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PURCHASING POWER PARITY AND THE REAL EXCHANGE RATE IN BANGLADESH: A NONLINEAR ANALYSIS

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Purchasing Power Parity and the Real Exchange Rate in Bangladesh: A Nonlinear Analysis†

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Abstract

The long-run purchasing power parity (PPP) hypothesis is examined using data for Bangladesh and its major trading partners - the US, Euro area, Japan and India - during the period 1994 to 2002. We apply recently developed nonlinear econometric techniques and provide strong evidence for highly nonlinear mean-reversion of real bilateral Bangladesh taka exchange rates toward a stable long-run equilibrium. Our findings imply strong support for the validity of long-run PPP as well as for the theoretical models which predict nonlinear adjustment in real exchange rates.

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1 Introduction

This study examines the validity of long-run PPP using data for Bangladesh and its four major trading partners during the period 1994 to 2002. The PPP hypothesis postulates that national price levels expressed in a common currency should be equal or, equivalently, that the nominal exchange rate should be proportional to the ratio of national price levels. Although PPP has variously been viewed by the international finance profession as a theory of exchange rate determination, as a short- or long-run equilibrium condition, or as an efficient arbitrage condition in either goods or asset markets, the recent literature appears to take the view that PPP is a valid long-run equilibrium condition at least in industrialized economies, which holds due to arbitrage in international goods markets (see the survey of Taylor, 1995; Froot and Rogoff, 1995; Rogoff, 1996; Sarno and Taylor, 2002; Sarno 2003). Conversely, empirical evidence on the validity of long-run PPP for developing countries is rather mixed (see, for example, Gan, 1994; Calvo, Reinhart and Vegh, 1995; Doganlar and Ozmen, 2000; Bahmani-Oskooee and Mirzai, 2000; Luintel, 2000; Yunus, 2000; Basher and Mohsin, 2004).

PPP may be legitimately tested by carrying out nonstationarity tests on the real bilateral exchange rate since the latter can be interpreted as a measure of deviation from PPP. While the real exchange rate may be subject to short-run variation, a necessary condition for PPP to hold in the long-run is that the real exchange rate be covariance stationary, and thus has a tendency to revert to a stable equilibrium level over time. In fact, nonstationarity of the real exchange rate implies invalidity of long-run PPP as the divergence of purchasing power across the countries considered (expressed in the same currency) would become theoretically infinite.

Whether long-run PPP holds or the real exchange rate is stationary has important economic implications. First, the degree of persistence in the real exchange rate can be used to infer the principal impulses driving exchange rate movements. In particular, if the real exchange rate is highly persistent (for example close to a random walk), then the shocks are likely to be supply-side, whereas if there is little persistence, then the shocks may principally be aggregate demand-based (Rogoff, 1996). Second, since the real exchange rate is commonly regarded as a measure of international competitiveness, it reflects an important policy-relevant variable particularly in developing countries, like Bangladesh, where exports have been the principal source of economic growth in the 1990s. In fact, as Bangladesh's export base is hardly diversified with textiles and clothing being the dominant export goods, Bangladesh is particularly vulnerable to increased competition from other Asian countries that produce labour intensive garments.[1] Finally, estimates of PPP exchange rates are often used for practical purposes, such as determining the degree of misalignment of the nominal exchange rate and the appropriate policy response, the setting of exchange rate parities, and the international comparison of national income levels. These practical uses of the PPP concept, and in particular the calculation of PPP exchange rates, would obviously be affected if the real exchange rate would be non mean-reverting, thereby containing a unit root.

