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The Yen Real Exchange Rate May Be Stationary After All: Evidence from Nonlinear Unit-Root Tests

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Abstract

The empirical literature that tests for purchasing power parity (PPP) by focusing on the stationarity of real exchange rates has so far provided, at best, mixed results. The yen real exchange rate behavior, as compared to other major currencies, has most stubbornly challenged the PPP hypothesis and deepened this puzzle. This paper contributes to this discussion by providing new evidence on the stationarity of bilateral yen real exchange rates. We employ a non-linear version of the Augmented Dickey-Fuller test, based on an exponentially smooth-transition autoregressive model (ESTAR) that enhances the power of the tests against mean-reverting nonlinear alternative hypotheses. Our results suggest that the bilateral yen real exchange rates against the other G7 and Asian currencies were mean reverting during the post-Bretton Woods era. Thus, the real yen behavior may not be so different after all but simply perceived to be so due to the use of a restrictive alternative hypothesis in previous tests.

Key Words: PPP, yen, real exchange rates, nonlinear models, ESTAR models.

JEL Classification: C23, F31

1 Introduction

Testing for purchasing power parity (PPP) by focusing on real exchange rate stationarity has been a major focus of empirical international finance. However, the existing evidence cannot be considered definitive. The empirical consensus in the 1980s that viewed real exchange rates as random walks shuttered the Casselian view of PPP. Although this consensus has

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been challenged itself from more recent studies that employ relatively new methodologies,¹ the weak connection between exchange rates and national price levels is not resolved. On the contrary, it seems that the division of researchers between the “whittling down half lives” and the “whittling up half lives” camps, as termed by Taylor (2001), is as vivid as ever before. The yen real exchange rate emerges as one of the major currencies refuting the proposition that the real exchange rate is mean reverting in the long-run. Japan is often considered as the typical example of PPP failure. Thus, it is not surprising that the behavior of the yen real exchange rate has been the focus of many studies using different data sets and statistical methodologies (e.g., Lothian 1990, Ceglowski 1996, Cheung and Lai 2001).

The present paper contributes to this discussion by using new nonlinear stationarity tests to provide evidence about the time series properties of the bilateral yen real exchange rates. Recent highlights in the empirical literature investigate the presence of nonlinear adjustment dynamics in the real exchange rate process (e.g., Michael et. al. 1997, Taylor 2001), thus motivating the use of such tests. We employ a nonlinear version of the Augmented Dickey Fuller (ADF) tests statistic that is based on a exponential smooth-transition autoregressive (ESTAR) model as developed by Kapetanios, Shin and Snell (2001).² If the true model involves nonlinear dynamics that can be adequately approximated in some metric by a STAR (smooth-transition autoregressive)-type model then such nonlinear tests are better equipped to reject the null hypothesis of a unit root as compared to the standard ADF test or other variants of it. In other words, our tests have greater power than the standard (ADF) tests for a particular mean-reverting nonlinear alternative hypothesis. Such tests allow for the presence of a “corridor regime” within which mean-reversion does not occur, a feature consistent with the recent theoretical models where the presence of nonlinearities implies a “band of inaction” (for example, see Sercu et al. 1995).

The results of our analysis suggest that when allowing for a non-linear alternative the bilateral yen real exchange rate against the other G7 currencies appears stationary for our full sample (1960Q1-2000Q4). Moreover this result does not change when we restrict our attention to the post-Bretton Woods era (with the exception of the yen/DM real exchange rate). Those results become more interesting when compared to those from corresponding ADF tests that provide limited evidence of stationarity. More importantly, the yen/dollar real exchange rate appears stationary during the current float -in contrast to most of the existing literature that has failed to produce evi-

¹For surveys see Boucher Breuer (1994), Froot and Rogoff (1995), and Mark (2001).

²See also Blake and Kapetanios (2002) and Chortareas, Kapetanios and Shin (2002)

dence in favor of its stationarity. In addition, we consider the bilateral real exchange rate of the yen with the Asian and Pacific rim currencies as motivated by various considerations -including geographic proximity and trade intensities. The nonlinear tests are able to reject the nonstationarity null in at least twice the number of cases than the standard ADF tests during the current float period. The results of the analysis indicate that the inability to reject the unit-root null in the bilateral yen real exchange rates may not necessarily reflect some extraordinary feature of this currency that refutes PPP, but rather the low power of tests against some alternative hypotheses or the presence of a “corridor regime” in the real yen adjustment process.

The next section provides a literature review on evidence for the yen real exchange rate stationarity. Section 3 describes the testing methodology of Kapetanios, Shin and Snell (2001) and the data. Section 4 presents the results of our analysis. Section 5 discusses those results emphasizing how they differ from existing evidence and why. Finally, section 6 concludes.

