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Abstract

This paper examines the statistical properties of the real exchange rates of G-5 countries for the post Bretton-Woods period, and draw implications on the purchasing power parity (PPP) hypothesis. In contrast to most previous studies that consider only unit root and stationary processes to describe the real exchange rate, this paper also considers two in-between processes, the locally persistent process and the fractionally integrated process, to complement past studies. Seeking to be consistent with the ample evidence of near unit root in the real exchange rate, this paper suggests that a local persistence process describes the real exchange rate movements very well. This finding implies that: 1) the real exchange movement is more persistent than the stationary case but less persistent than the unit root case; 2) the real exchange rate is non-stationary but the PPP reversion occurs and the PPP holds in the long run; 3) the real exchange rate does not exhibit the secular dependence of the fractional integration; 4) the real exchange rate evolves over time in a way that there is persistence over a range of time, but the effect of shocks will eventually disappear over time horizon longer than order $O(n^d)$, that is, at a finite time horizon; 5) shocks dissipation is faster than predicted by the fractional integration, and the total sum of the effects of a unit innovation is finite, implying that a full PPP reversion occurs at finite horizons. These results may explain why past empirical studies could not provide a clear-cut conclusion on the real exchange rate processes and the PPP hypothesis.

1. Introduction

Many past studies have examined statistical properties of the real exchange rate to draw the implications on the empirical relevance of the purchasing power parity (PPP) in the long run. Earlier studies found less favorable evidence of PPP in industrial countries for the post-Bretton Woods period, by reporting that the null hypothesis of unit root against stationarity of the real exchange rates was not rejected. These tests were mostly implemented through the augmented Dickey-Fuller (ADF) and/or Phillips-Perron (PP) statistics.

As suggested by Frankel (1986), Froot and Rogoff (1995), and Rogoff (1996), a popular explanation for the failure to reject the unit-root null has been the low power of the unit-root tests, which is aggravated by the high persistency (a root near unity) of the real exchange rate and the short time-span of the post-Bretton Woods sample. Many past researchers have tried to resolve this low power problem by extending data points. First, some researchers such as Frankel (1986) and Lothian and Taylor (1996), extended the sample period, which is comprised of both flexible and fixed exchange rate regime periods. Second, others studies such as Frankel and Rose (1996), Lothian (1997), Choi (2000), and Papell (1997), employed the panel data method to gain power by exploiting both cross-sectional and time-series variation.

While these studies employ the statistical procedure that the real exchange rate is assumed to be either a stationary or a unit root process, this paper explores the possibility that the real exchange rate process lies between stationary and unit root. Seeking to be consistent with the ample evidence of near unit root in the real exchange rate and to complement past studies[1], we apply recent econometric developments on modeling and testing near unit root processes to the real exchange rate movements. In particular, we examine whether the locally persistent process, which was introduced by Phillips et al (2001), can describe the real exchange rate movements, and draw the implications for the PPP. Local persistence in the real exchange rate implies that the real exchange rate is more persistent than the stationary case but less persistent than the unit root case. It also implies that the real exchange rate is non-stationary but a mean reversion occurs in the long run and the PPP holds (although the full reversion would take more time than the stationary case).

To examine the issue, we adopt the modified Dickey-Fuller (DF-GLS) test suggested by Elliot et al. (1996) and Lima and Xiao (2002) test. The former tests the null hypothesis of the unit root against the point near unity alternative while the latter tests the null hypothesis of the stationarity against the local persistence. We apply these tests to the US-based and Japanese-based bilateral real exchange rates of G-5 countries for the Post Bretton-Woods period, and found that the local persistence approximates the real exchange rate movements better than the stationary and the unit root process.

Throughout the paper, we compare the locally persistent process to the fractionally

[1] Evidence of a root near unity in the real exchange rates has been frequently reported in the literature. In fact, Darby (1980), MacDonald (1985), Enders (1988), and Mark (1990) found that the autoregressive coefficient (which is the coefficient of mean reversion) of various real exchange rates is around 0.97, and this implies that real exchange rates have a root near unity but not necessarily an exact unit root.

integrated process (Granger and Joyeux, 1980; Hosking, 1981) since some past studies such as Cheung and Lai (1993, 2001) and Diebold et al (1991) suggest that the real exchange rate follows the fractionally integrated process, which can be also regarded as the process that lies between stationary and unit root. An interesting feature of the fractionally integrated process is its pattern of temporal dependence. Fractional models have been largely used in physical sciences to model time series with secular cycles. The presence of secular cycles implies the existence of non-zero autocorrelation between two events observed many years apart, or a secular dependence between time series observations. Compared to local persistence, we expect longer temporal dependence in the fractional integration case. On the other hand, the fractionally integrated process also displays a mean-reverting behavior. Therefore, the fractional integration in the real exchange rate would imply that PPP holds in the long run as in the local persistence case. However, due to the hyperbolic decaying of the impulse response coefficient, a full PPP reversion may not occur at a finite horizon, or within a reasonable long period of time, if the real exchange rate is fractionally integrated. Thus, the PPP reversion would take far longer than the case of local persistence. To support the local persistence in the real exchange rate against the fractional integration, we show that the temporal dependence pattern found in the real exchange rate data is more consistent with the local persistence than the fractional integration.

This paper is organized as follows. Section 2 presents the local persistence model, developed by Phillips et al (2001). Section 3 compares the locally persistent process with the fractionally integrated process, based on their temporal dependence and the impulse responses. In doing so, we also draw the implications on the PPP hypothesis. Section 4 presents various statistical procedure to test whether a time series is a stationary, a local persistence, or a unit root process. Section 4 presents the empirical results. Finally, Section 5 concludes.

A word on notation. We will use the symbols " $\Rightarrow$ ", " $\rightarrow$ " and " $:=$ " to signify weak convergence, convergence in probability and equality in distribution, respectively. Following the standard stochastic order of magnitude notation, we write $X_n = O_p(1)$ and $X_n = o_p(1)$ to signify that the sequence of random variable X_n is bounded and converges to zero, respectively, as the sample size, n, goes to infinite.

2. A Model with Local Persistence

2.1. Locally Persistent Process Without Drift

The time series that we consider is the following process

$$y_i = ay_{i-1} + u_i,\ a = e^{\frac{c}{n^d}} \sim 1 + \frac{c}{n^d},\ i = 1, ..., n \tag{2.1}$$

where the coefficient "a" in the autoregression is near unity and measures the persistency in time series y_i. We allow the errors u_i to be a linear process, which includes many general covariance stationary time series that satisfies an invariance principle.

It can be shown that by a re-parameterization ($m = n^d$) we may re-write the above time series in the format of a block local to unity model which was first introduced by Phillips et al (2001). For more discussions on regularity assumptions and asymptotic properties of the above time series, see Phillips et al (2001).