Although the professional view on the validity of long-run PPP has shifted several times, some recent studies using nonlinear econometric methods have provided

[1] The following quotation emphasizes this point: "However, an appreciating real effective exchange rate has threatened to undermine Bangladesh's export competitiveness, particularly vis-à-vis South-East Asian garment manufacturers, and therefore constitutes a threat to future export-led growth." (World Trade Organization, 2000, p. 3).

fairly convincing evidence that deviations from PPP dissipate over time, which, in turn, implies that the real exchange rate is a stationary process with a unique long-run equilibrium level (e.g. Goldberg, Gosnell and Okunev, 1997; Michael, Nobay and Peel, 1997; Sarno, 2000; Taylor, Peel and Sarno, 2001; Sarno and Taylor, 2001). These empirical studies motivate the adoption of nonlinear econometric methods on the basis of recently developed models of real exchange rate determination under transaction costs of international arbitrage. The key idea of these theoretical models - cited in the next section - is that deviations from PPP (the real exchange rate) will be more rapidly mean reverting for larger deviations from the PPP equilibrium level, since the larger the deviations from equilibrium the greater the net benefit of arbitrage in international goods markets.

Indeed, this study provides strong evidence that deviations from PPP obtained between Bangladesh and important trading partners - the US, the Euro area, Japan and India - do not contain a statistically significant permanent component and dissipate in a nonlinear fashion, consistent with the emerging theoretical literature on nonlinear real exchange rate adjustment in the presence of international arbitrage costs.

The remainder of the paper is as follows. Section 2 discusses the theoretical rationale for nonlinear mean reversion in the real exchange rate. Section 3 presents the econometric methodology. In Section 4 we describe the data set and report the results from the empirical analysis. A final section briefly summarizes and concludes.

2 Motivating Nonlinear Dynamics in the Real Exchange Rate

Several studies have tested the validity of long-run PPP applying unit root tests to real exchange rate data based on a linear auxiliary regression. The literature applying linear unit root tests of this kind or variants of it to real exchange rate data has generally provided evidence that real exchange rates are nonstationary processes, implying the invalidity of long-run PPP (see Sarno, 2003, and the references therein).[2] However, some researchers have been able to reject the null of nonstationarity either using long spans of real exchange rate data (e.g. Kim, 1990; Diebold, Husted and Rush, 1991; Lothian and Taylor, 1996) or using panel unit root tests (e.g. Abuaf and Jorion, 1990; Wu, 1996; O'Connell, 1998; Taylor and Sarno, 1998; Breuer, McNown and Wallace, 2001) in order to increase the power of conventional unit root tests. Although some of these studies suggest that PPP may be viewed as a valid long-run international parity condition, they also report a high degree of real exchange rate persistency, which, in turn, seems difficult to reconcile with the observed high volatility of real exchange rates, therefore generating a puzzle (Rogoff, 1996).

Among the possible explanations of the violation of PPP suggested by the literature, transport costs, tariffs and non-tariff barriers play a somewhat dominant role. In fact, the custom tariff is a key instrument of Bangladesh's trade policy and reflects the government's primary source of revenue, accounting for almost one third of total taxes. In recent years Bangladesh has successfully simplified and rationalized its tariff structure by reducing the number of tariff bands and lowering the maximum tariff rate substantially. Moreover applied most-favoured-nation

[2] An exception is the study of Edison, Gagnon and Melick (1997), who provide some evidence in favour of PPP during the recent floating period for a number of countries. See also Cheung and Lai (1993a,b, 1994, 1998).

(MFN) tariffs have fallen by more than half.[3] Even though tariffs have been considerably reduced over time, non-tariff barriers - for example in the form of ad-hoc surcharges or strict inspection requirements - play a relevant role in Bangladesh.

Frictions in international arbitrage have important implications and, in particular, imply potential nonlinearities in real exchange rate dynamics, as shown by a number of authors who have developed theoretical models of real exchange rate determination under transaction costs (e.g. Benninga and Protopapadakis, 1988; Dumas, 1992; Sercu, Uppal and Van Hulle, 1995; Obstfeld and Rogoff, 2000). In most of these models, proportional or "iceberg" transport costs ("iceberg" because a fraction of goods are presumed to `melt' when shipped) create a band for the real exchange rate within which the marginal cost of arbitrage exceeds the marginal benefit. Assuming instantaneous goods arbitrage at the edges of the band then typically implies that the thresholds become reflecting barriers. In general, these models suggest that, as a result of international trade costs, small deviations from PPP may be persistent since they are left uncorrected as long as they are small relative to the costs of trading. However, deviations from PPP follow a nonlinear mean-reverting process such that the speed of adjustment towards equilibrium varies directly with the extent of the deviation from PPP. In some models the jump to mean-reverting behaviour is sudden, whilst in others it is smooth, and Teräsvirta (1994) and Dumas (1994) suggest that, even in the former case, time