2 Literature Review

A stylized fact of the recent float is the difficulty to distinguish real exchange rates from random walks and provide evidence for PPP. Various methodological approaches have been used including cointegration tests for exchange rates and prices, variance ratios tests, and unit root tests on real exchange rate series. Despite the theoretical appeal of PPP and the voluminous literature trying to uncover evidence of real exchange rate stationarity by considering longer time-series, low frequency data, and new statistical methodologies, studies on PPP remain inconclusive. Less ambiguity seems to exist, however, for the yen real exchange rate which is shown on balance to be nonstationary (e.g., Kim 1990, Cheung and Lai 1998, Koedijk et. al. 1998). Not only researchers have not been able to reject the null of a unit root with drift in the yen real exchange rate (Rogoff 1996) but also encountered difficulties in finding evidence for PPP when the yen was used as a base currency (e.g., Papell and Theodoridis, 1998).

A number of possible reasons can be put forward for the failure to find evidence for PPP. These include traditional forms of price stickiness (Dornbusch, 1976) as well as explanations based on trade costs (e.g., Dumas, 1992) and price discrimination (e.g., Chari, Kehoe, and McGratan, 2000). The empirical literature on the time-series properties of the Japanese real exchange rate has primarily focused on real factors, and in particular on cross-country differences in growth productivity. The trend-like appreciation of the Japanese

real exchange rate has been viewed as a typical case of the Balassa-Samuelson effect (e.g., Rogoff 1996, and Ito et. al. 1999).³

Departing from similar concerns Ceglowski (1996) attempts to capture the effects of relative productivity differentials on yen real exchange rates using the procedures suggested by Perron (1990) and Perron and Vogelsang (1992). Those tests and their modified versions enrich the ADF tests by allowing consideration of breaks in the mean as well as in the trend. The tests are applied to data spanning from the 1910s to 1991 and the non-stationarity result of the ADF tests is reversed in up to three out of five bilateral real exchange rates. All identified breaks in means, however, take place before the collapse of the Bretton Woods system, while the most recent break in trend is shown to take place in 1960. Since the current float is free from structural breaks the yen real exchange rate puzzle only deepens motivating further our analysis which is based on alternative methodology.

Sequential extensions of the ADF tests suggested by Banerjee et. al. (1992) have been considered as better equipped to capture the trends breaks and shifts in the data. Cheung and Lai's (1998) analysis that uses such tests covers four bilateral yen real exchange rates among others. When the mean shift is included in the alternative hypothesis they can reject the unit root null in the bilateral yen real exchange rates in only one out of four cases (yen/DM) in the post-Bretton Woods period when the mean shift is included in the alternative hypothesis and in none case when the trend-shift is included.

In general it has been difficult to reject the null hypothesis of a unit root in the real exchange rates when short data spans are considered -a difficulty more pronounced under the current float.⁴ Cheung and Lai (1998), however, show that relatively short sample sizes, such as the post-Bretton Woods era, do not necessarily render the rejection of non-stationarity impossible. They consider the bilateral real exchange rates of five industrialized countries including Japan using modified Dickey-Fuller tests suggested by Park and Fuller (1995) and Elliott et. al. (1996) to avoid the low-power problem. While both tests indicate that in the presence of a trend the yen/FF and yen/DM real exchange rates are stationary, they still cannot reject the null of a unit root in the yen/sterling and -more importantly- the yen/dollar real exchange rates.

³Obstfeld (1993) finds that the yen real exchange rate is characterized by a time trend and attributes it to productivity growth differentials between tradables and nontradables sectors. Marston (1987) also focuses on labor productivity differentials.

⁴Lothian and Taylor (1996) show that for sample sizes as that of the current float the power of standard unit root tests is extremely low resulting to an inability of rejecting the potentially false null.

The most popular recent method for circumventing the low-power problem of stationarity tests, however, is the use of panel methods. Applying such methods has typically allowed for the production of more evidence in favor of real exchange rate mean-reversion (e.g., Frankel and Rose 1996, McDonald 1996, Oh 1996, Wu 1996, Papell 1997).⁵ Panel unit root tests, however, are not free from potential drawbacks including their excessive sensitivity to the country groupings and the panel size (e.g., see Papell, 1997).

Cheung and Lai (2001) focus on the possibility of long-memory dynamics. They consider eight bilateral yen real exchange rates examining for evidence of fractional integration and find that the order of integration of all series considered is between zero and one. The use of fractionally integrated processes allows for long-cycles and longer-term memory and provides a flexible enough framework to simultaneously describe large swings and mean-reverting dynamics that may characterize real exchange rate behavior. Long-memory modelling, however, is essentially a purely statistical construct and the results should be interpreted with caution. The major difficulty is associated with giving a direct and meaningful economic interpretation to the non-integer order of integration.