The process provides a new form of persistent behavior. The above device provides a statistical model for what may be described as "locally persistent behavior" for macroeconomic time series. Many macroeconomic time series are now well known to display a form of persistence whereby economic shocks have long-run effects. However, it is possible that shocks may affect an economy for a long period of time but not forever. In other words, the effects of a shock may be highly persistent over a certain range (the range of persistent behavior), but then may begin to disappear outside this range. In the above model, the autoregressive root of time series y_i is close to unity, and thus persistency can be found in y_i. On the other hand, the series evolves over time in such a way that there is persistency over a range of time (of order $O(n^d)$, compared to the full sample range n), but the effect of shocks will eventually disappear over time horizon longer than order $O(n^d)$. The region of persistent behavior may constitute a 'little infinity' relative to the full sample. Since there is persistent memory within a time horizon of order $O(n^d)$, but there is only short memory over longer periods, we call this type of memory "local persistence" or "regional persistence". For this reason, we call the above process a locally persistent process with persistent parameter d.

We are especially interested in the case that $0 < d < 1$. In this case, the process has autoregressive coefficient near unity $\alpha = e^{c/n^d}, c < 0$ but it is not the conventional stationary or unit root process. It is a locally persistent process. In particular, $y_{[nr]} = O(n^{d/2})$, implying that the process will ultimately diverge at rate $n^{d/2}$ as $n \to \infty$.

However, the model in equation (2.1) does include the traditional stationary process and the unit-root type persistency as two special extreme cases. In particular,

(i) When $d = 1$, it reduces to the traditional near-integrated processes and has the conventional unit-root type persistency. In particular, shocks has permanent effects, with $y_{[nr]}$ diverging at rate $n^{1/2}$, which is the same rate of divergence presented by a unit root process. See, inter alia, Phillips (1988), for properties of this kind of process

(ii) When $d = 0$, the process becomes standard stationary process.

Locally persistent processes are not covariance-stationary and, as we will see soon, they may be used to model the dynamics of economic time series that display persistence as well as transitory shocks. Overall, we can view the local persistent process, $0 < d < 1$, as a process that lies between a covariance-stationary and unit root.

2.2. Locally Persistent Process With A Deterministic Trend

The locally persistent process can be extended to include a deterministic time trend. Such an extension is important because many economic time series display tendency of growth. We may consider a locally persistent process y_i^τ with trend component

$$y_i^\tau = \tau_i + y_i;\ y_i = ay_{i-1} + u_i,\ a \sim 1 + \frac{c}{n^d},\ i = 1, ..., n. \tag{2.2}$$

where τ_i is a deterministic trend, the leading case being a linear time trend where $\tau_i = \varphi_o + \varphi t = \varphi_y' \Upsilon_i$. The stochastic part represented by equation (2.2) corresponds to (2.1), a locally persistent process without trend.

Note that the trend coefficients are unknown and thus, in practice, appropriate detrending is needed. We may estimate y_i from the residuals of the following detrending regression

$$y_i^{\tau} = \widehat{\varphi}_y' \Upsilon_i + \widehat{y}_i \tag{2.3}$$

where $\widehat{\varphi}_y$ is the least squares estimator of φ_y. The detrended time series $\widehat{y}_i$ has properties similar to the process with no drift y_i (see Phillips et al, 2001).

3. Local Persistence and Long Memory in Real Exchange Rate and PPP Hypothesis

The real exchange rate behavior has been the object of intense study in past works. By examining the statistical properties of the real exchange rate, past studies have tried to draw some implications on the PPP hypothesis. The PPP hypothesis suggests that national price levels of two countries are equalized when expressed in the common currency unit. Even though this relation is not likely to hold in the short run due to price stickiness, it may hold in the long run. In past studies, the most popular way of testing the PPP hypothesis in the long run was examining the statistical properties of the real exchange rate. The real exchange rate, by definition, is the relative national price level of two countries expressed in common currency units. If the real exchange rate is stationary, deviations from the PPP are transitory and a mean reversion occurs. In such a case, the relative national price level is equalized in the long run and the PPP hypothesis holds in the long run. However, if the real exchange rate has a unit root, deviations from the PPP are permanent and a mean reversion does not occur. In such a case, the PPP does not hold in the long run.

Although many past studies have focused on the stationary and the unit root process, what if the real exchange rate process is better described by the locally persistent process? When the real exchange rate process is locally persistent (equations 2.1 and 2.2 with $0 < d < 1$), interesting implications are obtained for the PPP reversion and the PPP hypothesis. Real exchange rates have a root near unity and are highly persistent within a certain region, but such persistence disappears at longer horizons. Therefore, the PPP hypothesis holds in the long run, even though local persistence implies a near-unity autoregressive root and non-stationarity.

A long memory process (or a fractional integration) can be thought of as another process that lies between the conventional covariance stationary and unit root processes. Since some past studies such as Cheung and Lai (1993, 2001) and Diebold et al (1991) suggested that the real exchange rate follows the fractionally integrated process, we compare local persistent process with the fractionally integrated process, and try to draw implications on the PPP hypothesis. Importantly, the fractionally integrated process (FI(d)) is more appropriate to capture long-range dependence, and our locally persistent process (LP(d)) characterizes persistence within ranges of persistency. As we will explain, a fractional integration in the real exchange rate also imply the parity reversion and the validity of long-run PPP hypothesis, but shock dissipation would take a much longer time than in the case of local persistence, and we might say that, for a reasonably long period of time, or for finite horizons, full PPP reversion may not occur.

3.1. Temporal Dependence

As a comparative example, assume that real exchange rate is modelled by a fractional white noise process, which is a discrete time version of fractional Brownian motion. The process is defined as

$$(1-L)^d(y_i - \mu) = u_i, \tag{3.1}$$

where $E(u_i) = 0$, $E(u_i^2) = \sigma^2$, $E(u_i u_s) = 0$ for $s \neq i$, and where the parameter d is a noninteger. One says that the real exchange rate is weakly stationary for $d < 1/2$. When $d \in (0, 1)$, then we say that real exchange rates displays long-range dependence or long memory, whereas if $d < 0$ the process exhibits anti-persistency. Following the literature on parity reversion, we will focus only on cases where the parameter of fractional integration is $d \in (0, 0.5)$.[2]

We use the autocovariance function to describe the pattern of temporal dependence displayed by an economic time series. If we look at the lag h autocovariance function, the long-memory process has its autocovariance function proportional to $|h|^{2d-1}$, but the locally persistent process has autocovariance function proportional to $(e^{-c/n^d})^h$. If you consider the standardized locally persistent process[3], in which $c = -1$, then the autocovariance function will be proportional to e^{-h/n^d}. Figures 1 and 2 show the

[2] When d refers to the parameter of local persistence, the it may be less than unity and larger than zero.

