**ASIAN REAL INTEREST RATES, NON-LINEAR DYNAMICS AND INTERNATIONAL
PARITY**

*Mark J. Holmes and Nabil Maghrebi**

December 2002

Acknowledgement. Permission to access the financial database at the Institute for Economics and Econometrics at Wakayama University is gratefully acknowledged. Any remaining errors are our own.

* Mark J. Holmes and Nabil Maghrebi are respectively, Senior Lecturer at the Department of Economics, Loughborough University, UK and Associate Professor at the Faculty of Economics, Wakayama University, Japan.

ASIAN REAL INTEREST RATES, NON-LINEAR DYNAMICS AND INTERNATIONAL PARITY

Abstract

This study tests for non-linearities in the real interest differentials of four South East Asian economies with respect to Japan and the U.S. The logistic and exponential smooth transition regression models are applied to monthly data over the sample period 1977M1-2000M3. There is evidence of non-linearities in Asian real interest differentials where non-linearities are often captured by the logistic smooth transition autoregressive model. The extent of non-linearities varies across the sample with the Singapore-Japan and Thailand-Japan differentials exhibiting the sharpest transition from one regime to another. Large shocks to real interest parity are more likely to lead to the reestablishment of parity at a faster rate than small shocks. Modeling the non-linear stochastic dynamics of real interest parity can thus be useful for policy-making purposes in recovering information on monetary and financial crises.

The extent to which real interest rates are equalized across countries has occupied researchers for a number of reasons. In an open economy, real interest rates play a key role in influencing real activity through saving and investment behavior. Confirmation or rejection of real interest parity (RIP) provides an indication of whether countries are financially integrated or autonomous. However, since RIP requires that *ex ante* purchasing power parity (PPP) holds, it can be viewed as a more general indicator of

macroeconomic integration or convergence. RIP is also important because it is a key working assumption in various models of exchange rate determination and the focus of many studies as early as Frenkel (1976), Mussa (1976) and Frankel (1979). The purpose of this paper is to examine how plausible such an assumption is and whether or not adjustments towards RIP between South East Asian economies and Japan and the U.S. can be characterized as non-linear.

The investigation of non-linearities and asymmetries in the behavior of macroeconomic variables constitutes an increasingly important area of empirical finance. Many recent studies including Taylor (2001), Iannizzotto (2001), McMillan and Speight (2001), Serletis and Gogas (2000), Sarno (2000a, 2000b), Sarantis (1999) and Michael *et al.* (1997), provide strong argument and empirical evidence of non-linearities mainly with respect to OECD real exchange rates. The importance of evidence on non-linearities in real exchange rates derives from the fact that it may be reflect the degree of heterogeneity of foreign exchange market participants in terms of the formulation of objective functions and formation of expectations.¹ It is against the background of recent research trends and the importance of non-linearity in understanding of the behavior economic agents and economic convergence that the present study should be of interest for many reasons.

As a contribution to the debate concerning RIP, this study undertakes the analysis of real interest parity in four South East Asian economies using monthly data for the period 1977M1 to 2000M3. Since RIP is predicated on uncovered interest parity (UIP) and *ex ante* PPP, useful insights into how economies adjust in response to a shock to RIP can be obtained. Moreover, adjustment towards RIP requires both financial and goods market arbitrage where recent studies by Balke and Wohar (1998) along with the recent work on real exchange rates mentioned above, suggest that non-

linearities can arise through the presence of transactions costs. An implication of non-linear modeling of RIP is that the speed of adjustment towards RIP following a shock is likely to be positively related to the size of the shock. Such an insight is not available in the hitherto *linear* tests for mean-reversion in real interest differentials, or Engle-Granger and Johansen cointegration tests between domestic and foreign real interest rates. Furthermore, if real interest differentials can be characterised by non-linearities, then linear approaches to nonlinear problems of monetary policy and integration would be inappropriate, though non-linear modeling poses forecasting difficulties of its own which derive from inaccuracies in the initial values and the stability of parameter values.² Whether non-linearity can be a complicating factor from the perspective of policymakers in attempts to alter real economic activity or coordinate economic integration will depend indeed on the ability of econometric models to describe the dynamics of real interest rate adjustments.

Testing for non-linearity is based on the Smooth Transition Autoregression (STAR) methodology advocated by Granger and Terasvirta (1993). While GARCH modeling is useful in describing volatility clustering, focus is made here on the use of STAR models to test for the possible existence of non-linearity in the form of Markov regimes in real interest differentials, which allows for smooth transition in the adjustment dynamics. The application of two variants of STAR modeling- logistic smooth transition autoregression (LSTAR) models and exponential smooth transition autoregression (ESTAR) models- enables us to explore the proposition that adjustments and alignments to RIP are non-linear. To the knowledge of the authors, there are no empirical studies applying STAR models to Asian real interest differentials.³

The paper is organized as follows. The following section discusses the relevant literature on real interest parity. The third section discusses the data and econometric methodology. The fourth section reports and analyses the results. The final section concludes.

I. Real Interest Parity

The attainment of *ex ante* RIP requires both uncovered interest parity (UIP) and *ex ante* relative PPP. The relationship between domestic and foreign interest rates on appropriate financial assets with the same maturities can be expressed as

$$s_{t+1}^e - s_t = i_t - i_t^* \quad (1)$$

where s is the natural logarithm of the spot exchange rate (domestic price of foreign currency), i is the nominal interest rate, superscript e and $*$ respectively refer to an expected and foreign value. *Ex ante* PPP requires

$$s_{t+1}^e - s_t = \Delta p_{t+1}^e - \Delta p_{t+1}^{*e} \quad (2)$$

where p is the natural logarithm of the price level. The familiar Fisher closed conditions expressing changes in exchange rates for the home and foreign country as functions of perfectly foreseen changes in inflation can be written as

$$r_t = i_t - \Delta p_{t+1}^e \quad (3)$$

$$r_t^* = i_t^* - \Delta p_{t+1}^{*e} \quad (4)$$

where r denotes the real interest rate. RIP in the *ex ante* representation is obtained by combining (1) and (2) along with Fisher closed conditions (3) and (4) as