One must also note the existence of a relatively small⁶ but growing literature on modeling exchange rates using nonlinear models such as the TAR and STAR models (e.g., Sarantis 2001, Baum et. al. 2001). These papers, however, do not consider the interplay between non-linearity and non-stationarity since they simply assume that the series (or their differences) are stationary.

Despite the widespread consensus on real yen yen's nonstationarity there exist studies providing evidence in favor of the opposite direction. Phylaktis and Kassimatis (1994) use black market exchange rates and find evidence for PPP but the results are stronger when Japan is excluded from their sample. Lothian (1990) finds evidence that the yen real exchange rates are mean reverting in very long samples. Interestingly, Lothian (1990) indicates that this adjustment process may not be continuous. The possibility of such a discontinuous adjustment process further motivates the use of a threshold model which can be approximated by a STAR-type model as the one that we consider.

⁵Exceptions include O'Connell (1997) and Chortareas and Driver (2001).

⁶For example, Sarantis (1999) mentions that "...there has been no attempt to investigate and model nonlinearities in real exchange rates and particular in effective exchange rates" (p. 28).

3 Methodology and Data

3.1 Testing for a Unit Root Against the Nonlinear *STAR* Model

Here we give an account of the testing methodology of Kapetanios, Shin and Snell (2001). Consider a univariate smooth transition autoregressive of order 1 (*STAR*(1)) model,

$$y_t = \beta y_{t-1} + \beta^* y_{t-1} \Phi(\theta; y_{t-d}) + \varepsilon_t, \quad t = 1, \dots, T; \quad d \geq 1, \quad (1)$$

where $\varepsilon_t \sim iid(0, \sigma^2)$, and β, β^* are unknown parameters. It is assumed that y_t is a mean zero stochastic process. (For the case with non-zero mean and/or with a linear time trend see the discussion below.) $\Phi(\theta; y_{t-d})$ denotes the transition function. We also assume that $\theta \geq 0$, and the delay parameter $d \geq 1$ is given. If θ is positive, then it effectively determines the speed of mean reversion. The representation (1) makes economic sense in that many economic models predict that the underlying system tends to display a dampened behavior towards an attractor when it is (sufficiently far) away from it, but that it shows some instability within the locality of that attractor. A classic example is the floor and ceiling model of output analyzed by Hicks (1950).

Following the literature on the *STAR* models the popular exponential transition function is considered.

$$\Phi_E(\theta y_{t-d}) = 1 - \exp(-\theta y_{t-d}^2). \quad (2)$$

The exponential transition function is bounded between zero and 1, *i.e.* $\Phi: \mathbb{R} \rightarrow [0, 1]$, has the properties

$$\Phi_E(0) = 0; \quad \lim_{x \rightarrow \pm\infty} \Phi_E(x) = 1,$$

and is symmetrically U-shaped around zero.

Using (2) in (1) gives an exponential *STAR* (*ESTAR*) model

$$y_t = \beta y_{t-1} + \beta^* y_{t-1} [1 - \exp(-\theta y_{t-d}^2)] + \varepsilon_t. \quad (3)$$

The null hypothesis of a unit root is considered, which in terms of the above model implies that $\beta = 1$ and $\theta = 0$ (and thus $\Phi_E(\cdot) = 0$). Under the null, then (3) becomes the nonstationary linear *AR*(1) model:

$$y_t = \beta y_{t-1} + \varepsilon_t. \quad (4)$$

Under the alternative of stationarity, θ is strictly positive and (3) becomes

$$y_t = \{\beta + \beta^* \Phi_E(\theta y_{t-d})\} y_{t-1} + \varepsilon_t, \quad 0 < \Phi_E(\theta y_{t-d}) < 1. \quad (5)$$

This clearly shows that y_t would locally follow a unit root in the region of $y_{t-d} = 0$ in which case $\beta + \beta^* \Phi_E(\theta y_{t-d})$ would be unity. Large values of y_{t-d} on the other hand would result in an approximately linear $AR(1)$ process with the stable root $\beta + \beta^*$ provided that $-2 < \beta^* < 0$. It is assumed that the latter is the case.

Explicitly then the null hypothesis is

$$H_0 : \theta = 0, \quad (6)$$

against the alternative⁷

$$H_1 : \theta > 0. \quad (7)$$

Obviously, testing the null hypothesis (6) directly is not feasible, since β^* is not identified under the null (see Davies 1977, 1987).