[3] Throughout this paper, we will use the standardized locally persistent process to model the dynamic of real exchange rates

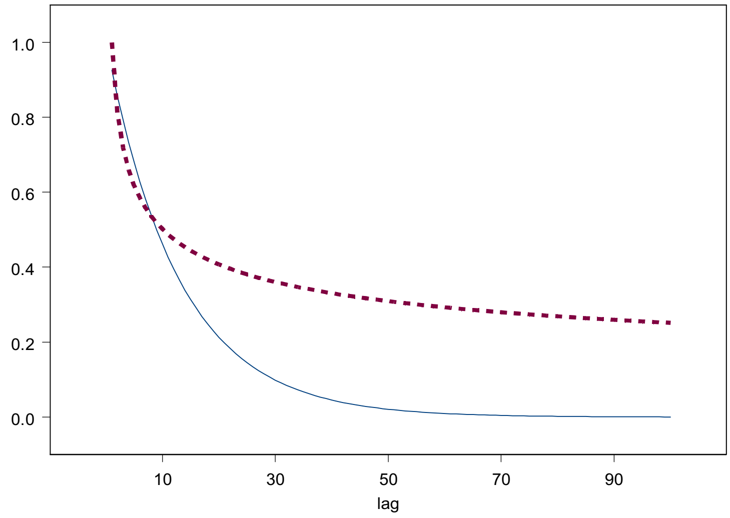


Figure 3.1: Temporal Dependence: FI(0.35) - dotted line; LP(0.35) - solid line.

graph of the functions $|h|^{2d-1}$(dotted line) and e^{-h/n^d} (solid line). We considered the cases $d = 0.35$ and $d = 0.45$ for the fractionally integrated process, where d is the parameter of fractional integration. As for the standardized locally persistent processes, we also considered $d = 0.35$ and $d = 0.45$, but now d represents the parameter of local persistence.[4]

The above figures reveal important facts on local persistence and long memory. First notice that the autocovariance of a standardized LP process converges to zero as the lag increases. Thus, if real exchange rate is locally persistent it will have temporal dependence (i.e., non-zero autocovariance) over a certain range, but not forever and it will eventually disappear. If we increases the parameter of local persistence from

[4] For the standardized locally persistent process, we set $n = 1500$, but the conclusion is not affected if you consider other sample sizes.

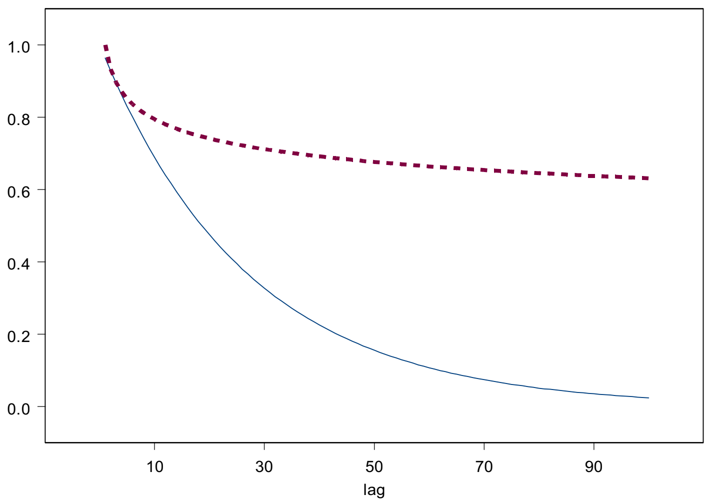


Figure 3.2: Temporal Dependence: FI(0.45) - dotted line; LP(0.45) - solid line.

d=0.35 to d=0.45, then we notice that the autocovariance function equals zero at a higher lag, indicating that the extension of the region of temporal persistency is determined by the magnitude of the local persistence parameter.

As for the long memory process, we can see that its autocovariance function decays hyperbolically and, therefore, much slower. Thus, if the real exchange rate exhibits long memory, then it will display temporal dependence over a long time period: the time series will have secular dependence. We believe that although real exchange rates have high temporal dependence, in the sense that they are highly correlated with past observations, this dependence may not be secular and, therefore, they are likely to disappear after some finite time. If it is the case, then the locally persistent process seems to be more adequate to model the dynamic of the real exchange rate.

3.2. Impulse Response Function

We discuss the impulse response function of a locally persistent and a fractionally integrated process, since the notion of PPP reversion is based upon the behavior of the impulse response function of the process describing the real exchange rate movement.

An impulse response function traces the effect of a shock in the innovation u_i on current and future values of the endogenous variables y_i . If the real exchange rate is stationary, $d = 0$, then its impulse response will converge to zero as the response horizon k increases and we say that the shocks are transitory and, therefore, the PPP hypothesis holds in the long run. On the other hand, when the real exchange rate has a unit root the impulse response never converges to zero. As an illustration, consider the following models:

Model 1: $y_i = \alpha y_{i-1} + u_i$, $u_i \sim iid$ and $|\alpha| \leq 1$.

Model 2: $(1-L)^d(y_i - \mu) = u_i$, $u_i \sim iid$ and $\alpha = 1$.

Model 3: $y_i = \alpha y_{i-1} + u_i$, $u_i \sim iid$, $\alpha = e^{-1/n^d}$.

In Model 1, the real exchange rate, y_i, is stationary when $|\alpha| < 1$. If the response horizon equals k, one can show that the impulse response at k-period horizon is: $IR_k = \dfrac{\partial y_k}{\partial u_1} = \alpha^{k-1} \rightarrow 0$ as $k \rightarrow \infty$. Therefore, if y_i is stationary, the shocks will be totally absorbed as k increases, and the LR PPP holds. Also notice that, according to model 1 $\sum\limits_{k=0}^{\infty} IR_k < \infty$, implying that total sum of the effects on the future values of the real exchange rate is a finite number. Now we consider Model 1 where y_i has a unit root, that is, $\alpha = 1$. In this random walk specification, the impulse response is: $IR_k = 1$ for any k. The shocks never vanish when there is a unit root. Thus, the LR PPP holds. The summation of the effects over time does not converge to a finite number, that is, $\sum_{k=0}^{\infty} IR_k = \infty$.

Model 2 represents a fractional white noise process. This process can be expressed as an infinite order moving average representation, or Wold decomposition,

$$y_i = \sum_{j=0}^{\infty} \Psi_j u_{i-j}$$

Therefore $IR_k = \dfrac{\partial y_k}{\partial u_1} = \Psi_k \approx 1/(k^{1-d})$ for $0 < d < 1$. So, the effect of innovations vanishes in the long run, but vanish slow because of the hyperbolically decaying of the impulse-response coefficient. Also notice that since $0 < 1 - d < 1$, we have $\sum_{k=0}^{\infty} IR_k = \infty$ so that the summation of the effects over time of a unit innovation is infinite. Therefore, at a long horizon, the total sum of the effects in the case of fractional integration is as strong as those in the unit root case.