[3] A MFN tariff is known as the tariff level that a member of the World Trade Organization charges on a good to other members. Hereby, imports from a country are treated on the same basis as that given to the most favoured other nation. That is, and with some exceptions, every country gets the lowest tariff that any country gets, and reductions in tariffs to one country are provided also to others.

aggregation and non-synchronous adjustment by heterogeneous agents is likely to result in smooth aggregate regime switching.

Goldberg, Gosnell and Okunev (1997) develop a model of mean reversion of exchange rates to PPP where exchange rates are assumed to follow a mean reverting random walk towards a stochastic PPP rate. The model allows for the possibility that mean reversion towards PPP is nonlinear, and regression equations consistent with the theoretical model are derived. The model is tested using data for six countries, producing evidence that the mean reversion process is not linear for some countries. Michael, Nobay and Peel (1997) and Taylor, Peel and Sarno (2001) propose an econometric modelling framework for the empirical analysis of PPP which essentially provides the empirical counterpart of the models of real exchange rate determination under transaction costs discussed above.[4] Using real exchange rate data for the recent float alone since 1973, Taylor, Peel and Sarno (2001) illustrate that four major real bilateral dollar exchange rates are well characterized by nonlinearly mean-reverting processes over the sample. Their estimated models imply an equilibrium level of the real exchange rate in the neighbourhood of which the behaviour of the log-level of the real exchange rate is close to a random walk, becoming increasingly mean reverting with the absolute size of the deviation from equilibrium, consistent with the theoretical literature on real exchange rate dynamics in the presence of international arbitrage costs.

[4] Other studies use similar techniques in the context of modelling deviations from the law of one price (see, for example, Obstfeld and Taylor, 1997; Sarno, Taylor, and Chowdhury, 2004).

3 Modelling Nonlinear Adjustment in the Real Exchange Rate

A nonlinear model that is able to capture the kind of properties required to model nonlinear mean-reversion in real exchange rates is the smooth transition autoregressive (STAR) model proposed by Granger and Teräsvirta (1993) and Teräsvirta (1994, 1998). A STAR model for the real exchange rate, q_t, may be written as

$$q_t = \phi_1' X_t + \left(\phi_2' X_t\right) G[\gamma; q_{t-d} - \mu] + \varepsilon_t, \tag{1}$$

where $X_t = \left(1, q_{t-1}, \ldots, q_{t-p}\right)'$ and $\phi_i = \left(\phi_{io}, \phi_{i1}, \ldots, \phi_{ip}\right)'$ is a vector of parameters for $i = 1,2$, $\varepsilon_t \sim iid\left(0, \sigma^2\right)$ and $(\gamma \quad \mu) \in \{\Re^+ \times \Re\}$, where $\Re$ denotes the real line (-∞,∞) and $\Re^+$ the positive real line (0,∞). The transition function $G[\gamma; q_{t-d} - \mu]$ determines the degree of mean reversion and is itself governed by the transition parameter γ, and the "location" parameter μ. The transition parameter γ essentially determines the speed of mean reversion, while μ can be interpreted as the equilibrium level of $\{q_t\}$; the integer $d > 0$ denotes a delay parameter and reflects the possibility that market participants react to deviations from the equilibrium with a lag. Different choices of the transition function $G[\gamma; q_{t-d} - \mu]$ give rise to STAR models with different dynamic properties. A popular transition function for modelling real exchange rate dynamics is the following exponential function (see Michael, Nobay and Peel, 1997; Taylor, Peel and Sarno, 2001; Sarno 2003):

$$G[\gamma; q_{t-d} - \mu] = 1 - \exp\left[-\gamma (q_{t-d} - \mu)^2\right], \tag{2}$$

in which case the resulting model would be an exponential STAR or ESTAR model. The exponential transition function is bounded between zero and unity, G:$\mathfrak{R}\rightarrow$[0,1] and is symmetrically inverse-bell shaped around zero. The transition parameter γ determines the speed of transition between the two extreme regimes, G($\cdot$)=0 and G($\cdot$)=1, with lower absolute values of implying slower transition.

