sum_{i=1}^{p} \phi_i y_{t-i}$ and $\partial F(x_t;\hat{\theta}) / \partial\theta = (y_{t-1},......y_{t-p})'$.

Testing for remaining nonlinearity

Another diagnostic check is to test whether the estimated model successfully captured the nonlinear features of the time series entirely. To do this, we can apply a test for no remaining nonlinearity to an estimated auxiliary model. The natural approach is to specify the alternative hypothesis of remaining nonlinearity as the presence of an additional regime. For instance, testing the hypothesis that a two-regime model is adequate against the alternative that a third regime is necessary. Eitrheim and Teräsvirta (1996) develop an LM statistic to test a two-regime STAR model against the alternative of the following additive STAR model:

$$y_t = \varphi' x_t + (\theta' x_t) G(s_t;\gamma_1, c_1) + (\psi' x_t) H(s_t;\gamma_2, c_2) + u_t \tag{21}$$

The two-regime model that has been estimated is assumed to have $G(.)$ as transition function. Therefore, the hypothesis to be tested concerns the question whether or not extending the model with $(\psi' x_t) H(.)$ is appropriate. The null hypothesis of a two regime model is either $H_0 : \gamma_2 = 0$ or $H_0' : \psi = 0$. Again, this test suffers from a similar identification problem as encountered in testing the null hypothesis of linearity against the alternative of a two regime STAR model in section (3.1). Similarly, the solution to this identification problem is replaced the transition function $H(s_t;\gamma_2, c_2)$ by a Taylor series approximation around the point $\gamma_2 = 0$. Using a third-order approximation, the resultant approximation to model (21) is

$$y_t = \beta_0' x_t + (\theta' x_t) G(s_t;\gamma_1, c_1) + \beta_1' x_t s_t + \beta_2' x_t s_t^2 + \beta_3' x_t s_t^3 + e_t \tag{22}$$

where the parameters β_i, $i = 1,\ 2,\ 3$, are functions of the parameters ψ, γ_2 and c_2. The null hypothesis of no additional nonlinear structure or $H_0 : \gamma_2 = 0$ in (21) is equivalent to $H_0' : \beta_1' = \beta_2' = \beta_3' = 0$ in (22). The test statistic can be computed as nR^2 from the auxiliary regression of the residuals (obtained from estimated model under the null hypothesis) on the partial derivatives of the regression function with respect to the parameters in the two-regime model θ, γ_1 and c_1, evaluated under the null hypothesis, and the auxiliary regressors $x_t s_t^i$, $i = 1,2,3$. The resultant F statistic has an asymptotic χ^2 distribution with $3p$ degrees of freedom.

6. Nonlinear Adjustment to the Long Run Purchasing Power Parity

As has been argued before, because of the transaction costs the adjustments to positive and negative deviations from the long-run PPP equilibrium are expected to be same. Michael, Nobay and Peel (1997) argued that the ESTAR model is more appropriate for modelling PPP deviations, since it has symmetric adjustments to positive and negative deviations of the same magnitudes. Incorporating the equations (1) and (3), the ESTAR model for the deviations from the PPP is modelled as follows:

$$\nu_t = k + \sum_{j=1}^{p} \pi_j \nu_{t-j} + (k^* + \sum_{j=1}^{p} \pi_j^* \nu_{t-j}) \times \left\{1 - \exp\left[-\gamma (s_t - c)^2\right]\right\} + u_t \qquad (23)$$

where ν_t is a stationary and ergodic[1] process, $u_t \sim n.i.d(0, \sigma^2)$, and $\gamma > 0$. As mentioned above, the transition function $G(.)$ is U-shaped and the parameter γ determines the speed of the transition process between the two extreme regimes. The middle regime corresponds to $s_t = c$, yielding $G = 0$, and then (23) becomes a linear $AR(p)$ model:

$$\nu_t = k + \sum_{j=1}^{p} \pi_j \nu_{t-j} + u_t \qquad (24)$$

The outer regime corresponds to $s_t = \pm\infty$, yielding $G = 1$, and then (23) again becomes an $AR(p)$ model, but with a different set of parameters:

$$\nu_t = k + k^* + \sum_{j=1}^{p} (\pi_j + \pi_j^*) \nu_{t-j} + u_t \qquad (25)$$