Now, we analyze the behavior of the impulse response function of standardized locally persistent processes. Consider Model 3 with $0 < d < 1$. The k period impulse response is given by $IR_k = (e^{-1/n^d})^{k-1}$ and we can see that $IR_k \to 0$ if $k = O_p(n^{d+\epsilon})$ for any $\epsilon > 0$ ($IR_k \nrightarrow 0$ if $k = O_p(n^d)$ or $o_p(n^d)$). Thus, the impulse response does not go to zero in relatively shorter period, but does go to zero in longer horizon. In this sense, we say that if the real exchange rate is locally persistent, its impulse response eventually converges to zero and, therefore, the shocks are globally transitory and the LR PPP holds. In addition, the summation of the effects over time is finite, that is, $\sum_{k=0}^{\infty} IR_k < \infty$, implying that, at long horizon, the total sum of the effect is weaker in the case of local persistence than that in the case of the fractional integration.

Moreover, one can easily see that the degree of persistence is determined by the parameter of local persistence, d: the larger d is, the more persistent are the shocks and, therefore, we can affirm that shocks take much more time to die out when the real exchange rate is locally persistent ($0 < d < 1$) than when it is stationary ($d = 0$). That is, local persistency lies between the stationary and the unit root cases in terms of persistence of shocks.

Overall, we may say that a locally persistent process and a fractionally integrated process are similar in the sense that the are sitting in between the stationary and unit root extremes and their impulse responses converge to zero, but they converge slower than when the process is ergodic stationary. However, shocks affecting a FI(d) process takes much longer time to die out and, consequently, have a larger total sum of the effects than shocks affecting a LP(d) process. Therefore, the full mean reversion would take much longer for fractional integration than local persistence. In fact, due to the hyperbolically decaying of the impulse response coefficient Ψ_k, the fractional integration may imply that the full reversion would not occur within a reasonably long period of time or at a finite period of time.

4. Statistical Procedures

In this section, we discuss various statistical methods to detect whether a time series has a unit root, is ergodic stationary, or locally persistent. First of all, note that the conventional ADF test does not tell much. Either rejection or non-rejection of the unit root null does not necessarily imply that the process is stationary or unit root since we cannot exclude the locally persistent process.

First, we discuss the modified Dickey-Fuller test (DF-GLS), introduced by Elliot et al. (1992). It has good power to test the null hypothesis of an exact unit root against the local to unity alternative of stationarity. Next, we discuss the Lima-Xiao (2002) method to test the null hypothesis of standard stationary against local persistence. Finally, we discuss the method to estimate the parameter d of the locally persistent process when a series is a locally persistent process. As shown in the previous sections, the parameter d represents the persistency of the process.

4.1. DF-GLS Test

The DF-GLS test, introduced by Elliot et al. (1996) has a power function that is close to the power envelope. This implies that the DF-GLS test has better power against local to unity alternative. The version of the test that allows for a linear trend is based on the regression below:

$$(1-L)y_i^{\tau} = \phi_o y_{i-1}^{\tau} + \sum_{j=1}^{p} \phi_j (1-L) y_{i-j}^{\tau} + v_i \tag{4.1}$$

where v_i is an error term and y_i^{τ} is the locally detrended data process under the local alternative of $\alpha = 1 + c/n$. This is illustrated by:

$$y_i^{\tau} = y_i - z_i \beta_i \tag{4.2}$$

where $z_i = (1, i)$ is a linear trend and β is the least square regression coefficient of $\widetilde{y}_i$ on $\widetilde{z}_i$, where $\widetilde{y}_i = [y_1,\ (1-aL)y_2,\ ...,\ (1-aL)y_n]'$ and $\widetilde{z}_i = [z_1,\ (1-aL)z_2,\ ...,$ $(1-aL)z_n]'$. The DF-GLS statistic is given by the t-ratio, when testing $H_o : \phi_o = 0$ against $H_1 : \phi_o < 0$. Elliot et al (1996) recommended the localizing parameter, c, to be set equal to -13.5. For the specification without a linear trend but still with a constant, one should go over the same steps by replacing y_i^{τ} with the locally demeaned series y_i^{μ} and $z_i = 1$. For this case, one should now set $c = -7$.

4.2. Lima-Xiao Test

The Lima-Xiao test (Lima and Xiao, 2002) was designed to test the null hypothesis $H_o : d = 0$ against the alternative $H_a : 0 < d \leq 1$ in the local persistence model. The null hypothesis is tested by looking at the fluctuation of the real exchange rate, y_i. The measure of fluctuation used by them is the one suggested by Ploberger et al. (1986), that is:

$$Q_n = \underset{1 \leq v \leq n}{Max} \frac{1}{\sqrt{n}} \frac{1}{\widehat{w}_y} \left| \sum_{i=1}^{v} y_i - \frac{v}{n} \sum_{i=1}^{n} y_i \right| \tag{4.3}$$

where $\widehat{w}_y = \sum_{h=-q}^{q} (1 - \frac{|h|}{q+1}) \widehat{\gamma}(h)$, $\widehat{\gamma}(h) = \frac{1}{n} \sum_{i=1}^{n-h} y_i y_{i+h}$ is the long-run variance with truncation parameter q.

If the real exchange rate is not directly observable due to the presence of time trend, then the test statistic 4.3 will be evaluated at the detrended value of the real exchange rate $\widehat{y}_i$, that is

$$\widehat{Q}_n = \underset{1 \leq v \leq n}{Max} \frac{1}{\sqrt{n}} \frac{1}{\widehat{w}_y} \left| \sum_{i=1}^{v} \widehat{y}_i - \frac{v}{n} \sum_{i=1}^{n} \widehat{y}_i \right| \tag{4.4}$$

where $\widehat{y}_i$ is the detrended value of the real exchange rate. Under the null hypothesis, Lima and Xiao(2002) show that

$$\begin{aligned} Q_n &= \underset{1 \leq v \leq n}{Max} \frac{1}{\sqrt{n}} \frac{1}{\widehat{w}_y} \left| \sum_{i=1}^{v} y_i - \frac{v}{n} \sum_{i=1}^{n} y_i \right| \\ &\Rightarrow \underset{0 \leq r \leq 1}{\sup} \left| W_y(r) - r W_y(1) \right| \end{aligned} \tag{4.5}$$

$$
\begin{aligned}
\widehat{Q}_n &= \underset{1 \le v \le n}{Max} \frac{1}{\sqrt{n}} \frac{1}{\widehat{w}_y} \left| \sum_{i=1}^{v} \widehat{y}_i - \frac{v}{n} \sum_{i=1}^{n} \widehat{y}_i \right| \qquad (4.6) \\
&\Rightarrow \sup_{0 \le r \le 1} \left| W(r) - rW(1) + 6(1-r) \left\{ \frac{1}{2} W(1) - \int_0^1 W(s)ds \right\} \right|
\end{aligned}
$$

where $W_y(r) - rW_y(1)$ is a standard Brownian bridge. As showed in Billingsley (1968), the corresponding distribution function of the limiting variate, $\sup_{0 \le r \le 1} |W_\mu(r) - rW_\mu(1)|$, has the well-known Kolmogoroff form:

$$
\begin{aligned}
F(x) &= Pr\{ \sup_{0 \le r \le 1} |W_\mu(r) - rW_\mu(1)| \le x \} \qquad (4.7) \\
&= 1 + 2 \sum_{j=1}^{\infty} (-1)^j \exp(-2j^2 x^2) \text{ for } x \ge 0 \\
&\quad \textit{and} \\
F(x) &= 0 \text{ for } x < 0
\end{aligned}
$$