$$r_{t+1}^e = r_{t+1}^{*e} \quad (5)$$

The deviations from *ex ante* RIP may be denoted as v where

$$r_{t+1}^e - r_{t+1}^{*e} = v_{t+1} \quad (6)$$

When expectations concerning inflation are assumed to be formed rationally,

$$\Delta p_{t+1} = \Delta p_{t+1}^e + \omega_{t+1} \quad (7a)$$

$$\Delta p_{t+1}^* = \Delta p_{t+1}^{*e} + \omega_{t+1}^* \quad (7b)$$

where ω is the forecast error that is serially uncorrelated with a zero mean. The *ex post* RIP condition can be written as

$$r_{t+1} - r_{t+1}^* = \pi_{t+1} \quad (8)$$

where π , which constitutes the deviation from *ex post* RIP, is a composite term incorporating v and $\omega - \omega^*$. If both domestic and foreign *ex post* real interest rates are non-stationary but π is stationary then *strong RIP* (*ex post* real interest rate equality) holds within a long-run cointegrating relationship. Strong RIP can be violated because of transactions costs, non-traded goods, non-zero foreign exchange premia, differential national tax rates and so on. It is therefore appropriate to also define *weak RIP* where domestic and foreign *ex post* real interest rates are cointegrated as

$$r_{t+1} = \lambda + \psi r_{t+1}^* + x'_{t+1} \quad (9)$$

where $\lambda \neq 0$ and/or $\psi \neq 1$ and x'_{t+1} is $I(0)$. If x'_{t+1} is $I(1)$, however, then no long-run relationship exists.

Early studies of RIP include Mishkin (1984) and Cumby and Mishkin (1986) who employ classic OLS regression analysis and find evidence against strong RIP. More recent studies that utilize the cointegration methodology include Goodwin and Grennes (1994) and Moosa and Bhatti (1996a) who find support for weak RIP for various OECD countries. However, Moosa and Bhatti (1996b) and Wu and Chen (1998) find that a series of alternative unit root tests that are more powerful than the conventional ADF tests leads to the rejection of non-stationarity. Fountas and Wu

(1999) find evidence in favor of strong RIP in the European Union member countries using unit root tests that allow for structural breaks in the series.

Other methods of investigation include Fraser and Taylor (1990) and MacDonald and Taylor (1989) who evaluate RIP by analyzing the ability of nominal interest differentials to predict future inflation differentials and conclude for the rejection of RIP. Chinn and Frankel (1995) also provide evidence that with a few exceptions, RIP does not hold for Pacific Rim rates. While many studies assume rational expectations and work in *ex post* terms, Cumby and Mishkin (1986) and Cavaglia (1992) work in *ex ante* terms having derived expectations on the basis of a subset of relevant contemporaneous data. Al Awad and Goodwin (1998) use *ex ante* measures of expected inflation in the computation of real interest rates for the purposes of conducting cointegration tests whereas Phylaktis (1999) applies Granger causality and impulse response analyses to examine the lead-lag relationship and assess the speed of adjustment of real interest rates following exogenous shocks to long-run equilibrium. Cavaglia (1992) uses time-varying parameters to find evidence of real interest rate convergence among OECD countries. In an *ex ante* expectational approach similar to Al Awad and Goodwin (1998), Mancuso, Goodwin and Grennes (2002) consider the use of threshold autoregression and nonparametric regression modeling to analyze the real interest rate linkages among a group of OECD countries. In contrast to these studies, our approach is based on STAR modeling, which aids at understanding variants of regime-dependent non-linearity.

II. Methodology and Data

The above mentioned studies by Iannizzotto (2001), McMillan and Speight (2001), Serletis and Gogas (2000), Sarno (2000a, 2000b), Sarantis (1999) and Michael *et al.*

(1997) of exchange rates imply non-linearities with distinct characteristics associated with different real exchange rate regimes. A variety of empirical models have been developed to capture these regime-dependent properties. The main approaches to modeling such non-linearities include the Markov regime-switching models, where the switch between regimes is described by a probabilistic function (see Hamilton (1989) and others), and the threshold class of models where the regime switch can be triggered by any variable, especially past values as far as parity is concerned (see, for example, Tsay (1989), Tong (1990)). Both these classes of model imply that the economy must be within a single regime in each time period where there is a sharp switch between regimes. Alternatively, there are the models based on a smooth transition generalization of threshold class (see, for example, Granger and Terasvirta (1993), Terasvirta (1994), Terasvirta and Anderson (1992)). These models allow for the possibility that deviations from *ex post* RIP occur in some intermediate state between regimes where the nature of adjustment varies with the extent of deviation from equilibrium.

The smooth transition methodology is followed in this paper for various reasons. The smoothness of adjustment between regimes is estimated and one can judge the sharpness of switching from one regime to another. One could certainly argue that the threshold class of models is more appropriate in the case of financial market arbitrage. For example, Balke and Wohar (1998) find non-linearities in the adjustment towards covered interest parity (CIP) between the UK and U.S. where the non-linearity is dependant on the size of the shock to CIP in relation to the transactions cost bandwidth. Mancuso, Goodwin and Grennes (2002) also find that generally the larger the magnitude of the shock, the faster the adjustment. The methodology followed in these studies is based on threshold autoregression on the

grounds that arbitrage towards CIP will suddenly occur once the transactions cost band is breached. RIP, however, is not based only on conditions prevailing in financial markets but on goods market adjustment as well. The extent of threshold versus STAR models will depend on the relative importance of forces governing the adjustment process in both markets. The overall nonlinearity is likely to depend on which forces are the more dominate in the determination of real interest rates and to differ across countries. Whereas adjustments to deviations in financial markets are rapid, they are as noted by Dumas (1992), rather sluggish, gradual and costly in the goods markets. In the case of goods market arbitrage, there might exist a series of thresholds straddling the equilibrium value of PPP so that as one moves further away from central parity, more and more arbitrage opportunities arise against a background of transactions costs (see, *inter alia*, Obstfeld and Taylor (1997)). The rationale behind the use of STAR methodology over the threshold class of models derives from the fact that it can be viewed as 'nesting' any threshold adjustment through the explicit estimation of the smoothness of adjustment parameter. Furthermore, whereas threshold approach may be appropriate for models with single agent and fixed transactions costs, smooth transition modeling describes better the adjustments with different transactions costs and multiple agents.