Considering that the standard linear ADF test is not expected to be very powerful when the true process is stationary but nonlinear, the direct testing framework can be developed. This involves estimating the following auxiliary regression

$$\Delta y_t = \delta y_{t-1}^3 + error, \quad (8)$$

and using the following t-test statistic for (6)

$$NLADF = \frac{\hat{\delta}}{s.e.(\hat{\delta})}, \quad (9)$$

where $\hat{\delta}$ is the OLS estimate obtained from the auxiliary regression and $s.e.(\hat{\delta})$ is the standard error of $\hat{\delta}$. The test is motivated by the fact that the auxiliary regression is testing the significance of the score vector from the quasi-likelihood function of the *ESTAR* model, evaluated at $\theta = 0$. The *LM* test of (6) against (7) also tests the significance of this term and is thus intimately related to the t-test that we consider. For a more thorough discussion of the test see Kapetanios, Shin and Snell (2001). Unlike the case of testing the linearity against the nonlinearity for the stationary process the *NLADF* test does not have an asymptotic standard normal distribution. Kapetanios, Shin, and Snell (2001) provide details on the asymptotic properties of the test.

⁷Under the alternative (7) y_t is stable as $1 + \beta_* \Phi_E(\theta y_{t-d}) < 1$ for all y_{t-d} .

The *NLADF* test for the above case needs modifications to accommodate the process with a non-zero mean and a linear trend. In the case where the data have non-zero mean such that $x_t = \mu + y_t$, the demeaned data $x_t - \bar{x}$ are used in (1), where \bar{x} is the sample mean. Similarly, for the case with non-zero mean and non-zero linear trend such that $x_t = \mu + \delta t + y_t$ the de-meaned and de-trended data $x_t - \hat{\mu} - \hat{\delta}t$ are used in (1), where $\hat{\mu}$ and $\hat{\delta}$ are the OLS estimators of μ and δ .

Asymptotic critical values of the *NLADF* statistics for the above three cases, denoted *NLADF1*, *NLADF2*, *NLADF3*, respectively, have been tabulated via stochastic simulations with $T = 1,000$ and $100,000$ replications, and presented in Table 1 below.

Table 1:

| Asymptotic Critical Values | | | |
|----------------------------|---------------|---------------|---------------|
| sig. level | <i>NLADF1</i> | <i>NLADF2</i> | <i>NLADF3</i> |
| 99% | -2.82 | -3.48 | -3.93 |
| 95% | -2.22 | -2.93 | -3.40 |
| 90% | -1.92 | -2.66 | -3.13 |

A relevant issue that emerges is the possibility of serial correlation in the error term. The presence of serial correlation may be dealt with by an augmentation similar to that undertaken for the Dickey-Fuller tests. In particular, lagged differences of the dependent variable may be included in the regression and the asymptotic distribution of the t-test does not change.⁸ Furthermore, many results that are valid for the linear case such as the data-dependent selection of the lag order of the augmentation extend to the nonlinear test. In this paper we use a sequential testing procedure to determine the optimal lag order starting from a maximum lag order of 4.

3.2 Data

We construct bilateral yen real exchange rate against the *i*-th currency at time *t* ($q_{i,t}$) as $q_{i,t} = s_{i,t} + p_{J,t} - p_{i,t}^*$, where $s_{i,t}$ is the corresponding nominal exchange rate (*i*-th currency per yen), $p_{J,t}$ the price level in Japan, and $p_{i,t}^*$ the price level of the *i*-th country. Thus, a rise in $q_{i,t}$ implies a real yen appreciation against the *i*-th currency. The price levels are consumer price

⁸For a proof and detailed discussion see Kapetanios, Shin, and Snell (2001).

indices and all variable are in logs. All data are from the International Monetary Fund's *International Financial Statistics* in CD-ROM. The data are not seasonally adjusted. The first set of bilateral real exchange rates consists of those of the other G7 countries, i.e., the US, the UK, Germany, Italy, France, and Canada. The second set of bilateral real exchange rates are those of Australia, Korea Singapore, Malaysia, Philippines, New Zealand, Hong Kong, Indonesia, Thailand, and Sri Lanka.

The Pacific Basin countries of our sample represent approximately one fourth (24%) of Japan's total trade, the US represents another one fourth, and the remaining five G7 the 12%.⁹ Furthermore, the experience of Japan may be of relevance for the Pacific Basin countries. Ito et. al. (1999) suggest that the pattern of industrial development in Asia may be similar to what has been called a "flying geese pattern" with Japan as a leader followed by Hong Kong and Singapore, which are followed by Korea, followed in turn by Taiwan and Thailand, then by Indonesia and so on.