Consequently, if there is no time trend, then the above distribution can be used to derive the critical values for the Lima-Xiao test . Table II gives critical values for the test statistic Q_n.

Table II: Upper Tail Critical Values for Q_n

Level of Significance	0.1	0.05	0.01
Critical value	1.22	1.36	1.63

On the other hand, the test statistic $\widehat{Q}_n$ converges to the *sup* of the Brownian bridge plus a second term brought of a time trend t. Xiao (1999) calculated, via simulation, the critical values for the test statistic $\widehat{Q}_n$ which is reproduced in the Table below.

Table III: Upper Tail Critical Values for $\widetilde{Q}_n$

Level of Significance	0.1	0.05	0.01
Critical value	0.827	0.901	1.041

Lima and Xiao (2002) also showed that Q_n and $\widehat{Q}_n$ will diverge under the alternative hypothesis $0 < d \leq 1$, as $n \to \infty$.

4.3. Estimation of the Local Persistence Parameter.

According to locally persistent process, the extension of region of persistency is given by the magnitude of the parameter d. The greater the value of d, the longer the persistent range and the longer the persistent effect will last. Therefore, it turns out to be important to estimate the parameter d in order to identify the degree of local persistence of the stochastic process. In order to do so without changing the level of persistency (magnitude of d), Lima and Xiao (2002) suggested to standardize the localizing parameter. For convenience, they consider a standardized LP process, in which the localizing parameter, c, equals -1. Notice that $1 - \alpha = n^{-d}$, and after taking the logarithm , one obtains

$$d = -\frac{\ln(1-\alpha)}{\ln(n)} \tag{4.8}$$

After standardization, we have that:

$$n^{\frac{1}{2}+\frac{d}{2}}(\widehat{\alpha} - \alpha) \Rightarrow \xi = N(0, 2). \tag{4.9}$$

i.e.

$$n^{\frac{1}{2}-\frac{d}{2}}[n^d(\widehat{\alpha} - 1) + 1] \Rightarrow \xi \tag{4.10}$$

where

$$\widehat{\alpha} = \widehat{\alpha}_{OLS} - \frac{n\widehat{\lambda}}{\sum y_{i-1}^2}$$

Therefore, $\widehat{a}$ is made up of two components. The first one corresponds to the usual least-square estimator of a. The second one is a nonparametric correction that uses the consistent estimator of the one sided long run covariance parameter, $\widehat{\lambda}$.[5] The nonparametric correction is needed whenever we have a non i.i.d. innovation sequence. When the innovation sequence $\{u_i\}$ is independent and identically distributed, we have $\lambda = 0$.

The above result implies that

$$n^d(1-\widehat{\alpha}) \xrightarrow{P} 1 \qquad or \;\; \ln[n^d(1-\widehat{\alpha})] \xrightarrow{P} 0 \tag{4.11}$$

Hence, one can propose the following consistent estimator for d :

$$\begin{aligned} \widehat{d} &= -\frac{\ln(1-\widehat{\alpha})}{\ln(n)} \\ &= -\frac{\ln[n^d(1-\widehat{\alpha})] - d\ln(n)}{\ln(n)} \\ &= d - \frac{\ln[n^d(1-\widehat{\alpha})]}{\ln(n)} \\ &\rightarrow d \end{aligned} \tag{4.12}$$

[5] $\widehat{\lambda} = \frac{1}{2}(\widehat{\sigma}_\mu^2 - \widehat{\omega}_\mu^2)$, where $\widehat{\sigma}_\mu^2$ is a consistent estimator of the variance of μ_i and $\widehat{\omega}_\mu^2$ is a consistent estimator of the long-run variance of μ_i. In this paper, we consistently estimate ω_μ^2 by using nonparametric kernel smoothing.

5. Results

5.1. Basic Statistical Results

We used monthly data of the US-dollar based and Japanese-yen based bilateral real exchange rates. To construct the real exchange rate, the data on the nominal exchange rate and the price level (Consumer Price Index) are collected from the International Financial Statistics CD-Rom, which is made by the International Monetary Fund (IMF). The sample covers the Post-Bretton Woods period that runs from April 1973 to March 2001. Real exchange rates are used in logarithm form. Figure 5.1 shows graphs of 7 bilateral real exchange rates. Each series is plotted and centered with respect to its sample mean. One can see that US-dollar based real exchange rates display wide fluctuations, but there seems to be a mean reversion in most cases.

In order to control the possible forces that move the real exchange rate to a direction in the long run (such as Balassa-Samuleson effect), and to be consistent with past studies such as Cheung and Lai (2001), we decided to include a deterministic trend in the specification of Japanese-yen based real exchange rates. Figure 5.2 shows the detrended value for all the Japanese-yen based real exchange rates. One can see that all the detrended series display wide fluctuations but still are not clearly inconsistent with trend reversion.

Given the visual evidence of mean reversion in figures 3 and 4, we may expect the ADF unit root test to reject the null hypothesis of a unit root for most cases. However, as suggested by past studies, this visual impression of mean reversion (or trend reversion) has been hard to establish statistically using traditional unit root