It is important for the purposes of model estimation, that the real interest differential π_t is stationary. Following Granger and Terasvirta (1993), a smooth transition autoregressive (STAR) model of order k , for π_t has the following specification

$$\pi_t = \beta_0 + \beta_1'x_t + (\theta_0 + \theta_1'x_t)F(\pi_{t-d}) + w_t \quad (10)$$

where $x_t = (\pi_{t-1}, \pi_{t-2}, \dots, \pi_{t-k})$, $\beta_1 = (\beta_1, \beta_2, \dots, \beta_k)'$, $\theta_1 = (\theta_1, \theta_2, \dots, \theta_k)'$, $w_t \sim iid N(0, \sigma^2)$, $F(\cdot)$ is the continuous transition function, π_{t-d} is the switching variable, and d is the delay parameter. The differentiable $F(\cdot)$ is a monotonically increasing function with $F(-\infty) = 0$ and $F(\infty) = 1$ which yields a non-linear asymmetric adjustment. Consider the following differentiable LSTAR function

$$F(\pi_{t-d}) = \{1 + \exp[-\gamma(\pi_{t-d} - \mu)]\}^{-1} \quad (11)$$

where γ measures the smoothness of transition from one regime to another and μ is some threshold value for π that indicates the halfway point between the two regimes. The LSTAR model assumes that different regimes may have different dynamics with the speed of adjustment varying with the extent of the deviation from equilibrium. The transition function of LSTAR is S-shaped around μ , monotonically increasing in π_{t-d} yielding an asymmetric adjustment toward equilibrium in the model. Moreover, $F(\cdot) \rightarrow 0$ as $\pi_{t-d} \rightarrow -\infty$ and $F(\cdot) \rightarrow 1$ as $\pi_{t-d} \rightarrow +\infty$ thus $F(\cdot)$ is bounded between 0 and 1 where $F(\cdot) = 0.5$ if $\pi_{t-d} = \mu$. The smaller is γ , the smoother is the transition. In the extreme, $\gamma = 0$ means that $F(\cdot)$ becomes a constant and thus (10) reduces to a linear model, $F(\cdot) = 0.5$ at all times. On the other hand, as $\gamma \rightarrow \infty$ there is an ever sharper transition at $\pi_{t-d} = \mu$ where $F(\cdot)$ jumps from 0 to 1. In this latter case, (11) becomes the usual threshold transition model along the lines of Tong (1983). Whereas the logistic distribution closely approximates the bell-shaped normal, the alternative exponential (ESTAR) transition function, symmetric and U-shaped around μ , can also be considered as in Terasvirta and Anderson (1992)

$$F(\pi_{t-d}) = 1 - \exp\{-\gamma(\pi_{t-d} - \mu)^2\} \quad (12)$$

where, as before, γ measures the speed of transition from one regime to another and μ is some threshold value for π which indicates the halfway point between the two regimes. The ESTAR function in (12) defines a transition function about μ where $F(\cdot)$ is still bounded between 0 and 1. The differences between STAR models are reflective of discrepancies in the reaction of agents to shocks of opposite signs. ESTAR models imply a symmetric U-shaped response of the real interest rate differential about some threshold 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 reaction of agents to shocks of opposite signs.

The initial testing for the presence of non-linearities in π is based on three stages. First, a linear AR model for π is specified in order to determine the lag length k . The lag length selection is based on the Schwarz information criterion and Ljung-Box statistic for autocorrelation. The residuals are saved from the chosen AR model and denoted as v . Second, having determined k , the next stage is to test for the presence of non-linearities. This is achieved through the estimation of

$$v_t = \beta_0 + \beta'_1 x_t + \beta'_2 x_t \pi_{t-d} + \beta'_3 x_t \pi_{t-d}^2 + \beta'_4 x_t \pi_{t-d}^3 + w_t \quad (13)$$

where the basic linearity test is on the null $H_0: \beta'_2 = \beta'_3 = \beta'_4 = 0$. Equation (13) is estimated across a range of values for d where the lowest p -value attached to the linearity test determines d in the later estimation of (10). The third stage of the non-linearity test is to see which smooth transition model -LSTAR or ESTAR- is appropriate for the real interest differentials. For this purpose, the following null hypotheses are considered.

$$H_{04}: \beta'_4 = 0 \quad (14)$$

$$H_{03}: \beta'_3 = 0 / \beta'_4 = 0 \quad (15)$$

$$H_{02}: \beta'_2 = 0 / \beta'_4 = \beta'_3 = 0 \quad (16)$$

One possible approach to the identification of the appropriate STAR model is to run the following sequence of nested tests. The LSTAR model is selected if H_{04} is rejected. In the alternative, the ESTAR model is adopted if in addition, H_{03} is rejected. Accepting H_{04} and H_{03} but rejecting H_{02} implies selecting the LSTAR model. Having selected the form of appropriate model, this study considers the value of γ described in (11) and (12). There is evidence of linearity when $H_0: \beta'_2 = \beta'_3 = \beta'_4 = 0$ cannot be rejected. Caution in the implementation of this approach is warranted because higher-order terms in the Taylor expansion are not taken into account in the derivation of these tests as shown by Granger and Terasvirta (1993) and Terasvirta (1994). It is suggested, as in Sarantis (1999) as well, that the p -values for each of these F tests are computed and the choice of STAR model be made on the basis of the lowest p -value. If the rejection of H_{04} or H_{02} is accompanied by the lowest p -value then the LSTAR model is chosen. If the rejection of H_{03} is accompanied by the lowest p -value then the ESTAR model is chosen. In either case, the appropriate STAR model with the speed of transition parameter γ described in (11) and (12) is estimated through non-linear least squares regressions.

This study employs monthly end-of-month *International Financial Statistics* data for six countries: Japan, Korea, Malaysia, Singapore, Thailand and the U.S. covering the period January 1977 through March 2000. The onshore nominal interest rate data are three-month deposit rates. For the purposes of macroeconomic policy, it is the setting of domestic interest rates that is of most importance. Employing offshore rates in this study would prevent capital controls from exerting influence (if any) on the assessment of RIP. Data for annual inflation (Δp) are constructed from the

consumer price indices for each country and real interest differentials are defined with respect to Japan and the U.S. Though this is not the primary objective of the paper, performing parallel tests with respect to Japan and U.S. aids in assessing the relative degree of Asian financial integration with these countries. The existing literature including Chinn and Frankel (1995) and Phylaktis (1995, 1999) suggests that while there is continued integration with U.S., there is also a growing sphere of influence for Japanese interest rates over time.