All data are quarterly, spanning from 1960Q1 to 2000Q4 and the bilateral nominal exchange rates against the currencies other than the US dollar are cross-rates computed using the US dollar rates. Previous research on the stationarity of real exchange rates in general, and of the yen real exchange rate in particular has used monthly, quarterly, and annual data. Typically, lower frequency data were used when longer time spans were considered. Our sample is shorter as compared to studies whose data span measures in centuries or so (e.g., Lothian 1990, Ceglowski 1996) but almost one decade longer than the existing studies focusing on the post Bretton-Woods real yen behavior. Those analyses use annual and monthly data respectively. Using quarterly data, however, seems a good compromise, since it allows comparability of our results with most of the existing literature.¹⁰

⁹Data are from OECD's Statistical Comedium.

¹⁰Recently, Taylor (2001) points to the implications of considering data whose frequency does not match that of the underlying arbitrage process. Clearly, considering low frequency data is not the best strategy when the hypothesized adjustment process is a high frequency one. Temporal aggregation problems may result to biased coefficients and low power in standard unit root tests. While such considerations valid we proceed with the quartely data to make our results comparable with those of the existing literature. We should note, however, that if the true model is nonlinear indeed high frequency data will make it easier to pick up the nonlinear adjustment. On the other hand, if the adjustment process is a long-memory one then it will be picked up easier by longer data spans of low frequency.

4 Results

Figures 1 and 2 show the logarithms of the bilateral yen real exchange rates against the other G7 countries, and the Asia and Pacific countries in our sample respectively. In all cases the presence of a trend in the bilateral yen real exchange rate is apparent. Hong Kong seems to be an exception, but this is possibly because of the limited data span. Thus, we include a trend in the unit-root tests we perform subsequently.

Table 2 provides the results from applying the tests described in section 3.1 to the real exchange rates of the G7 countries, covering our full sample. These tests as well as the tests that we subsequently perform employ the version of the test that uses detrended data. The nonlinear version of the ADF (NLADF) tests show that all six bilateral real exchange rates are stationary either at the 1% or the 5% level of significance. Furthermore, the choice of lag augmentation does not affect the results. That is, all six series appear stationary when both a four-lag scheme is imposed and when the choice of the optimal number of lags occurs through an optimization routine that consecutively tests down from a higher to a lower number of lags. On the other hand, the standard ADF tests indicate stationarity for only two real exchange rates, namely the sterling and the Canadian dollar.

When we restrict our focus on the post-Bretton Woods era in Table 4 the results of the NLADF tests (of Table 2) remain robust. The only exception seems to be the yen/DM real exchange rate that displays a unit root under both the linear and nonlinear alternative hypotheses. The standard ADF tests provide evidence of stationarity for up to three bilateral real exchange rates, namely Italy, the UK, and Canada. Only the evidence for the yen/Canadian dollar real exchange rate, however, appears to be robust to the choice of the lag augmentation. On the contrary, the results of the NLADF tests that demonstrate stationarity in five out of the six real exchange rates are invariant to the lag length/order selection. It is interesting that our analysis allows rejection of the unit-root null for the yen/dollar real exchange rate despite the notable variability of the yen against the dollar during the sample which includes the turbulent dollar's 1980's.

Turning to the Asian and Pacific economies now we obtain evidence of stationarity for six out of the nine countries when we consider the full sample period. Those results are invariant to the lag length selection. The countries for which we fail to reject the null hypothesis of a unit root are Korea, Malaysia and Sri Lanka. The results of the ADF tests depend on the method of lag augmentation. When the endogenous lag selection method is used then the ADF test rejects the null hypothesis of a unit root in only three cases out of nine. The countries for which the null hypothesis is rejected are

Australia, Korea and the Philippines. Interestingly, the null hypothesis is rejected for Korea using the ADF test but not using the NLADF leading us to believe that there might be a degree of complementarity between the two tests. In other words the two tests seem to be powerful against different directions of deviation from the null hypothesis, as expected. Turning to the floating exchange rate period, the NLADF test rejects the null hypothesis in five out of nine cases (Table 7). These countries are Australia, Korea, Hong Kong, Indonesia and Philippines. This result is quite robust to the lag-length selection method employed. For example, when four lags are used, the result changes in only one country (Korea). The ADF tests, on the other hand, reject the nonstationarity null in up to two out of nine cases (and only when the automatic lag selection method is used).

5 Discussion

Our results indicate that when the alternative hypothesis in the unit root tests can accommodate the presence of nonlinear adjustment the yen bilateral real exchange rates appear to be mean-reverting. Those findings are consistent with recent theoretical analyses that point to various rigidities including transportation costs and monopolistic price-setting. For example, Sercu et. al. (1995) modify the Lucas (1982) model to allow for the presence of transactions costs, and show that the international imbalances between marginal utilities of consumption (the ratio of which corresponds to the real exchange rate) do not adjust if they are sufficiently small relative to the transaction costs. Dumas (1992) provides a model where spatial market separation and transaction costs give rise to a band of inaction. Deviations from PPP are persistent but mean-reversion through a nonlinear adjustment process finally occurs. Uppal (1993) provides a similar nonlinear channel of real exchange rate adjustment.