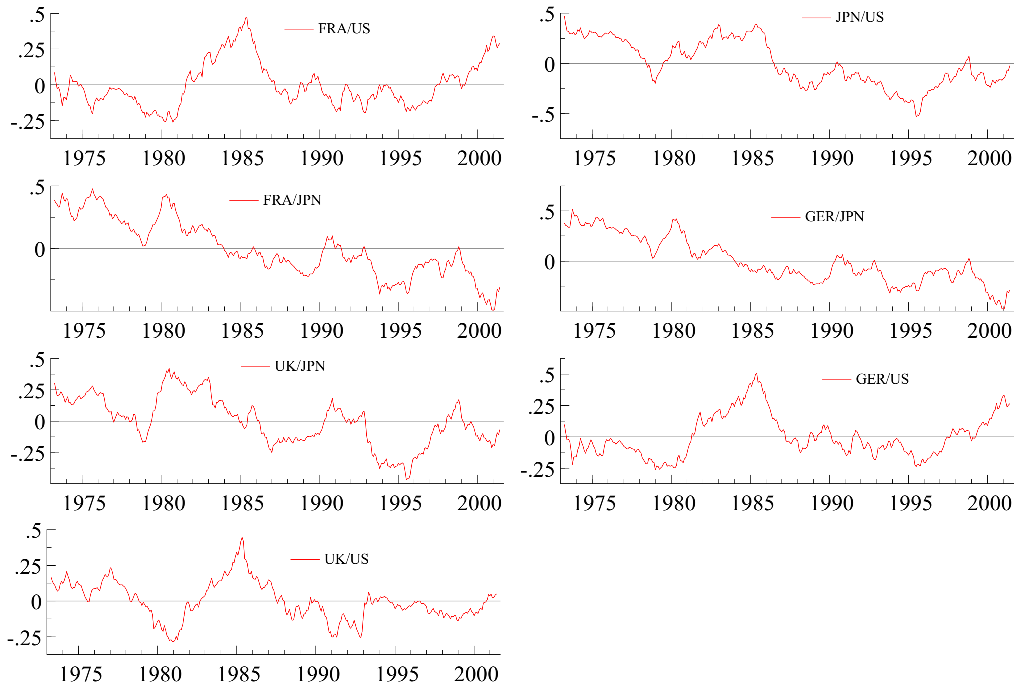


Figure 5.1: Log of Real Exchange Rates (centered)

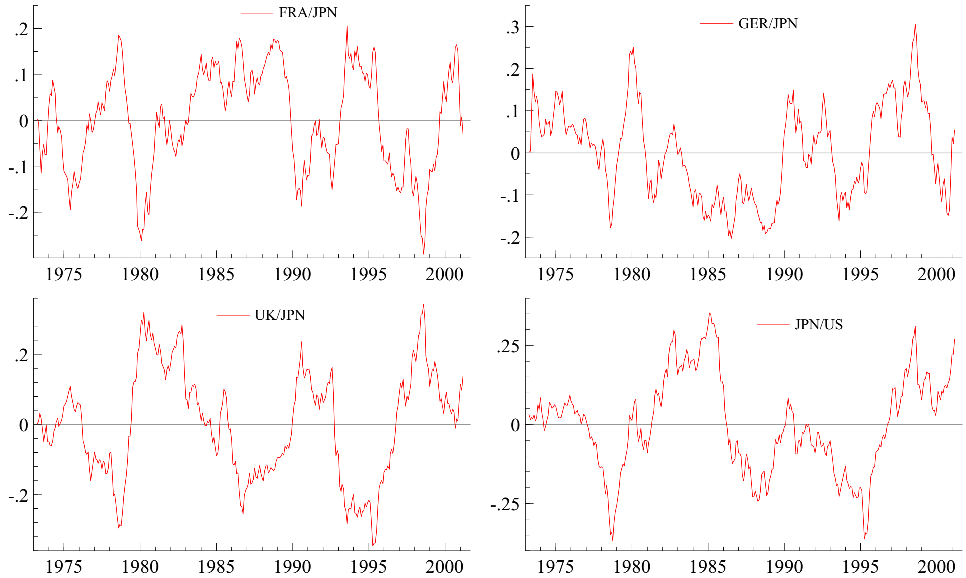


Figure 5.2: Log of Real Exchange Rates (detrended)

tests. The specification with trend and intercept is used for Japanese series, as we did when we plot the graphs. Table IV reports the results of the ADF and DF-GLS tests.[6] The notations '*', '**', and '***' suggest that the null hypothesis of a unit root is rejected at 10%, 5%, and 1%, respectively. The results show that the unit-root hypothesis is rejected in the ADF test for only the France/Japan and Germany/Japan real exchange rates.

However, Table IV indicates that the autoregressive roots are near unity and, as suggested by past studies like Campbell and Perron (1991), Rogoff (1996) and Froot, and Rogoff (1995), the failure to reject the unit root null might be due to the low power of the ADF test when the process has root near unity. In fact, when we apply the DF-GLS test, we find much stronger evidence against unit root: DF-GLS test rejects the unit root in all cases. Rejection of an exact unit root in favor of a root near unity for real exchange rates has already been reported in the literature. Cheung and Lai (1998) used the DF-GLS test to show that the hypothesis of an exact unit root is not accepted for most real exchange rates. MacDonald (1985,1995) used the variance ratio test to suggest that real exchange rates are highly persistent, but do not follow a random walk process.

[6] We used the same lag choice for the ADF and DF-GLS tests, that is, the choice based on the Modified Information Criteria (MIC) suggested by Perron and Ng (2001). Unlike the Akaike or Schwarz criteria, MIC imposes the null hypothesis of unit root into the objective function used to calculate the optimal lag. As a matter effect, AIC and SIC tend to choose small lags whereas MIC tends to choose larger lags.

Table IV:Unit-Root Tests

Series	Specification	Lags	ADF Test	$\widehat{\alpha}$	DF-GLS
France/USA	Intercept	5	-1.85	0.98	-1.84**
Germany/USA	Intercept	8	-1.77	0.99	-1.74**
UK/USA	Intercept	6	-2.45	0.98	-1.89**
Japan/USA	Linear Trend	11	-2.57	0.98	-2.64*
UK/Japan	Linear Trend	8	-2.88	0.98	-2.91**
France/Japan	Linear Trend	7	-3.65**	0.96	-3.66***
Germany/Japan	Linear Trend	8	-3.73**	0.96	-3.74***

From the results of the DF-GLS test, we may tempt to say that the log of real exchange rates does not have a unit root, and the real exchange rate is a stationary process. However, we cannot simply conclude in that way since the real exchange rate may follow the process in between, for example, a locally persistence process.

Table V reports the results for test of the null $H_o : d = 0$ against the alternative $H_1 : 0 < d \leq 1$, as well as point estimates of the local persistence parameter, d. As suggested by Lima and Xiao (2002), we use the sample dependent truncation-lag $q = [\ln(n)]$, where [.] signifies an integer number. Again, the notation '*', '**', and '***' suggest that the null hypothesis, $d = 0$, is rejected at 10%, 5% and 1% level of significance, respectively. First, the results reported in Table V indicate that the data uniformly reject the stationarity null hypothesis, i.e. $d = 0$ against the alternative $0 < d \leq 1$. In addition, all the series have an estimated local persistence parameter, d, different from zero. Combining this result with the previous evidence displayed in

Table IV, our empirical analysis indicates that the real exchange rate process is likely to be local persistence, which lies between the conventional stationary and the unit root process.