Following (7a) and (7b), rational expectations are assumed where inflation expected one year ahead is measured by actual inflation one year ahead plus a random forecast error. Use of the term structure of interest rates to extract information on expected inflation is ruled out because this study is concerned with onshore rates where data availability rules out consistent runs of monthly data for the desired maturity range over the entire study period. This same comment also applies to data for index-linked securities. Furthermore, evidence on the information content of the term structure on inflation is mixed. For example, Mishkin (1991) analyses offshore rates for a sample of ten OECD economies and finds that the shorter maturity term structure does not contain a great deal of information about inflation.⁴

III. Results

Table I reports some estimates of the distributional moments and univariate ADF unit root tests on real interest rates and differentials with respect to the U.S. and Japan. Real interest rates are on average highest in Thailand and lowest in Japan. But so is volatility too. There is some evidence that departures of real interest rates from zero are likely to be associated with increases in volatility. Judging solely by the magnitude of average real interest differentials, financial integration in Asia seems to

be more pronounced with respect to the U.S. than Japan, the mean differentials with the latter being typically higher.

However, given the potential exhibited by Figure 1 for opposite signs in *ex post* U.S. differentials to net out over the long run, the result is not conclusive. Stable real differentials followed the volatile negative values in the early 1980s. The removal or relaxation of restrictions on international capital mobility and the liberalization of domestic financial systems played a role in determining the behavior of differentials over time. In fact, there is no synchronous implementation of liberalization programs across these Asian countries. Japan and Malaysia implemented measures of financial liberalization in the late 1970s following Singapore's earlier endeavors. It was towards the end of the 1980s that Korea started its liberalization programs but remained the least liberalized whereas liberalization efforts in Thailand were initiated only in 1990. The likelihood of arbitrage opportunities increases during periods of regime change but their realization remains dependent on transactions costs. But, the surge of Korea's and Thailand's real interest differentials to highest levels during the 1997-8 Asian financial crisis is more likely to be reflective of increased default risk premium. Shifts across positive and negative regimes of real interest differentials seem to be more likely to occur with respect to U.S. than Japan. Figure 2 reveals indeed the stronger tendency for interest differentials with respect to Japan to be positive over long stretches of the sample period. This may be reflective among others, of the tradition for substantially lower real interest rates in Japan compared to international levels.⁵

Table I

The results in Table I indicate also that with the exception of Korea, the null of non-stationarity is accepted for all real interest rates at the 5% significance level. In contrast, the results for *ex post* real interest differentials π_t indicate that the null of

non-stationarity is rejected at the 5% level in all cases except Thailand vis-à-vis Japan. Overall, these results may be interpreted as evidence in favor of strong RIP. They may be seen in the context of Moosa and Bhatti (1996b) who employ conventional and more sophisticated unit root tests on quarterly *ex ante* real interest differentials over the period 1980-94 and find evidence in favor of RIP for six Asian countries vis-à-vis Japan on the basis of strongly mean-reverting differentials.

Table II

Table II reports the tests for non-linearities in the real interest differential series, π_t together with the Ljung-Box statistic for serial correlation in the residuals. Following the selection of the lag length k for each AR process, the delay parameter d is constrained to be $1 \leq d \leq (D = 8)$. The *ad hoc* setting of the maximum delay to eight months leaves room for the sluggish adjustment in goods markets to take place in addition to the much faster corrections in financial markets. It remains close to the range implied by the half-lives of deviations running into several months in studies by Nakagawa's (2002) and Mancuso, Goodwin and Grennes (2002). The value of the delay parameter d^* that minimizes the p -value associated with H_0 in the auxiliary regression (13) is lowest (highest) for Thailand (Malaysia) with respect to both the U.S. and Japan. Given the maximum lag k and the delay parameter d^* , serial correlation tests indicate that the residuals are free from serial correlation. The minimum p -value estimates suggest that the null of linearity can be rejected at the 5% significance level for all cases except the Singapore-Japan differential for which the null is rejected at the 10% level.

Table III

As the type of non-linear dependency in the data is *a priori* undetermined, further specification tests are warranted. Table III reports the test results for the specific form of non-linearity present in real interest differentials. The results of hypothesis testing outlined in Equations (14)-(16), indicate that the LSTAR model is

the more adequate non-linear fit in all cases except Malaysia-U.S. and Thailand-Japan. The transition process governed by the logistic and exponential forms is a smooth and continuous function of the information set on lagged disturbances in real interest differentials. The LSTAR modeling prevailing for most countries suggests that the dynamics of real interest differential regimes are different. Moreover, the speed of adjustment towards parity is likely to depend not only on the magnitude of the shock but on its sign as well. The asymmetric adjustment might be attributable to the nature of goods market adjustment. For example, a negative shock to RIP may result from an increase in Asian inflationary pressures. It may be the case that the presence of Asian menu costs inhibits a required downward adjustment of Asian inflation. The ESTAR modeling of real interest differentials for Malaysia-U.S. and Thailand-Japan suggests in contrast, that regimes are likely to have similar dynamics but the transition properties may be different and duration-dependent. In these cases, the sign of the shock to RIP would not matter in particular when goods market adjustment is equally affected by the presence of menu costs at home and abroad.