Alternatively, a logistic STAR (LSTAR) model may be considered, which is characterised by a logistic transition function of the type $G_L = [1-\exp\{-\gamma_L(q_{t-d}-\mu_L)\}]^{-1}$. The transition function of the LSTAR is a monotonically increasing function of q_{t-d} and yields asymmetric adjustment towards equilibrium.[5] In the present context, however, the properties of the exponential transition function are more attractive because they allow a smooth transition between regimes and symmetric adjustment of the real exchange rate for deviations above and below the equilibrium level, consistent with the predictions of the theoretical literature on real exchange rate determination under transaction costs. For this reason, we have a prior in favour of the ESTAR model, although, as discussed below, we select the most appropriate transition function to model the real exchange rate on the basis of a testing procedure designed to discriminate between the LSTAR and ESTAR models merely on a statistical ground.

Before estimating a nonlinear real exchange rate model, we need to test the null hypothesis that the real exchange rate is governed by a linear process against the alternative hypothesis of a nonlinearly mean-reverting STAR model. This linear-

[5] Asymmetric behaviour of the real exchange rate would imply that the speed of adjustment differs according to whether the real exchange rate is overvalued or undervalued, which does not seem plausible, particularly if one considers goods arbitrage as ultimately driving the impetus towards long-run equilibrium.

ity testing procedure is complicated, however, by the presence of unidentified nuisance parameters under the null hypothesis of linearity.[6] This problem can be overcome by using a low-order Taylor series expansion around the equilibrium, resulting in an auxiliary regression of the form:

$$q_t = \kappa_{00} + \kappa_0' X_t + \kappa_1' X_t q_{t-d} + \kappa_2' X_t q_{t-d}^2 + \kappa_3' X_t q_{t-d}^3 + e_t, \tag{3}$$

where $\kappa_i = (0, \kappa_{i1}, \ldots \kappa_{ip})'$ for i=0,1,2,3 (Saikkonen and Luukkonen, 1988; Luukkonen, Saikkonen and Teräsvirta, 1988). Teräsvirta (1994) derives LM-type tests of linearity against ESTAR and LSTAR models and also suggests a decision rule for choosing between ESTAR and LSTAR specifications. After specifying the autoregressive order p of q_t on the basis of the partial autocorrelation function (PACF) the linearity test against a STAR model consists of testing the null hypothesis that

$$H_{0L} : \kappa_1 = \kappa_2 = \kappa_3 = 0 \tag{4}$$

against the alternative that H_{0L} is not valid. In practice the ordinary F-test is recommended as an approximation to the LM-type test to obtain better size properties in small samples (Teräsvirta, 1994, 1998). In the present study this test statistic is termed F_L. Usually, the null hypothesis is tested for a set of plausible values of the delay parameter *d* and the value of *d* is chosen such that the marginal sig-

[6] Under $H_0 : \gamma = 0,$ the STAR model collapses to a linear AR(p) model and the parameter vector ϕ_2 and the location parameter μ can take any value without affecting the likelihood. Alternatively, the null hypothesis could be formulated as $H_0 : \phi_2 = 0,$ under which the STAR model reduces to the same linear AR(p) model as under $H_0 : \gamma = 0$; in the case of $H_0 : \phi_2 = 0,$ the transition parameter γ and the location parameter μ can take any value.