Since the real exchange rate displays local persistence, a shock dissipation will take place in the long run, and the long run PPP hypothesis holds as in the stationary case. However, shocks would take more time to die out than the stationary case.

Table V. local persistence Analysis

RER	$\widehat{d}$	Model	$H_o : d = 0$
France/USA	0.80	Intercept	1.54**
Germany/USA	0.77	Intercept	1.68**
UK/USA	0.68	Intercept	1.87***
Japan/USA	0.63	Linear Trend	1.18***
Japan/UK	0.58	Linear Trend	0.98**
Japan/France	0.49	Linear Trend	0.94**
Japan/Germany	0.50	Linear Trend	1.17***

5.2. Local Persistence vs. Long Memory

In the previous section, we present a formal statistical evidence to support the local persistence in the real exchange rate, against the stationarity and the unit root. In this section, we present empirical evidence supporting local persistence in the real exchange rate, against long memory process that is another process that lies between stationary and unit root. Although a formal statistical procedure to test whether a series is locally persistent or fractionally integrated is not developed yet, we provide informal evidence by plotting the autocorrelogram functions.

The next figures show the autocorrelogram functions of the US-dollar based real exchange rates and the detrended values of the Japanese-yen based real exchange rates. As discussed in Section 2.1., the autocorrelogram of a locally persistent process converges to zero at a finite lag, whereas the one of a fractional white noise decays hyperbolically and, therefore, take a very long time to converge to zero, characterizing the existence of secular dependence. The graphs seem to suggest the presence of MA components, which is accounted for our model of local persistence since it assumes that the innovation process u_i is a linear process. Mostly importantly, the graphs suggest that the autocorrelogram converges toward zero at a finite lag, which is inconsistent with the existence of secular dependence in real exchange rate. This supports the existence of local persistence in the real exchange rate, rather than the fractional integration.

Local persistence implies that real exchange rates are nonstationary but still exhibit mean reverting behavior in the sense that the impulse response converges to zero. Moreover, local persistence behavior excludes the existence of secular dependence and infinite total impact of a unit innovation. Although fractional models also imply long run PPP reversion, the full reversion may take a very long time. Thus, for reasonable finite horizons, we might say that full PPP reversion does not occur if the real exchange rate is fractionally integrated. On the other hand, local persistency may rather imply PPP reversion within reasonably long finite horizons, though still it takes longer than the stationary case. Overall, our study suggest that PPP hypothesis holds in the long run, with PPP reversion occurring slower than in the stationary case but quicker than predicted by fractional integration and unit root. In fact, the visual impression in Figures 3 and 4 suggests that there appears to be a mean (trend) reversion in the real exchange rate, but it does not seem to take too long as predicted by fractional integration. Therefore, the presence of local persistence in real exchange rates seems to match with the visual evidence yielded by the real exchange rate.

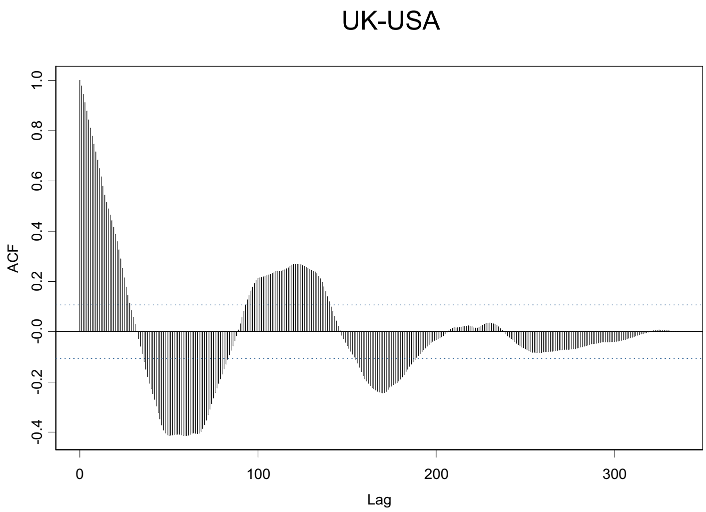


Figure 5.3:

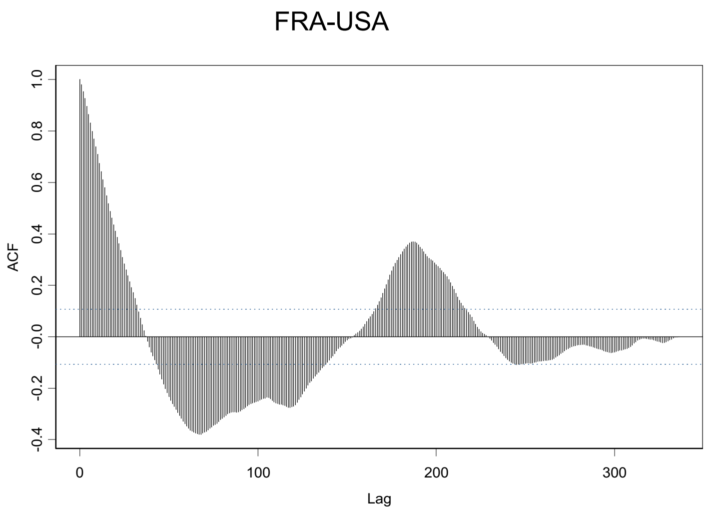


Figure 5.4:

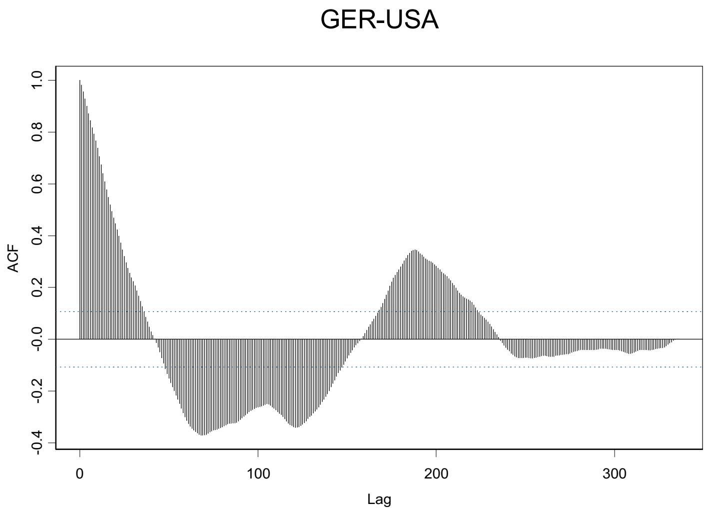
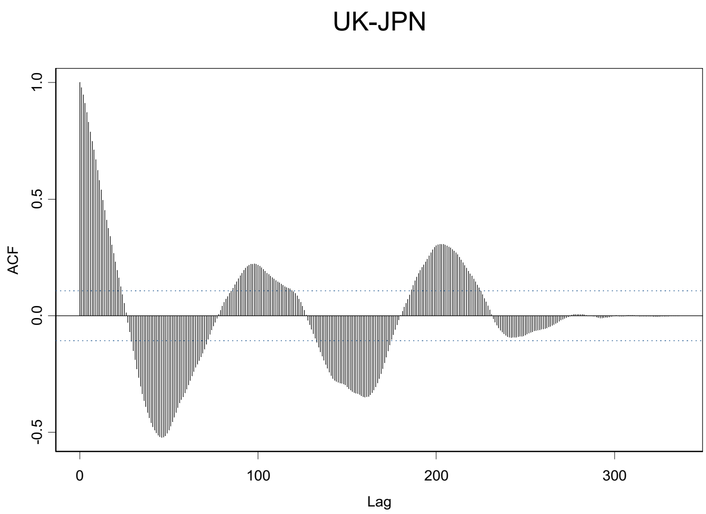