Having determined the nature of the STAR model that is appropriate for each of the series, we may now turn our attention to non-linear estimation. In line with earlier studies, the ESTAR and LSTAR models are scaled using the variance σ_π^2 and standard deviation σ_π , respectively. As well as assisting convergence during estimation, this normalizes the deviations in the switching variable and facilitates interpretation of the smoothness parameter. Thus (11) and (12) may be rewritten as

$$F(\pi_{t-d}) = [1 + \exp\{-\gamma(1/\sigma_\pi)(\pi_{t-d} - \mu)\}]^{-1} \quad (17)$$

$$F(\pi_{t-d}) = 1 - \exp\{-\gamma(1/\sigma_\pi^2)(\pi_{t-d} - \mu)^2\} \quad (18)$$

Table IV reports estimates of the transition parameters and diagnostic tests. These statistics are derived from the non-linear least squares estimates of (10). The

estimates of the standard error of the non-linear regression suggest that the average variation of observed differentials around the regression line is not substantial. Except for Singapore-Japan differentials, the p-values associated with the Ljung-Box statistics generally indicate the absence of ARCH effects in the residuals of STAR models. The results indicate that in all cases γ takes the anticipated sign and with the exception of Japan-U.S. differential and the Korean relationships with both countries, it appears to be significant at the 10% level. Terasvirta (1994) and Sarantis (1999) point to the usual difficulty in the precise estimation of γ . However, Sarno (2000b) argues that the statistical significance of γ is in a sense not questionable because linearity has already been rejected in the earlier tests.

It is also clear from the examination of LSTAR model estimates that the estimated γ values are varied. Small values for γ are indicative of a very slow and smooth transition from one regime to another. On the other hand, the Singapore-Japan differential exhibits a much larger value of γ , which implies sharper and more abrupt transitions. It would be useful here to comment on what the values of γ actually mean. Let us designate $F(\pi_{t-d})=0$ and $F(\pi_{t-d})=1$ as regimes of a pure “positive real interest differentials” and “negative real interest differentials.” In the case of the Malaysia-Japan real interest differential, for instance, the LSTAR model estimate of $\gamma=0.123$ means that a one standard deviation positive shock to π_{t-8} yields $F(\pi_{t-8})=0.594$. The new regime is therefore a linear combination of regimes 1 and 2 with the weights [0.594,0.406]. In the case of a two standard deviation shock to π_{t-8} , we have $F(\pi_{t-8})=0.623$ and so these weights become [0.623,0.377]. There is therefore a larger leaning towards $F(\pi_{t-8})=0$ on account of the larger shock. The much higher values for γ exhibited in Table IV mean that even minute deviations of

the switching variable from the threshold level can place real interest differentials entirely in one regime or the other and cause the transition function $F(\pi_{t-d})$ to traverse the interval (0,1) rather very quickly. In cases where γ is significant at the 10% level, Figure 3 plots some of the estimated logistic transition functions against the appropriate lagged values of real interest differentials π_{t-d} to reveal the anticipated asymmetries and discrepancies in curvature across countries. The real interest differential between Singapore and Japan, which features the largest value for γ , features a sharp switch between regimes whereas the Singapore-U.S. differential, which has a much smaller value for γ , demonstrates the much smoother adjustment.

Table IV also reports estimates of the halfway point μ between the two regimes. Given the scaling procedure described in Equation (17), the data for μ^* are ‘de-scaled’ estimates of the halfway point where μ has been divided by the standard deviation σ_π and variance σ_π^2 for LSTAR and ESTAR model estimates, respectively. There is little evidence that these estimates approach the sample means of real interest differentials, thereby diminishing the likelihood for observations to lie with equal probability on either side of the logistic transition function (as demonstrated in Figure 3). The logistic functions for the Malaysia-Japan and Singapore-Japan differentials indicate also a similar tendency for the transition probability to rise as past differentials increase. Such results imply that the larger the positive departure from RIP, the higher the likelihood of regime switching and reversion towards equilibrium. However, the fact that positive differentials are more likely to trigger regime switches than negative observations lends support to the proposition of asymmetric adjustment to deviations from RIP.

The only instances in which ESTAR modeling is most appropriate concern the Malaysia-U.S. and Thailand-Japan differentials. Figure 4 depicts the estimated exponential function for each relationship. The greater value for γ in the case of Thailand-Japan leads to a sharper transition function and is reflective of a faster speed in regime switching and mean reversion. In both cases, μ^* is higher than the sample mean. However, the different signs attached to μ^* suggest that smaller positive shocks to parity are able to generate non-linearities in the case of Malaysia-U.S. than is the case for Thailand-Japan.

A final analysis of our results involves an assessment of the ability of the estimated STAR models to forecast beyond certain structural breakpoints. The breakpoints are determined endogenously using Perron (1997) unit root tests with observations from the subperiod September 1994 through March 2000 centered on July-1997. Thus, the extended period does not force breakpoints to fall after the onset of the Asian financial crisis. It leaves room for judgement on the adequacy of nonlinear modeling to capture significant developments in real interest parity either before or after the crisis.⁶

The results indicate that with the exception of Japan, the endogenous structural breaks for US-related differentials fall within a very narrow range (March to December 1997) from the critical date. It is with respect to Japan-related differentials that breakpoints seem to occur within a much larger window. With the exception of the narrow rejection in the case of Thailand-Japan, the traditional Chow second test performed on the basis of these breakpoints indicates that the null of predictive adequacy is accepted at the 5% significance level. This interesting evidence suggests that the non-linear modeling can describe significant developments in real interest parity over time, providing useful signals and tracing back the onset of crises. Not

only does it reinforce the view that non-linearities can be found in the behavior of real interest differentials in South East Asian economies, it also shows that for policy-making purposes, smooth transition models provide useful tools in recovering information on monetary and financial crises.

5. Summary and Conclusion

This is the first study that has investigated the possibility of non-linearities in the adjustment of South East Asian real interest rates towards real interest rates prevailing in Japan and the U.S. The application of smooth transition autoregressive modeling to South East Asian real interest differentials confirms the presence of non-linear adjustment in most cases. Furthermore, there is considerable variation in the smoothness of adjustment from one interest rate regime to another across countries. Such behavior would not be detected in the hitherto linear tests for unit root processes.

The implications of our findings are fourfold. First, large shocks to real interest parity are more likely to lead to the reestablishment of parity at a faster rate than small shocks. Second, the macroeconomic models of exchange rate behavior that imply real interest parity may find more applicability in explaining a world dominated by large rather than small shocks to parity. Third, the evidence in favor of non-linear stochastic dynamics should be useful in understanding the complexities of economic integration and monetary crises. Fourth, because adjustment of real interest rates incorporates both goods market and financial market behavior, this analysis is useful in bridging the gap between earlier studies that have focussed on non-linearities in the context of either goods market or financial market adjustment.