nificance level (p-value) of F_L is minimized. If linearity is rejected in favour of STAR-type nonlinearity, the next step involves discriminating between ESTAR and LSTAR formulations using a decision rule suggested by Teräsvirta (1994). The rule is based on the following sequence of nested tests within (3):

$$H_{03} : \kappa_3 = 0; \tag{5}$$

$$H_{02} : \kappa_2 = 0 | \kappa_3 = 0; \tag{6}$$

$$H_{01} : \kappa_1 = 0 | \kappa_3 = \kappa_2 = 0 . \tag{7}$$

The corresponding test statistics, denoted as F_3, F_2 and F_1 respectively, can be carried out using F-tests. The rule is as follows: after linearity is rejected using F_L, the three hypotheses (5) - (7) are tested. If F_2 yields the smallest p-value, an ESTAR model is chosen; otherwise an LSTAR specification is selected.[7]

At the estimation stage, the STAR model implied by the results in the identification stage is estimated consistently by nonlinear least squares (NLS) using a conventional optimisation routine. Granger and Teräsvirta (1993) and Teräsvirta (1994, 1998) suggest standardizing the exponent of the transition function by dividing it by the sample variance of the transition variable. Setting the transition parameter γ equal to unity is suggested as a reasonable starting value for the iterative NLS estimation.

Finally, once the model has been estimated, its properties can be evaluated and its adequacy assessed, which also involves a thorough examination of the residuals using a battery of misspecification tests, including tests for residual serial correla-

[7] Teräsvirta (1994) provides Monte Carlo evidence that this decision rule works well in selecting d as well as in discriminating between ESTAR and LSTAR formulations, unless the two models are close substitutes.

tion and for no-remaining nonlinearity (see Eitrheim and Teräsvirta, 1996; Teräsvirta, 1998).[8]

4 Data and Empirical Analysis

The data set comprises monthly observations on consumer prices for Bangladesh, the US, the Euro area, Japan and India, and nominal bilateral exchange rates for the Bangladesh taka, the euro, Japanese yen and Indian rupee vis-à-vis the US dollar. All time series except data on the Euro area cover the sample period from January 1994 to December 2002, and were taken from the International Monetary Fund's *International Financial Statistics* data base. Euro area data were drawn from the European Central Bank's monthly bulletin and span the sample period January 1997 to December 2002. Since the euro only took office in January 1999, we use the Ecu for the pre 1999 period. Monthly real exchange rate series for taka-dollar (TK/USD), taka-euro (TK/EUR), taka-yen (TK/JPY) and taka-rupee (TK/INR) were constructed from the above data according to the identity

$$q_t = s_t + p_t^{\mp} - p_t \tag{8}$$

and using the triangular arbitrage condition, where s_t denotes the logarithm of the nominal exchange rate (domestic price of foreign currency) observed at time t, p_t and $p_t^{\mp}$ are the logarithms of the domestic and foreign price levels respectively.

As a preliminary exercise we carry out tests for nonstationarity of the (log) real exchange rate series and their differences, based on Dickey-Fuller statistics. The

[8] Standard tests for residual autocorrelation, such as the standard Ljung-Box test, are not applicable in this context, as the asymptotic null distribution is unknown when the residuals come from a nonlinear model.

test statistics reported in Table 1 suggest that for all real log-level exchange rates the null hypothesis of a unit root cannot be rejected at conventional significance levels, while the possibility of multiple unit roots is excluded as all differenced real exchange rate series are found to be I(0). Thus on the basis of simple unit root tests one is tempted to conclude that PPP for Bangladesh and its major trading partners is violated. However, as the discussion in the previous section has shown, the presence of transaction costs may imply a nonlinear process for the real exchange rate, such that the framework surrounding conventional unit root tests, which are based on a linear autoregressive process, may not be appropriate in this context.[9]

Figure 1 displays a plot of the series for the deviations from the long-run equilibrium for each of the four real exchange rate series. Notably, for all four series we can observe remarkable short-run deviations from the mean. The examined series display strong persistence for smaller deviations from equilibrium, while becoming increasingly mean-reverting for larger deviations from equilibrium.