For testing purposes, it is convenient to reparameterize the ESTAR model in (23) as follows:

$$\Delta \nu_t = k + \lambda \nu_{t-1} + \sum_{j=1}^{p-1} \phi_j \nu_{t-j} + (k^* + \lambda^* \nu_{t-1} + \sum_{j=1}^{p-1} \phi_j^* \nu_{t-j})$$

[1] A stationary process is ergodic if it is asymptotically independent, that is any two random variables positioned far apart in the sequence are almost independently distributed (see Hayashi, F., 2000: p.101, for details).

$$\times \left\{1 - \exp\left[-\gamma (s_t - c)^2\right]\right\} + u_t \tag{26}$$

The crucial parameters in (26) are λ and λ^* which determine whether or not the small and large deviations respectively are mean-reverting. The effect of transaction costs on the real exchange rates suggests that the larger the deviation from long-run PPP equilibrium, the stronger the tendency to move back to the equilibrium. This implies that while $\lambda \geq 0$ is possible, the conditions that $\lambda^* < 0$ and $\lambda + \lambda^* < 0$ should be satisfied for the process to be global stationary. Under these conditions, for small deviations, y_t may follow a unit root or even exhibit explosive behaviour, but for large deviations the process is mean-reverting (Michael, Nobay and Peel, 1997).

The analysis based on the ESTAR model above has implications for the conventional cointegration test of PPP, which is based on a linear AR(p) model, written below as an augmented Dickey-Fuller regression:

$$\Delta \nu_t = k' + \lambda' \nu_{t-1} + \sum_{j=1}^{p-1} \phi'_j \Delta \nu_{t-j} + u_t \tag{27}$$

Assuming that the true process for ν is given by the nonlinear model (26), then estimates of the parameter λ' in (27) will tend to lie between λ and ($\lambda + \lambda^*$). Hence, the null hypothesis $H_0 : \lambda' = 0$ (no linear cointegration) may not be rejected against the stationary alternative hypothesis $H_1 : \lambda' < 0$, even though the true nonlinear process is globally stable with $\lambda + \lambda^* < 0$. This shows that the failure to reject the unit root hypothesis on the basis of a linear model does not necessarily invalidate the long-run PPP (Michael et. al., 1997 and Taylor et. al. 2001).

7. Data Series

The data series used in this study are monthly observations from January 1979 to June 2003 taken from the IMF's International Financial Statistics CD-ROM. The time period covers the managed-floating and free-floating regimes in Indonesia and it appear to be long enough to test for the PPP condition as a long-run relationship. The nominal exchange rates used in this study are the Indonesian Rupiah against US-Dollar, Japanese Yen, and Singaporean Dollar. The domestic price is Indonesian CPI and the foreign prices are US, Japan and Singapore CPI series. The relative price is defined as the ratio of domestic price to foreign price.

As discussed in Lestari, Kim, and Silvapulle (2003), the PPP models can be classified into three different forms, namely, the univariate, bivariate and multivariate models, depending on the nature of restriction(s) imposed. The empirical results in this study are based on investigating the appropriateness of ESTAR models for the deviations from PPP – defined by all three univariate, bivariate and multivariate models. When empirical studies did not find support for the univariate model they studied less restricted forms, bivariate and multivariate models. The univariate model (as argued before) is the real exchange rate – which is the nominal exchange rate fully adjusted to offset excess domestic inflation over foreign inflation. In the bivaraite model, the nominal exchange rate is allowed to partially adjust to this excess inflation, while, in the multivariate model, the nominal exchange rate is allowed to respond to domestic and foreign inflation rates separately. Empirical studies of bivariate and multivariate models have emerged as there was no support found for the fully restricted form of PPP – the univariate model.

8. Empirical Results and Analysis

Empirical analysis is carried out in different stages, which are given below.