Although it is argued that the smooth transition modeling is adequate approach to the analysis of the dynamics of Asian real interest parity, very sharp transitions

from one regime to another are also reported in some cases. This leaves many avenues for future research. It would be interesting to consider other forms of non-linearities and deterministic chaotic dynamics, including other variants in the class of non-linear state-dependent models such as the threshold and exponential autoregression models. Such approaches can certainly shed more light on the dynamics of economic integration.

References

- Al Awad, M. and Goodwin, B. K. (1998) 'dynamic Linkages among Real Interest Rates in International Capital Markets', *Journal of International Money and Finance*, 17, 881-907.
- Balke, N., and Wohar, M. (1998), 'Nonlinear Dynamics and Covered Interest Parity', *Empirical Economics*, 23, 535-59.
- Cavaglia, S. (1992), 'The Persistence of Real Interest Differentials: a Kalman Filtering Approach', *Journal of Monetary Economics*, 29, 429-43.
- Chinn, M.D. and Frankel, J.A. (1995), 'Who Drives Real Interest Rates around the Pacific Rim: the USA or Japan?' *Journal of International Money and Finance*, 14, 801-821.
- Cumby, R. and Mishkin, F. (1986), 'The International Linkage of Real Interest Rates: the European-US Connection', *Journal of International Money and Finance*, 5, 5-23.
- de Brouwer, G. (1999), *Financial Integration in East Asia*, Cambridge, UK: Cambridge University Press.
- Dumas, B. (1992), 'Dynamic Equilibrium and the Real Exchange Rate in a Spatially Separated World,' *The review of Financial Studies*, 5, 153-180.
- Fountas, S. and Wu, J. (1999), 'Testing for Real Interest Rate Convergence in European Countries', *Scottish Journal of Political Economy*, 46, 158-74.
- Frankel, J. (1979), 'On the Mark: a Theory of Floating Exchange Rates based on Real Interest Differentials', *American Economic Review*, 69, 610-22.
- Fraser, P. and Taylor, M. (1990), 'Some Efficient Tests of International Real Interest Parity', *Applied Economics*, 22, 1083-92.
- Frenkel, J. (1976), 'A Monetary Approach to the Exchange Rate: Doctrinal Aspects of Empirical Evidence', *Scandinavian Journal of Economics*, 78, 200-24.
- Fuller, W.A. (1976), *Introduction to Statistical Time Series*, New York: John Wiley.
- Goodwin, B. and Grennes, T. (1994), 'Real Interest Rate Equalization and the Integration of International Financial Markets', *Journal of International Money and Finance*, 13, 107-24.
- Granger, C. and Terasvirta, T. (1993), *Modeling Nonlinear Economic Relationships*, New York: Oxford University Press.
- Hamilton, J.D. (1989), 'A New Approach to the Economic Analysis of Nonstationary Time Series and the Business Cycle', *Econometrica*, 57, 357-384.

- Iannizzotto, M. (2001), 'Exchange Rate Misalignment and Nonlinear Convergence to Purchasing Power Parity in European Exchange Rate Mechanism', *Applied Financial Economics*, 11, 511-26.
- Kaminsky, G. L., and Schmukler S. L. (1999), 'What Triggers Market Jitters? A Chronicle of the Asian Crisis', *Journal of International Money and Finance*, 18, 537-560.
- Lane, P.R. and Milesi-Ferretti, G.M. (2001), 'Long-Term Capital Movements', Working Paper No. 8366, July, NBER.
- Leybourne S. and Mizen, P. (1997), 'Disinflation and Central Bank Independence in Australia, Canada and New Zealand: Evidence from Smooth Transition Analysis', Discussion Paper 97/6, Department of Economics, University of Nottingham.
- MacDonald, R. and Taylor, M.P. (1989), 'Interest Rate Parity: some New Evidence', *Bulletin of Economic Research*, 41, 255-74.
- Mancuso, A.J., Goodwin, B.K. and Grennes, T.J. (2002), 'Nonlinear Aspects of Capital Market Integration and Real Interest Rate Equalization', *International Review of Economics and Finance*, forthcoming.
- McMillan, D. and Speight, A. (2001), 'Non-Linearities in the Black Market Zloty-Dollar Exchange Rate: some Further Evidence', *Scottish Journal of Political Economy*, 11, 209-20.
- Michael, P., Nobay, A.R. and Peel, D. (1997), 'Transactions Costs and Nonlinear Adjustments in Real Exchange Rates: an Empirical Investigation', *Journal of Political Economy*, 105, 862-79.
- Mills, T. (1995), 'Business Cycle Asymmetries and Non-Linearities in UK Macroeconomic Time-series', *Ricerche Economiche*, 49, 97-24.
- Mishkin, F. (1984), 'Are Real Interest Rates Equal across Countries: an Empirical Investigation of International Parity Conditions', *Journal of Finance*, 39, 1345-57.
- Mishkin, F. (1991), 'A Multi-Country Study of the Information in the Shorter Maturity Term Structure about Future Inflation', *Journal of International Money and Finance*, 10, 2-22.
- Mishkin, F. (1999), 'Lessons from the Asian Crisis', *Journal of International Money and Finance*, 18, 709-723.
- Moosa, I. and Bhatti, R. (1996a), 'Some Evidence on Mean Reversion in *ex ante* Real Interest Rates', *Scottish Journal of Political Economy*, 43, 177-91.