In order to execute the linearity test F_L we first chose the autoregressive order p on the basis of the PACF for each of the monthly real exchange rate series considered. For all real exchange rates the PACF indicated $p=1$. The linearity test, based on the auxiliary regression (3), was then performed for a set of values of the delay parameter $d \in \{1,2,\ldots,12\}$. As displayed in Table 2, for all real exchange rate series, the linearity test generally suggested rejections of the linearity hypothesis. In the case of the real TK/EUR and real TK/USD linearity is rejected

[9] Moreover, various studies have highlighted the low power of unit root tests in testing for mean reversion in the real exchange rate (see, for example, Lothian and Taylor, 1997; Sarno, Taylor and Chowdhury, 2004).

most strongly when $d=1$ and $d=2$, respectively, suggesting a rather fast response to shocks for these real exchange rates. Conversely, for the real TK/JPY and real TK/INR the linear null hypothesis was rejected with $d=4$ for the former and $d=5$ for the latter, indicating a somewhat slower adjustment to shocks. Having established STAR-type nonlinearity we next have to discriminate between an ESTAR or LSTAR model. Application of the Teräsvirta decision rule using the tests F_3, F_2 and F_1, reported in Table 3, further suggested an ESTAR model for all real exchange rates.

The estimation results, displayed in Table 4, indicate that the estimated ESTAR models have similar features. In particular, the speed of adjustment coefficients appear very similar, ranging between 0.98 and 1.07 and being strongly statistically significant in each case. The analysis of the stability properties of the estimated ESTAR models suggests that deviations from PPP are generally persistent or even explosive in the "lower" regime when the transition function $G(\cdot)=0$ and the ESTAR model reduces to a simple autoregressive process with the autoregressive coefficient greater than unity. Deviations from PPP are stationary in the "upper" regime when $G(\cdot)=1$ and the ESTAR model has a root inside the unit circle, implying global stationarity of the real exchange rate.

For each real exchange rate model the estimated transition function is plotted against the estimated transition variable, $q_{t-d} - \mu$, in Figure 2. Note that the limiting case of $G(\cdot)=1$ is never achieved. Moreover, the slope of the transition function indicates that the speed of transition across regimes is quite weak, providing a clear indication that slow, albeit significant, reversion towards long-run equilibrium characterizes most realized values of the real taka exchange rates over the investigated sample period. In general, a movement in the real exchange rate away from the equilibrium level μ as large as 10 percent would raise the transition function to about 0.10.

Goodness-of-fit statistics are also very satisfactory, with the coefficient of determination ranging between 0.85 and 0.89. The ratio of the residual variance of the estimated nonlinear model to the residual variance of the best fitting linear model *(V)* suggests, for all estimated ESTAR models, some reduction in the residual variance relative to the best fitting linear model. Note that since for the real TK/INR real exchange rate the test statistics for F_2 and F_1 were reasonably close, indicating either an LSTAR or an ESTAR, we also fit a LSTAR model, the estimation of which, however, produced insignificant estimates and a poor fit to the data.[10]

5 Conclusion

A number of studies have failed to provide evidence in support of long-run PPP. This may be due to a number of reasons, including the low power of conventional unit root tests and the assumption of linearity in real exchange rate dynamics. However, a recent literature has shown that the real exchange rate may be a nonlinearly mean-reverting process, such that adjustment towards the real exchange rate or PPP equilibrium level occurs at a speed that is a function of the size of the deviation from equilibrium itself.

[10] In particular, the estimated transition parameter was not found to be statistically significant, while the estimated residual variance of the LSTAR model exceeded the estimated residual variance of the best fitted linear autoregressive model (results are not reported but available upon request).

The empirical analysis carried out in the present study provides strong evidence that the real bilateral Bangladesh taka exchange rates between Bangladesh and its major trading partners are well characterized by processes which adjust nonlinearly towards their long-run equilibrium over the investigated sample period, and hence indicates the validity of long-run PPP consistent with the emerging theoretical literature on nonlinear real exchange rate adjustment in the presence of international arbitrage costs. The estimated models imply an equilibrium level of the real exchange rate in the neighbourhood of which the behaviour of the log-level of the real exchange rate is close to a random walk, while adjusting faster with the absolute size of the deviation from equilibrium.