Our findings are also consistent with other recent empirical approaches that consider nonlinear adjustment dynamics. Baum et. al. (2001) find evidence for weak PPP in the CPI-based yen/dollar real exchange rate. This study, as well as those by Balke and Fomby (1997) and Michael et. al. (1997), focuses on cointegration between nominal exchange rates and price indices using threshold cointegration techniques. All those analyses, however, assume that the real exchange rate processes themselves are stationary and ergodic and focus on modeling the nonlinear adjustment. Our analysis differs from the above in that we test for stationarity by examining for the presence of unit root in the real exchange rate process itself. Identifying the

exact source of nonlinearities in the yen real exchange rates is beyond the scope of this paper.

A typical feature in the relevant literature is the difficulty to produce evidence in support for PPP using the yen as the numeraire currency. For example, Papell and Theodoridis (1998) conduct both panel and univariate unit-root tests using the US dollar and the DM as base currencies and find stronger evidence for PPP in panels when the DM is used as the numeraire currency. Using the Yen as a base currency only increases the difficulty to find evidence of real exchange rate stationarity. Koedijk et. al. (1998) use an alternative panel methodology that allows consideration of individual country effects when testing for the relationship between exchange rates and prices. Similarly to the previous authors, they obtain stronger evidence for PPP when the DM rather than the US dollar is used as a numeraire. Using the Japanese yen, however, provides again the weakest evidence for PPP. Interestingly, the use of the the yen as the base currency throughout this study does not hinder the ability to produce evidence for stationarity for both the G7 bilateral real exchange rates and those of other Asian and Pacific countries.

The nonlinear unit-root tests results provide always more evidence in support of the bilateral real yen stationarity when considering the full sample as compared to the current float only. This is consistent with Rogoff's (1996) observation that the most convincing evidence on PPP comes from data sets employing at least some fixed exchange rates data. Mussa (1986) also suggests that given the relatively similar price level paths under fixed and floating exchange rates most of the variability in real exchange rates can be accounted for by changes in the nominal exchange rates. Whether the presence of this regularity in our findings is a feature of the longer data span or due to the fixity of the exchange rates, however, is a matter of further research.

6 Conclusion

If there is a G7 real exchange rate that makes the rejection of the unit-root hypothesis difficult, this is the yen real exchange rate. Existing research has responded either by accepting this results as a fact (typically attributing yen's behavior to Balassa-Samuelson effects) or by considering longer data spans and -more recently- by utilizing different statistical techniques that may account for the series behavior more appropriately. Focusing on possible nonlinearities has been one direction within the last strategy.

We contribute to this discussion by considering the stationarity of bilat-

eral yen real exchange rates using new unit root tests that incorporate a non-linear alternative hypothesis. That is, in contrast to other recent studies that assume real exchange rate stationarity and try to explain the nonlinear adjustment, we check for the presence of nonlinear dynamics in the process itself. Such considerations are consistent with recent theoretical developments that emphasize the role of transactions costs and monopolistic pricing as possible explanations for the presence of “inaction bands” and nonlinear mean-reversion of real exchange rates.

Our results show that the null of nonstationarity is rejected when the alternative hypothesis in ADF-type unit root tests is modified to allow for the presence of nonlinearities. The nonlinear ADF tests provide stationarity evidence for the Japanese real exchange rate in at least twice the cases as compared to the standard ADF tests during the post-Bretton Woods era. When considering our full sample (starting from 1960) all six bilateral yen real exchange rates against the other G7 countries emerge as mean-reverting. Similar results in favor of PPP emerge when we consider the bilateral yen real exchange rates with other Asian and Pacific countries. Those results become more intriguing, when one recalls that earlier research encountered extra difficulties in uncovering evidence for PPP when the yen was used as the base currency.

Providing an explicit explanation for the underlying sources of the nonlinear adjustment process in the yen real exchange rates is beyond the scope of this paper and constitutes the subject of further research. The present paper makes the point that the yen real exchange rate behavior may not be so particular after all. That is, it may be simply the use of tests that incorporate an inappropriate alternative hypothesis that is responsible for what has been considered “exceptional” real exchange rate behavior.