Linearity Test Results

Although the nonlinear ESTAR has been recommended for modelling the deviations from PPP, in this study, testing for the null hypothesis of linearity against STAR was done first to find out whether the non-linear framework is more appropriate to model the process than the linear counterpart. Having rejected the linearity, testing was then done against LSTAR and ESTAR models separately. The reason is to find empirical support for the ESTAR model, among others.

The linearity test is carried out with different values of delay parameter (d), with d ranging from 1 to 10. Table 1 reports the linearity test results of the hypotheses given in (9). In the univariate case, the results indicate that the ESTAR process with the delay parameter $d = 1$ is an appropriate representation of the adjustment of the deviations from the long-run PPP equilibrium for all three real exchange rates and all the unrestricted models with two exceptions. These being for the multivariate model of Indonesia-US exchange rate, ESTAR process with $d = 4$ and for the univariate model of Indonesia-Japan exchange rate, ESTAR process with $d = 2$.

Estimation of ESTAR Models

The three real exchange rate (demeaned) series are plotted in Figure 1. It can be seen that there are two jumps in the 80's due to the devaluation in March 1983 and September 1986. The big jump in the 90's is due to the Asian financial crisis. The movement of these series over the sample period clearly indicate that they are mean-reverting. However, employing standard and threshold unit root tests, Lestari, Kim and Silvapulle (2003) found that the real exchange rate series are non-stationary, mean-averting. Closely analysing the behaviour of the series, the small deviations from the long run PPP are found to be persistent, while the large ones to be reverting back to the mean very fast. Further, the movements are fairly symmetric around the mean. These observations are consistent with the ESTAR model discussed in Section 2. In what follows, ESTAR is fitted to all three exchange rate series under all three assumptions briefly discussed in the previous section.

In the univariate model, the transition parameter γ estimates for all series were found to be quite big (see Table 2). It was 14.68 for the Indonesia-US real exchange rate, 6.25 for the Indonesia-Japan real exchange rate and 9.33 for Indonesia-Singapore real exchange rate. These estimated values indicate that the real exchange rates have a high speed of adjustment towards the long-run PPP equilibrium. Figures 2(a), 3(a) and 4(a) show the estimated transition functions for the real exchange rates. Since all the γ values in the univariate model are found to be significantly different from zero, it can be said that the ESTAR model can represent the adjustment process towards the long-run equilibrium of PPP well. Furthermore, the residuals are found to follow a white noise process, as indicated by the *p*-value associated with Q-statistics at lag 6 for all real exchange rates. However, the ARCH effects appear to be present in all cases.

Testing for the mean-reversion property of the series, the Indonesia-Singapore real exchange rate was found to have explosive behaviour in the lower regime as $\lambda > 0$, while the Indonesia-U.S. and Indonesia-Japan real exchange rates were found to have unit roots in the lower regime as $\lambda = 0$ was found to be true in both cases (see Table 3). Further, all three real exchange rate series were found to have stationary behaviour in the upper regime as $\lambda^* < 0$ in all three cases. However, the stability condition $\lambda + \lambda^* < 0$ is satisfied in all cases. It can be said that all exchange rates have stationary mean-reverting behaviour overall.

In the bivariate model of PPP, Table 2 reports that the Indonesia-Japan exchange rate was found to be characterised by a small γ**,** 0.31, while the Indonesia-US and Indonesia-Singapore exchange rates were by large γs**,** 3.73 and 4.86 respectively. These results suggest that the

Indonesia-Japan exchange rate has a low speed of adjustment towards the long-run equilibrium of PPP, while Indonesia-US and Indonesia-Singapore exchange rates have high speeds of adjustments towards the long-run equilibriums of PPP. Figure 2(b), 3(b) and 4(b) show the patterns of transition functions for the bivariate models of PPP. The γ values for Indonesia-US and Indonesia-Singapore exchange rates were found to be statistically different from zero indicating that the ESTAR model can be used to model the adjustment process towards the long-run equilibrium of PPP of the real exchange rates, while the γ value for the Indonesia-Japan exchange rate is found to be statistically equal to zero. However, this result should not be interpreted as evidence against the ESTAR model, because the estimate standard error of γ is rather imprecise in general and often appears to be insignificant when judged by its t-statistic (see Frances and van Dijk, 2000, pp.91). Furthermore, the residuals are also found to follow a white noise process, indicated by the *p*-value associated with Q-statistics at lag 6 for all real exchange rates. As with univariate models, the ARCH effects appear to be present in all cases. Furthermore, Table 3 showed that all cases have explosive behaviour in the lower regime as $\lambda > 0$, and the stability condition $\lambda + \lambda^* < 0$ is satisfied. Clearly, all exchange rates have stationary mean-reverting behaviour.