- Moosa, I. and Bhatti, R. (1996b), 'Does Real Interest Parity Hold? Empirical Evidence from Asia', *Keio Economic Studies*, 33, 63-70.
- Mullineux, A. and Peng, W. (1993), 'Nonlinear Business Cycle Modeling', *Journal of Economic Surveys*, 7, 41-83.
- Mussa, M. (1976), 'The Exchange Rate, Balance of Payments, and Monetary and Fiscal Policy under a Regime of Controlled Floating', *Scandinavian Journal of Economics*, 78, 229-48.
- Nakagawa, H., 2002, 'Real Exchange rates and real interest differentials: implications of nonlinear adjustment in real exchange rates', *Journal of Monetary Economics*, 49, 629-649.
- Obstfeld, M. and Taylor, A.M. (1997), 'Nonlinear Aspects of Goods Market Arbitrage and Adjustment: Heckscher's Commodity Points Revisited', Discussion Paper No. 1672, CEPR.
- Ocal, N. and Osborn, D. (2000), 'Business Cycle Non-Linearities in UK Consumption and Production', *Journal of Applied Econometrics*, 15, 27-43.
- Perron, P., 'Further Evidence from Breaking Trend Functions in Macroeconomic Variables', *Journal of Econometrics*, 80, 1997, 355-385.
- Phylaktis, K. (1995), 'Capital Market Integration in the Pacific Basin Region: an Analysis of Real Interest Rate Linkages', IMF Working Paper, 95/133.
- Phylaktis, K. (1997), 'Capital Market Integration in the Pacific Basin Region: an Analysis of Real Interest Rate Linkages', *Pacific-Basin Finance Journal*, 5, 195-213.
- Phylaktis, K. (1999), 'Capital Market Integration in the Pacific Basin Region: an Impulse Response Analysis', *Journal of International Money and Finance*, 18, 267-287.
- Sarantis, N. (1999), 'Modeling Non-Linearities in Real Effective Exchange Rates', *Journal of International Money and Finance*, 18, 27-45.
- Sarno, L. (2000a), 'Systematic Sampling and Real Exchange Rates', *Weltwirtschaftliches Archiv*, 136, 24-57.
- Sarno, L. (2000b), 'Real Exchange Rate Behavior in High Inflation Countries: Empirical Evidence from Turkey, 1980-97', *Applied Economics Letters*, 7, 285-91.
- Serletis, A. and Gogas, P. (2000), 'Purchasing Power Parity, Non-Linearity and Chaos', *Applied Financial Economics*, 10, 615-22.

- Skalin, J. and Terasvirta, T. (1999), 'Another Look at Swedish Business Cycles, 1861-1988', *Journal of Applied Econometrics*, 14, 359-378.
- Taylor, A. M., (2001), 'Potential Pitfalls for the Purchasing-Power-Parity Puzzle? Sampling and Specification Biases in Mean-Reversion Tests of the Law of One Price', *Econometrica*, 69, 473-498.
- Terasvirta, T. (1994), 'Specification, Estimation, and Evaluation of Smooth Transition Autoregressive Models', *Journal of American Statistic Association*, 89, 281-12.
- Terasvirta, T. and Anderson, H. M. (1992), 'Characterizing Nonlinearities in Business Cycles using Smooth Transition Autoregressive Models', *Journal of Applied Econometrics*, 7, S119-36.
- Tong, H. (1990), *Threshold Models in Nonlinear Time Series Analysis*, New York: Springer-Verlag.
- Tsay, R. (1989), 'Testing and Modeling Threshold Autoregressive Processes', *Journal of the American Statistical Association*, 84, 231-40.
- Wu, J-L. and Chen, S-L (1998), 'A Re-examination of Real Interest Rate Parity', *Canadian Journal of Economics*, 31, 837-51.

Table I. ADF unit root tests on real interest rates and differentials*

Time series	Mean	Std. Dev.	ADF (no trend)	ADF (trend)
Real Interest Rates				
Korea	4.060	3.915	-3.556 ^a	-3.682 ^b
Malaysia	3.071	2.908	-2.287	-2.259
Singapore	2.199	2.704	-2.681 ^c	-2.751
Thailand	5.032	5.037	-2.852 ^c	-2.805
Japan	0.647	1.700	-2.583 ^c	-2.565
U.S.	3.545	3.315	-2.346	-2.440
Differentials vis-à-vis U.S.				
Korea	0.515	3.841	-4.072 ^a	-4.503 ^a
Malaysia	-0.474	3.269	-3.521 ^a	-3.653 ^b
Singapore	-1.346	2.055	-4.064 ^a	-4.053 ^a
Thailand	1.487	4.124	-3.181 ^b	-3.307 ^c
Japan	-2.898	3.080	-3.651 ^a	-3.741 ^b
Differentials vis-à-vis Japan				
Korea	3.413	3.242	-3.929 ^a	-4.123 ^a
Malaysia	2.423	2.636	-4.115 ^a	-4.103 ^a
Singapore	1.552	2.338	-2.981 ^b	-3.130
Thailand	4.385	4.026	-2.672 ^c	-2.593

* These are Augmented Dickey-Fuller (ADF) unit root tests conducted on the real interest rates and differentials with respect to the U.S. then Japan. The full sample period is 1977M1-2000M3 with 279 monthly observations. For each test, the lag length was chosen using the Schwarz information criterion.

^a, ^b and ^c indicate rejection of the null of non-stationarity at respectively, the 1, 5 and 10% levels of significance in the ADF tests. Relevant ADF critical values taken from Fuller (1976) are -3.51, -2.89 and -2.58, while for regressions including a trend, these are -4.04, -3.45 and -3.15 respectively. In all cases, the time trend was insignificant at the 5% level.

Table II. Tests for non-linearities*				
Differentials	k	d^*	p -value	$Q(4)$
vis-à-vis U.S.				
Korea	2	6	0.008	0.548
Malaysia	2	7	0.000	0.734
Singapore	2	5	0.014	0.207
Thailand	2	1	0.000	0.362
Japan	2	6	0.000	0.168
vis-à-vis Japan				
Korea	2	7	0.011	0.409
Malaysia	5	8	0.003	0.971
Singapore	1	3	0.076	0.450
Thailand	2	1	0.000	0.952

* These tests are based on the real interest differential with respect to the U.S. and then Japan. The null of linearity is based on Equation (12): $v_t = \beta_0 + \beta_1'x_t + \beta_2'x_t\pi_{t-d} + \beta_3'x_t\pi_{t-d}^2 + \beta_4'x_t\pi_{t-d}^3 + w_t$. The column headed 'p-value' corresponds to the test where the null is linearity $H_0: \beta_2' = \beta_3' = \beta_4' = 0$. It should be noted that the Schwarz criterion is used to determine lag length k of AR process. The residuals from AR processes were then saved. Having determined k , a range of delay parameters d (d is between 1 and $D = 8$) were employed. The value of d^* chosen is that which gives rise to the lowest p -value of the linearity test using the data for the residuals of the AR process. The linearity test is itself a variable-deletion F test on the restriction applied to Equation (12). The column headed $Q(4)$ refers to the p -value associated with the Ljung-Box statistic for serial correlation among the residuals.