Figure 5.5:



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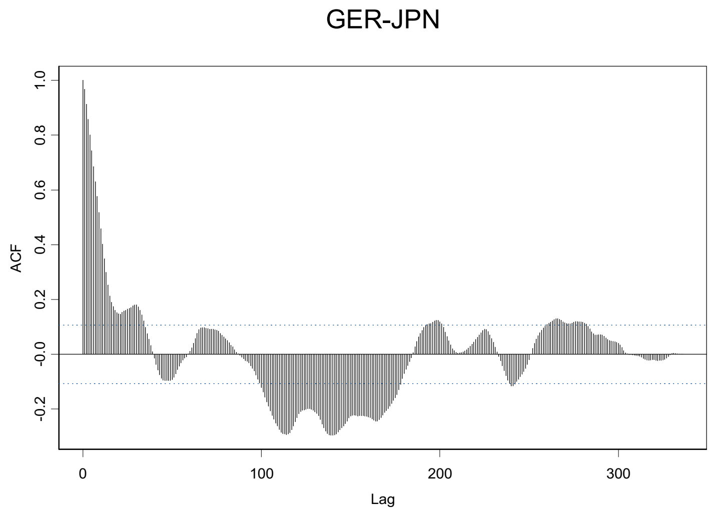

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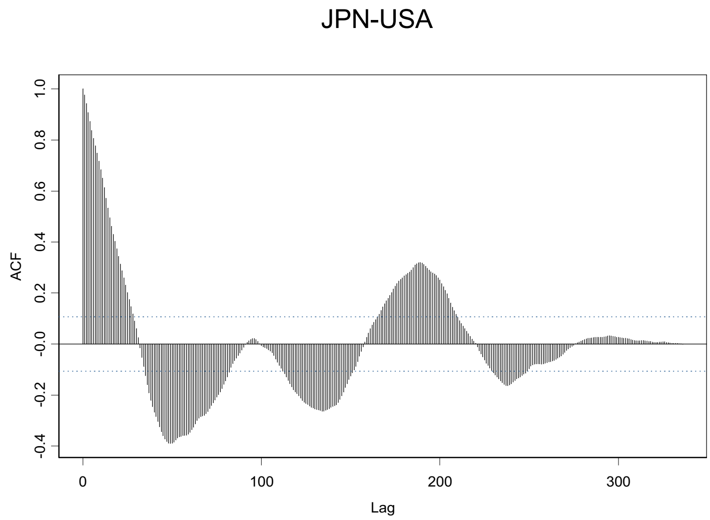

Figure 5.8:

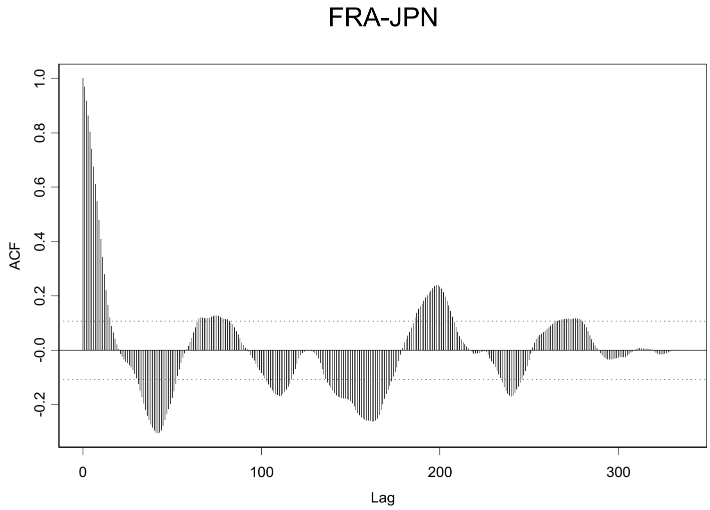

Figure 5.9:

6. Conclusion

We examine whether the PPP hypothesis holds in the long run by examining the statistical properties of the bilateral real exchange rates of G-5 countries for the post Bretton-Woods period. In contrast to most past studies that restricted their focus to the unit root and the stationary process, we also consider some processes in between, such as the locally persistent and the fractionally integrated processes. Both processes imply that a mean reversion occurs in the long run. Therefore, if the real exchange rate is described by either process, the PPP would hold in the long run, although a full reversion would take more time than the stationary case. For fractional integration, as displayed in the hyperbolic decaying of its autocovariance and impulse responses, a full reversion does not occur for the finite time horizon or within a reasonably long period of time, which is an important difference between the implication of two processes on the PPP.

The result suggests that the real exchange rate may be better described by a locally persistent process; the persistence found in the real exchange rate is stronger than the stationary case but weaker than the unit root case. Although most past studies focused on the unit root and the stationary process, the real exchange rate movements may be described by neither a clear-cut unit root nor a clear-cut stationary process. This result may explain why past studies could not provide a clear-cut conclusion on the real exchange rate processes and the PPP hypothesis.

The empirical result implies that the PPP holds in the long run, although the persistence found in the real exchange rate is stronger than the persistence of a stationary process. By further supporting the local persistence in the real exchange rate against the fractional integration based on autocorrelogram of the real exchange rate, we suggest that the persistence in the real exchange rate is not extreme like the fractional integration case; we exclude the possibility of the secular dependence, hyperbolic decaying of the impulse response coefficient, and an infinite total sum of the effects of a unit innovation. Instead, we suggest that the real exchange rate evolves over time in such a way that there is persistency over a range of time, but the effect of shocks will eventually disappear over time horizon longer than order $O(n^d)$, that is, at a finite time horizon.

From the theoretical econometricians' perspective, this paper can be regarded as an initial attempt to apply locally persistent process to the empirical analysis. Further theoretical developments of the local persistence, in addition to further empirical applications, would be fruitful. Finally, further investigations on the PPP hypothesis are also important, to accumulate more evidence to end the continuing debate on the empirical relevance of the PPP hypothesis.

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