Now, turning to the results of estimating the multivariate (fully unrestricted) models of PPP, the Indonesia-Japan exchange rate is found to be characterised by a small γ of 0.43, while Indonesia-U.S. and Indonesia-Singapore exchange rates by large γs**,** 2.16 and 10.89 respectively. These results suggest that the Indonesia-Japan exchange rate has low speeds of adjustments towards the long-run equilibrium of PPP, while Indonesia-US and Indonesia-Singapore exchange rates have high speeds of adjustment towards the long-run equilibriums of PPP (see Table 2). Figure 2(c), 3(c), and 4(c) show the transition functions for the multivariate models of PPP. However, not all the γ values in the multivariate model are significantly different from zero, such as in the Indonesia-Japan exchange rate which has γ value equal to zero. The ESTAR model, still, can represent the adjustment process towards the long-run equilibrium of PPP for the multivariate model. As in other models, the residuals are found to follow a white noise process as indicated by the *p*-value associated with Q-statistics at lag 6 for all cases and the ARCH effects are present in all cases. Further, an explosive behaviour in the lower regime ($\lambda > 0$) is found in Indonesia-Japan and Indonesia-Singapore exchange rates, while the Indonesia-US exchange rate has a unit root in the lower regime ($\lambda = 0$). All exchange rates show mean reverting behaviour in the overall adjustment process as indicated by $\lambda + \lambda^* < 0$ (see Table 3).

9. Conclusion

This paper models the dynamics of the adjustment process to Indonesian long run purchasing power parity relative to US, Japan and Singapore by employing a non-linear framework, which is recently shown to be appropriate in the presence of transaction costs associated with international trade. Using monthly observations from January 1979 to June 2002 (post-Bretton Woods period), covering the managed- and free-floating regimes in Indonesia, the real exchange rates were tested for mean-reverting properties. The data series used includes the domestic price (which is Indonesian CPI) and the foreign prices (the US, Japan and Singapore CPI series). The relative price is defined as the ratio of domestic price to foreign price. The real exchange rate is defined as the difference between the nominal exchange rates and the relative price ratio. A large number of studies found that the real exchange series are mean-averting and persistent, creating PPP puzzles. Using the linear framework many attempted to resolve these puzzles unsuccessfully. Motivated by the success of recent studies on PPP, applying the non-linear ESTAR to model the adjustment process, we tested for mean-reverting properties of all three real exchange rates for small and large deviations from the long-run equilibria.

We find that the small deviations are non-stationary, persistent and it can even be explosive, while the large deviations stationary with fast adjustment, making the overall adjustment process mean-reverting. These results are consistent with the previous findings. Further, the real exchange rate implied by PPP is a very restricted form of PPP condition. We also examined less restricted and fully unrestricted forms of PPP and found the results are stronger than those for the restricted form. It is noteworthy that the nonlinear ESTAR model helps to resolve the two PPP puzzles, which many empirical studies made considerable efforts to resolve for many decades unsuccessfully.