Table III. Specification of the non-linear model*

Real Interest differentials	k	d*	H₀₄	H₀₃	H₀₂	Model Type
vis-à-vis U.S.						
Korea	2	6	0.002*	0.787	0.119	LSTAR
Malaysia	2	7	0.039	0.005*	0.000	ESTAR
Singapore	2	5	0.406	0.077	0.010*	LSTAR
Thailand	2	1	0.136	0.042	0.000*	LSTAR
Japan	2	6	0.002*	0.009	0.032	LSTAR
vis-à-vis Japan						
Korea	2	7	0.001*	0.975	0.270	LSTAR
Malaysia	5	8	0.005*	0.142	0.092	LSTAR
Singapore	1	3	0.049*	0.831	0.086	LSTAR
Thailand	2	1	0.006	0.003*	0.028	ESTAR

* The lag length k is determined using Schwarz criterion and Ljung-Box statistic for autocorrelation.

The value of d^* is that of the delay parameter which gives rise to the lowest p -value of the linearity test using the data for the residuals of the AR process. In the nested hypothesis testing, the rejection of

$H_{04}: \beta'_4 = 0$ in Equation (12) results in LSTAR selection; the acceptance of H_{04} and rejection of

$H_{03}: \beta'_3 = 0 / \beta'_4 = 0$ in Equation (14) implies ESTAR selection and the acceptance of both H_{04} and

H_{03} , combined with the rejection of $H_{02}: \beta'_2 = 0 / \beta'_4 = \beta'_3 = 0$ in Equation (15) indicate the

appropriateness of LSTAR modeling. The asterisk denotes the lowest p -value associated with the

variable-deletion tests and therefore the determination of the relevant STAR model.

Table IV. The smoothness of adjustment*

Differentials	Model Type	Smooth γ	F-test	μ	μ^*	Q(4)	PF	Std. err
vis-à-vis U.S.								
Korea	LSTAR	0.388	0.182	-34.156	-8.892	0.627	0.425	0.182
Malaysia	ESTAR	6.902	0.000	-0.669	-0.205	0.599	0.845	0.975
Singapore	LSTAR	1.438	0.002	-3.178	-1.546	0.311	0.999	0.833
Thailand	LSTAR	0.344	0.070	9.116	2.210	0.163	0.888	1.237
Japan	LSTAR	4.365	1.000	9.431	3.062	0.720	1.000	0.945
vis-à-vis Japan								
Korea	LSTAR	0.031	1.000	-16.386	-5.054	0.192	0.201	1.238
Malaysia	LSTAR	0.123	0.053	-2.091	-0.793	0.689	0.179	0.869
Singapore	LSTAR	13.920	0.000	9.913	3.933	0.012	0.615	0.756
Thailand	ESTAR	21.151	0.000	8.199	2.036	0.680	0.048	1.025

* Non-linear least squares estimation of Equation (10) is made using the Gauss-Newton method. The column headed F-test refers to the p -value associated with a variable-deletion F test on the coefficient of adjustment smoothness γ . The μ statistic represents the estimated threshold value for the switching variable (see Equations (17) and (18)) while μ^* is the estimate after “descaling”, dividing by the standard deviation (variance) for LSTAR (ESTAR) models. Q(4) refers to the p -value associated with the Ljung-Box statistic with four lags testing for serial correlation in the residuals. PF is the p -value of the Chow (second) test for predictive failure using the endogenously determined breakpoints of February 1996 (Malaysia-Japan, Japan-US), March 1997 (Singapore-US), May 1997 (Thailand-Japan), October 1997 (Korea-US, Korea-Japan), November 1997 (Thailand-US), December 1997 (Malaysia-US), and August 1998 (Singapore-Japan). Each of these follows an F -distribution on the null of predictive adequacy. $St. err$ is the standard error of the non-linear regression.

**Figure 1. Real interest differentials series vis-à-vis the US
(Monthly observations January 1977 - March 2000)**

**Figure 2 Real interest differentials series vis-à-vis Japan
(Monthly observations January 1977 - March 2000)**

Figure 3. LSTAR Transition Functions

Transition Function for Malaysia-Japan Real Interest Differential

Transition Function for Singapore-Japan Real Interest Differential

Figure 3 (continued). LSTAR Transition Functions

Transition Function for Singapore-US Real Interest Differential

Transition Function for Thailand-US Real Interest Differential

Figure 4. ESTAR Transition Functions

Transition Function for Thailand-Japan Real Interest Differential

Transition Function for Malaysia-US Real Interest Differential

FOOTNOTES

¹ Further discussion can be found in Sarantis (1999) and references therein.

² Such difficulties are discussed for instance by Mullineux and Peng (1993) in the context of business cycle modelling.

³ Earlier examples of studies employing the STAR methodology include Sarantis (1999) who examines the real exchange rates of the G10 countries, Leybourne and Mizen (1997) who examine consumer prices, Mills (1995) and Ocal and Osborn (2000) who examine a range of UK macroeconomic series that includes industrial production, and Skalin and Terasvirta (1999) who examine the Swedish business cycle.

⁴ The limitations on term structure and survey data preclude the use of *ex ante* real interest rates. It is possible to predict *ex ante* rates using instrumental variables; this approach however guarantees consistency but not unbiasedness. Our recourse to *ex post* rates is founded on the assumption of rational expectations, a plausible hypothesis that is usually applied in financial markets. This is also consistent with a large literature on real interest parity.

⁵ It is noted that the mean of quarterly *ex post* real interest differentials with respect to the U.S. over the period 1980-94 are found by de Brouwer (1999) to be statistically insignificant for Malaysia and Japan. The RIP is primarily rejected through the rejection of UIP in the cases of Korea and Thailand, and PPP in the case of Singapore.

⁶ As with any other financial turmoil, the starting point of the Asian currency crisis cannot be determined with precision. The developments in Thailand culminated by the official decision under increasing speculative pressures by the monetary authorities to float the baht in July 1997 makes this date a prime candidate. Several studies refer to this event as the trigger to the Asian financial crisis including Mishkin (1999) and Kaminsky and Schmukler (1999).