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Table 2: 1960Q1 2000Q4

| | DF | ADF(4) | ADF(A) | NLDF | NLADF(4) | NLADF(A) |
|---------|--------|---------|----------|--------|----------|----------|
| US | -1.666 | -2.768 | -2.675 | -2.618 | -3.928* | -3.850* |
| Germany | -2.669 | -2.960 | -2.904 | -3.222 | -3.728* | -3.504* |
| France | -2.412 | -3.115 | -3.072 | -2.939 | -3.947** | -3.828* |
| Italy | -2.015 | -2.853 | -3.011 | -3.230 | -4.606** | -4.848** |
| UK | -2.112 | -3.050 | -4.079** | -2.811 | -3.986** | -4.603** |
| Canada | -2.309 | -3.748* | -4.082** | -2.537 | -3.848* | -3.838* |

Table 3: 1974Q1 2000Q4

| | DF | ADF(4) | ADF(A) | NLDF | NLADF(4) | NLADF(A) |
|---------|--------|---------|---------|--------|----------|----------|
| US | -1.785 | -2.776 | -2.692 | -2.825 | -4.227** | -4.181** |
| Germany | -2.234 | -2.367 | -2.460 | -2.639 | -3.084 | -2.894 |
| France | -2.192 | -2.687 | -2.784 | -2.770 | -4.019** | -4.495** |
| Italy | -2.220 | -3.082 | -3.664* | -3.242 | -4.758** | -5.305** |
| UK | -1.982 | -2.812 | -3.937* | -2.571 | -3.798* | -4.770** |
| Canada | -2.019 | -3.436* | -3.801* | -2.590 | -4.042** | -4.237** |

Table 4: 1960Q1 2000Q4

| | DF | ADF(4) | ADF(A) | NLDF | NLADF(4) | NLADF(A) |
|-------------|---------|----------|----------|----------|----------|----------|
| Australia | -2.634 | -3.587* | -4.652** | -3.793* | -5.106** | -5.812** |
| Korea | -2.931 | -3.880* | -3.651* | -2.621 | -3.331 | -3.106 |
| Hong Kong | -1.472 | -2.030 | -1.977 | -2.853 | -4.228** | -4.199** |
| Singapore | -1.944 | -3.435* | -3.329 | -2.105 | -3.612* | -3.592* |
| Malaysia | -2.373 | -3.565* | -3.325 | -2.359 | -3.328 | -3.190 |
| Indonesia | -2.338 | -2.756 | -2.521 | -2.906 | -4.624** | -4.283** |
| Thailand | -2.532 | -3.281 | -3.243 | -2.975 | -3.777* | -3.721* |
| Philippines | -3.566* | -4.216** | -3.761* | -3.938** | -5.036** | -6.472** |
| Sri Lanka | -0.782 | -1.282 | -1.578 | -1.128 | -1.761 | -2.229 |

Table 5: 1974Q1 2000Q4

| | DF | ADF(4) | ADF(A) | NLDF | NLADF(4) | NLADF(A) |
|-------------|--------|--------|---------|----------|----------|----------|
| Australia | -2.136 | -2.944 | -3.541* | -3.075 | -4.186** | -4.644** |
| Korea | -2.879 | -3.215 | -3.441* | -2.678 | -3.132 | -3.448* |
| Hong Kong | -1.472 | -2.030 | -1.977 | -2.853 | -4.228** | -4.199** |
| Singapore | -1.707 | -3.040 | -3.145 | -1.827 | -3.095 | -3.274 |
| Malaysia | -1.894 | -2.797 | -2.643 | -2.087 | -2.961 | -2.878 |
| Indonesia | -2.934 | -3.226 | -2.902 | -3.004 | -4.569** | -4.186** |
| Thailand | -2.305 | -2.830 | -2.848 | -2.663 | -3.302 | -3.328 |
| Philippines | -3.196 | -3.257 | -3.188 | -4.767** | -5.407** | -6.721** |
| Sri Lanka | -1.902 | -2.631 | -3.254 | -1.586 | -2.232 | -2.868 |