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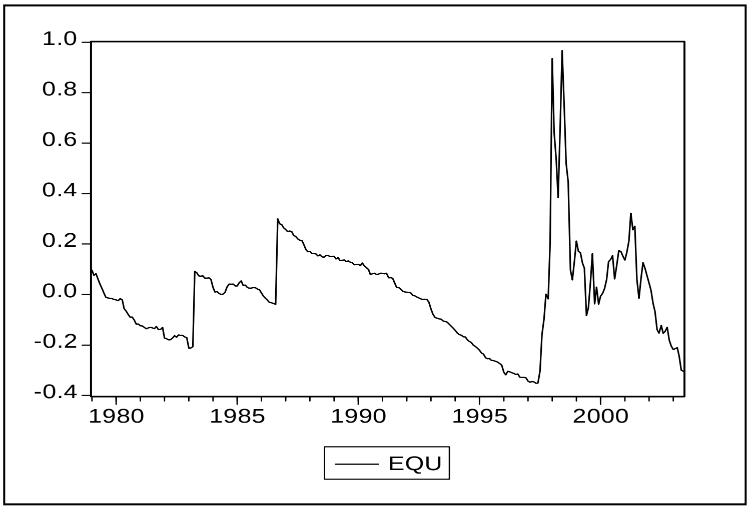
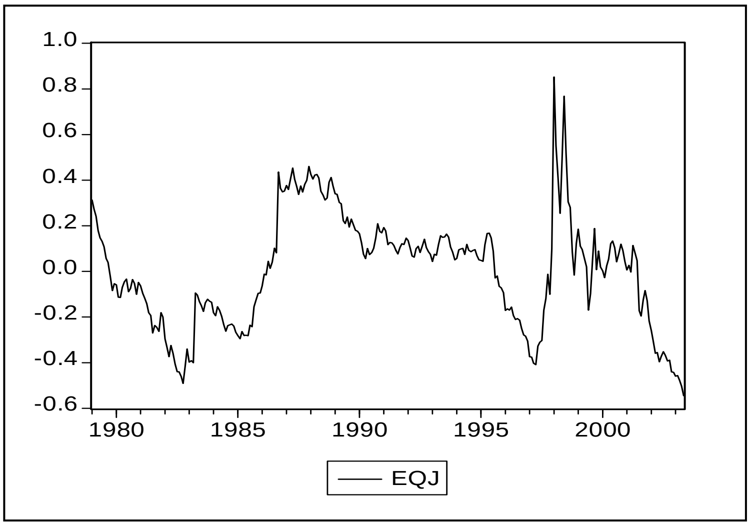
Figure 1:

(a) Demeaned and detrended Indonesia-U.S. Real Exchange Rate



(b) Demeaned and detrended Indonesia-Japan Real Exchange Rate



(c) Demeaned and detrended Indonesia-Singapore Real Exchange Rate

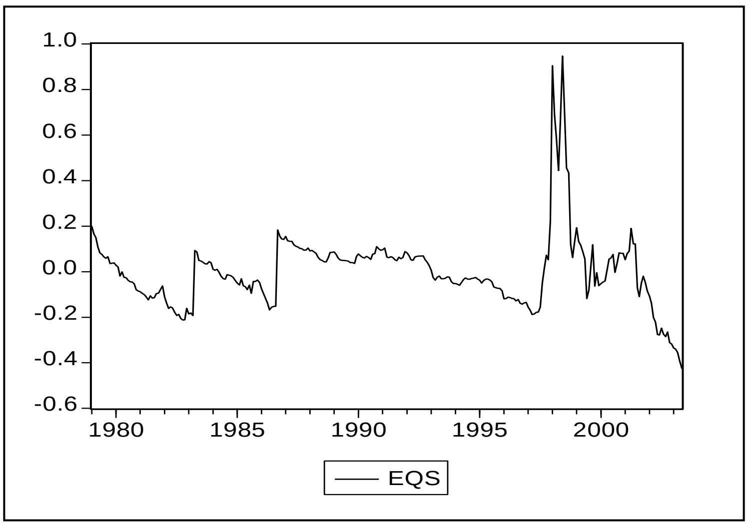


Table 1: Linearity Tests Result

PPP Models	*d*	*p*-value	F- Test	Model
US as foreign country				
Univariate	1	0.000001	4.593126	ESTAR
Bivariate	1	0.000000	11.866165	ESTAR
Multivariate	4	0.000000	7.444207	ESTAR
Japan as foreign country				
Univariate	2	0.000246	3.219213	ESTAR
Bivariate	1	0.000000	5.010695	ESTAR
Multivariate	1	0.000000	5.043555	ESTAR
Singapore as foreign country				
Univariate	1	0.000000	8.639613	ESTAR
Bivariate	1	0.0000000	20.922625	ESTAR
Multivariate	1	0.0000000	13.977329	ESTAR