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Table 1: Dickey-Fuller (DF) Test Statistics

	$q(t)$	$q^{\tau}(t)$	$\Delta q(t)$	$\Delta^2 q(t)$
TK/USD	0.2569	-2.4026	-8.1151	-9.0471
TK/EUR	-1.3525	-0.7416	-8.3299	-8.7104
TK/JPY	-1.8060	-1.9864	-10.1045	-12.2231
TK/INR	-1.8339	-3.0033	-9.7491	-7.8072

Notes: q denotes the log-level of the real exchange rate as defined in the text; the superscript τ indicates that a linear time trend is also included in the DF regression, and Δ denotes the first-difference operator. The 5% critical value for the DF test statistics is -2.88 (-3.43) if no (a) linear trend is allowed for.

Table 2: Linearity Tests

	TK/USD	TK/EUR	TK/JPY	TK/INR
d = 1	0.0325	0.0427	0.3932	0.4702
d = 2	0.0230	0.0988	0.1139	0.3711
d = 3	0.0576	0.8458	0.0164	0.3495
d = 4	0.0667	0.9943	0.0055	0.2751
d = 5	0.0874	0.6721	0.1717	0.0251
d = 6	0.0851	0.7404	0.8036	0.0710
d = 7	0.0755	0.7811	0.8778	0.6988
d = 8	0.1788	0.5424	0.9202	0.3348
d = 9	0.1472	0.3921	0.9872	0.1502
d = 10	0.0502	0.0664	0.7371	0.5176
d = 11	0.0591	0.6989	0.7851	0.6089
d = 12	0.0947	0.1395	0.5076	0.7076

Notes: F_L is the Lagrange-multiplier test statistic for linearity constructed as described in the text assuming that the order of the autoregression is $p=1$; d denotes the delay parameter. All marginal significance levels (p-values) are calculated using the appropriate F-distribution

Table 3: Tests F_3, F_2 and F_1

	TK/USD	TK/EUR	TK/JPY	TK/INR
d	2	1	4	5
F_3	0.8909	0.3783	0.1675	0.7883
F_2	0.0054	0.0076	0.0068	0.0179
F_1	0.1783	0.6119	0.0641	0.0543

Notes: The test statistics F_3, F_2 and F_1 are test statistics for discriminating between ESTAR and LSTAR formulations, constructed as described in the text; figures reported are marginal significance levels (p-values), calculated using the appropriate F-distribution; d denotes the delay parameter.

Table 4: Estimated STAR Models

Real TK/USD

$$\hat{q}_t = \underset{(24.62)}{1.08}\, q_{t-1} + \left(\underset{(2.29)}{0.06} - \underset{(-2.37)}{0.76}\, q_{t-1} \right)\left[1 - \exp\left\{ -\underset{(9.54)}{1.01}\left(q_{t-2}\right)^2 \right\}\right]$$

R^2=0.89; V=0.92; AR(1)=0.14; AR(3)=0.31; AR(6)=0.59; ARCH(1)=0.97; ARCH(3)=0.98; ARCH(6)=0.96; ET=0.85

Real TK/EUR

$$\hat{q}_t = \underset{(13.67)}{1.07}\, q_{t-1} - \underset{(-2.21)}{1.24}\, q_{t-1}\left[1 - \exp\left\{ -\underset{(3.83)}{0.98}\left(q_{t-1}\right)^2 \right\}\right]$$

R^2=0.88; V=0.93; AR(1)=0.04; AR(3)=0.14; AR(6)=0.16; ARCH(1)=0.80; ARCH(3)=0.77; ARCH(6)=0.79; ET=0.51

Real TK/JPY

$$\hat{q}_t = \underset{(25.01)}{1.07}\, q_{t-1} - \underset{(-2.68)}{0.31}\, q_{t-1}\left[1 - \exp\left\{ -\underset{(3.57)}{1.01}\left(q_{t-4}\right)^2 \right\}\right]$$