Figure 1: Bilateral yen real exchange rates against the other G7 currencies

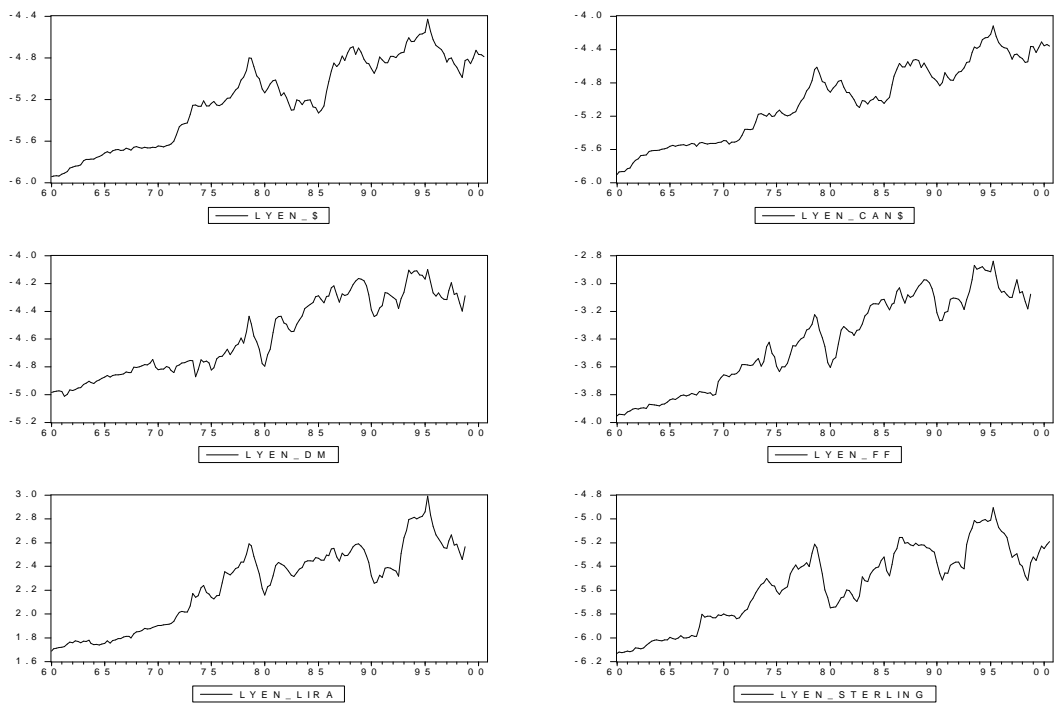
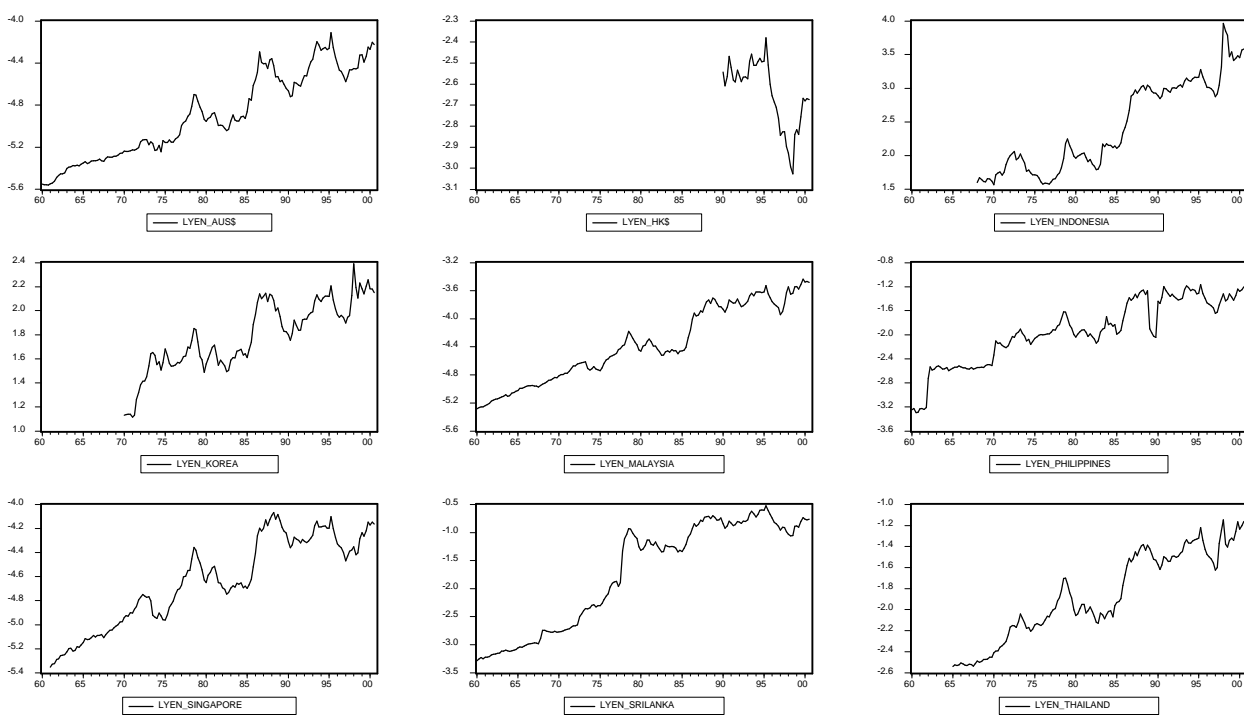


Figure 2: Bilateral yen real exchange rates



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