Notes:
- The optimal *d* chosen is which gives the lowest the *p*-value of the linearity test over the range $1 \leq d \leq 10$

Table 2: Estimation of the ESTAR models

Series	γ	Std. Err	c	Q(6)	ARCH(6)	AR(*p*)
US as foreign country						
Univariate	14.6753	0.0467	0.4914	18.916	16.463	3
Bivariate	3.7299	0.0779	0.4914	7.3672	12.5530	6
Multivariate	2.1581	0.1163	0.3538	4.9604	17.5143	6
Japan as foreign country						
Univariate	6.2509	0.0505	0.0853	14.712	16.0062	3
Bivariate	0.3114	0.6459	0.0643	10.710	10.9611	4
Multivariate	0.4253	0.3613	0.0551	5.2928	11.6541	5
Singapore as foreign country						
Univariate	9.3304	0.0414	0.3111	9.8413	14.9904	5
Bivariate	4.8555	0.0299	0.5316	16.538	12.9517	5
Multivariate	10.8859	0.0000	0.4228	11.274	16.6295	5

Notes:
- Q(6) is Ljung-Box statistics for residual autocorrelation for lag six.
- ARCH(6) is Engle’s ARCH-LM test for ARCH with lags six.
- ARp(*p*) ischosen on the basis of serial correlation tests.

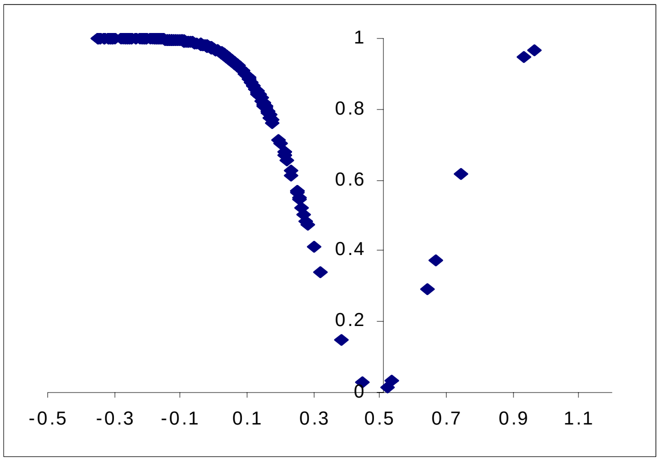
Table 3: Estimation of the ESTAR models (continued)

PPP Models	λ' *(s.e)*	λ *(s.e)*	λ^* *(s.e)*	$\lambda + \lambda^*$ *(F_stat)*
US as foreign country				
Univariate	0.0148 *(0.0104)*	0.4840 *(0.2927)*	-1.5900 *(0.3242)*	-1.7395 *(6.4695)*
Bivariate	-0.0539 *(0.0271)*	0.5510 *(0.1295)*	-2.1192 *(0.2045)*	-1.5682 *(5.7282)*
Multivariate	-0.0795 *(0.0307)*	0.1051 *(0.1376)*	-1.4035 *(0.4621)*	-1.2984 *(6.7626)*
Japan as foreign country				
Univariate	-0.0130 *(0.0084)*	-0.0903 *(0.0748)*	-0.6716 *(0.2022)*	-0.9361 *(7.5527)*
Bivariate	-0.0214 *(0.0177)*	0.2230 *(0.0832)*	-3.2795 *(0.7730)*	-3.0565 *(4.1807)*
Multivariate	-0.0525 *(0.0224)*	0.2398 *(0.0847)*	-3.7356 *(0.6125)*	-3.4958 *(5.0436)*
Singapore as foreign country				
Univariate	-0.0148 *(0.0100)*	0.4276 *(0.1783)*	-1.6339 *(0.2170)*	-6.2826 *(7.0308)*
Bivariate	-0.0474 *(0.0282)*	0.7370 *(0.1487)*	-2.3486 *(0.8161)*	-1.6116 *(5.8615)*
Multivariate	-0.0824 *(0.0324)*	0.3050 *(0.0858)*	-1.5056 *(0.3271)*	-1.2006 *(4.9086)*