R^2=0.89; V=0.94; AR(1)=0.83; AR(3)=0.98; AR(6)=0.88; ARCH(1)=0.17; ARCH(3)=0.52; ARCH(6)=0.85; ET=0.09

Real TK/INR

$$\hat{q}_t = \underset{(20.60)}{1.03}\, q_{t-1} - \underset{(-2.34)}{1.49}\, q_{t-1}\left[1 - \exp\left\{ -\underset{(4.34)}{1.07}\left(q_{t-5} - \underset{(2.01)}{0.03} \right)^2 \right\}\right]$$

R^2=0.85; V=0.92; AR(1)=0.43; AR(3)=0.48; AR(6)=0.64; ARCH(1)=0.36; ARCH(3)=0.85; ARCH(6)=0.69; ET=0.40

Notes: A hat denotes the fitted value. Figures in parentheses below coefficient estimates denote t-ratios. R^2 denotes the coefficient of determination; *V* is the ratio of the residual variance from the estimated ESTAR model to the best fitting linear model for q_t; *AR(j)* is a Lagrange multiplier test statistic for up to jth-order serial correlation in the residuals and *ET* is a test statistic for no remaining nonlinearity as constructed in Eitrheim and Teräsvirta (1996); *ARCH(j)* is a Lagrange multiplier test statistic for up to jth-order autoregressive conditional heteroskedasticity in the residuals. For *ARCH(j)*, *AR(j)* and *ET* we only report *p*-values.

Figure 1: Real Exchange Rates (in de-mean)

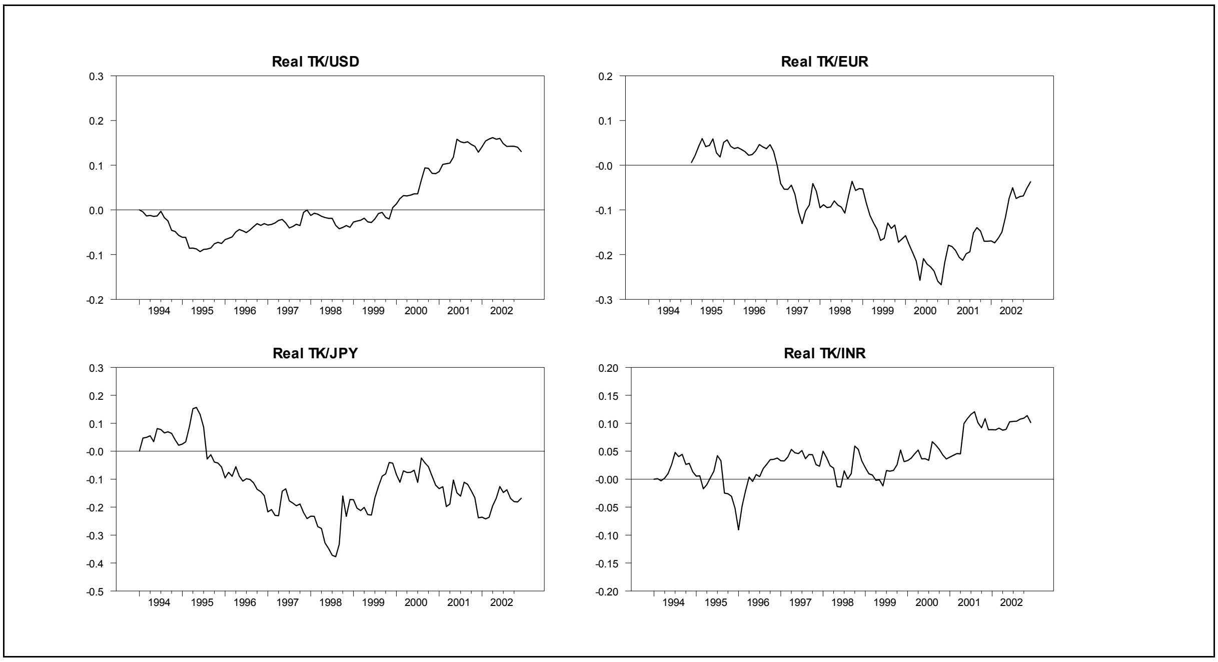

Figure 2: Estimated Transition Functions