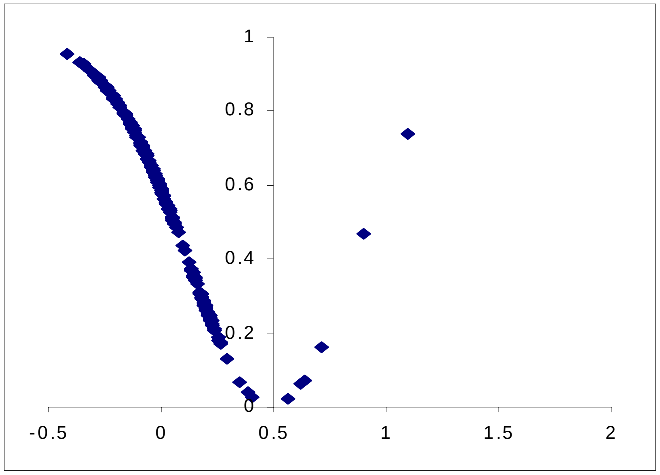
Notes:
- λ' is obtained from eq. (5.27)
- λ and λ^* are obtained from eq. (5.26)

Figure 2. Transition Function for Indonesia-U.S. Exchange Rate

(a) Univariate Model (γ=14.68)



(b) Bivariate Model (γ=3.72)



(c) Multivariate Model (γ=2.16)

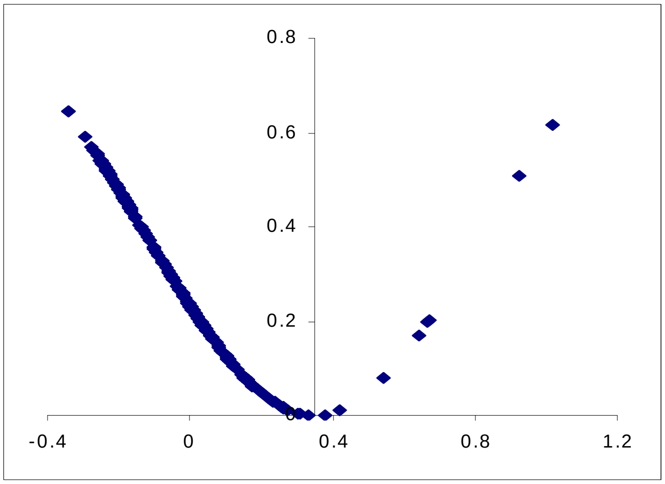
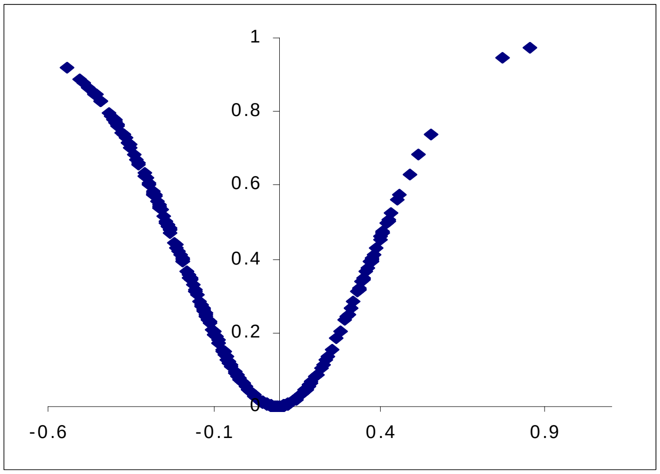
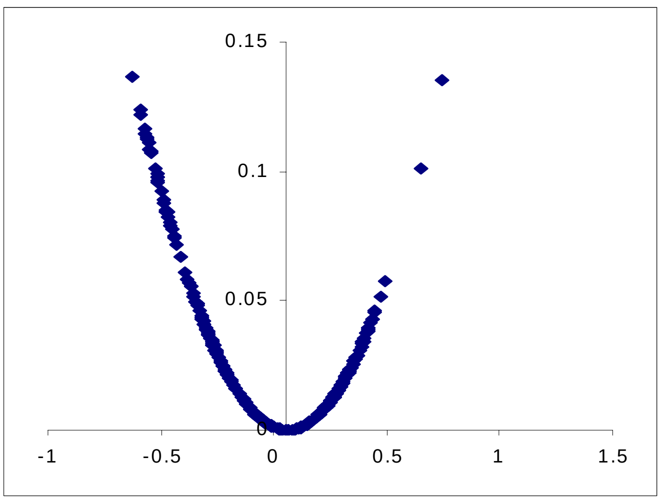


Figure 3. Transition Function for Indonesia-Japan Exchange Rate

(a) Univariate Model (γ=6.25)



(b) Bivariate Model (γ=0.31)



(c) Multivariate Model (γ=0.43)

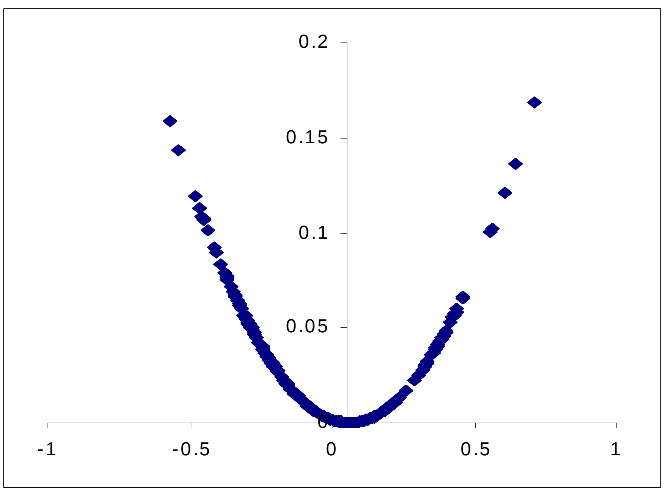
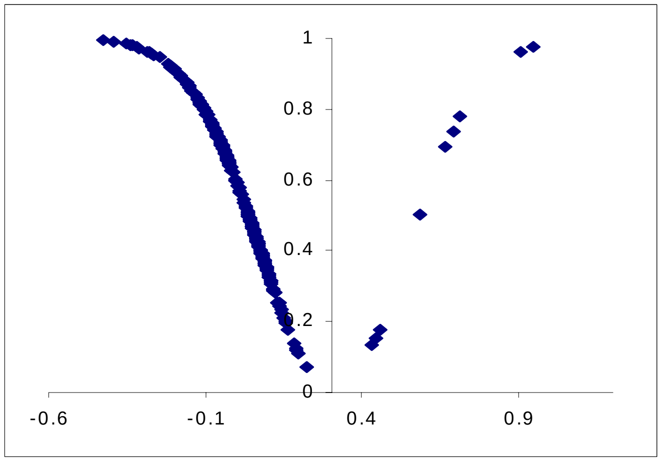
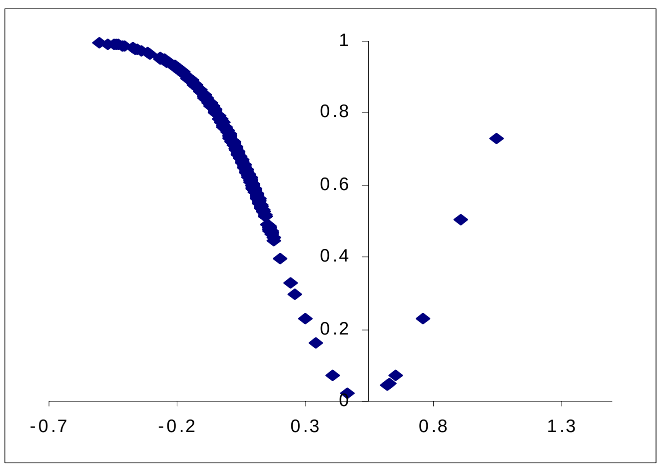


Figure 4. Transition Function for Indonesia-Singapore Exchange Rate

(a) Univariate Model (γ=9.33)



(b) Bivariate Model (γ=4.86)



(c) Multivariate Model (γ=10.89)