Long-Term Dependence in the Foreign Exchange Markets: International Evidence

SANG HOON, KANG*

This article examines nine Pacific-rim country currency markets for evidence of long-term dependence in terms of two different techniques: the modified R/S analysis and the GPH test. The empirical tests are conducted using weekly and monthly data during the post-Bretton Woods period from January 1974 to December 2004. Although there is evidence of long-term dependence in some country exchange returns, it is not convincing. This finding is supportive of the efficient market hypothesis (EMH), but it is inconsistent with previous literature, which found significant evidence of long-term dependence in the major foreign currency markets. Therefore, the result of long-term dependence is sporadic and unstable in foreign exchange rate markets.

JEL Classifications: C22, F31

Key words: Exchange rate; Long-term dependence; Modified R/S analysis; GPH test

* **Correspondence**: School of Commerce, University of South Australia, GPO BOX 2471, Adelaide, SA, 5001, Australia, Phone: +61 8 83020577 Fax: +61 8 83020992, E-mail: sang.kang@unisa.edu.au

I. INTRODUCTION

Exchange rate dynamics have significant economic implications for both theoretical and empirical levels. For instance, exchange rate dynamics are important determinants of international trading flows, prices of goods and services, and international asset portfolios. Despite hedging some short-term foreign exchange risks in derivatives markets, market participants may be confronted by the fluctuations and uncertainty of long-term exchange rates. Such long-term exchange rate behaviours would affect investment decisions and distort the optimal allocation of resources (Barkoulas et al. 2003).

Early studies of exchange rates question whether exchange rates follow a random walk or not. Hokkio (1986) indicates that exchange rates follow a random walk under the specific condition that the market is fundamental. On the other hand, Baillie and Bollerslev (1989) test univariate properties of exchange rates and find a unit root in the major exchange rates, indicating that exchange rates follow a random walk. Such previous studies focus only on the short-term dependence in foreign exchange rates. Recently, researchers have become interested in long-term dependence due to several factors associated with the currency crisis such as central bank interventions, monetary changes and exchange regime changes. Long-term dependence describes a correlation structure of series which displays persistent dependence between observations distant, discrete and far apart in time. In contrast, if correlations among observations become negligible at long lags, the series is said to exhibit short-tem dependence. Thus, long-term dependence exhibits that shocks to returns tend to decay much more slowly than short-term dependence where shocks to returns tend to decay rapidly.

Long-term dependence is also a special form of nonlinear dynamics that describes the autocorrelation structure of returns. As far as linear behaviour is concerned, all market participants react to information instantaneously as soon as information is received. This linear behaviour assumes that the returns are identically and independently distributed (i.i.d.). However, non-linear behaviour states that each market participant has heterogeneous trading horizons. For example, when the information arrives at foreign exchange markets, some people react to the information immediately, while others delay their reactions until a trend is well established (Skjeltorp 2000). Such non-linear behaviour gives rise to a question about the distribution of returns. It is widely accepted that returns have a leptokurtic distribution which has fatter tails and is more peaked at the mean than a normal distribution due to the interaction between market participants with different investment horizons. Moreover, the presence of long-term dependence indicates evidence of a predictable component for future returns. Predictability would question the validity of the weak-form efficiency of foreign exchange markets, because exchange rate returns having long-term dependence would allow market participants to anticipate future price movement and consistently earn abnormal returns. Thus, the presence or absence of long-term dependence may test the efficient market hypothesis (EMH).

Much literature has been dedicated to support long-term dependence in exchange rate dynamics since the breakdown of the Bretton Wood system in 1973. Initially, Booth et al. (1982) examine three major daily spot exchange rates, namely the British pound, French franc, and German mark, using a classical rescaled range analysis. They find evidence of positive long-term dependence during the flexible exchange rate regime and negative long-term dependence during the fixed exchange rate regime. Cheung (1993) finds convincing evidence of long-term dependence in five major weekly

nominal exchange rates from 1974-1987 using the popular GPH method. Pan et al. (1996) employ Lo's modified rescaled range analysis in five major weekly exchange rates. They provide evidence of long-term dependence in weekly nominal exchange rates. Mulligan (2000) also applies the modified rescaled range analysis in twenty-two monthly foreign exchange rates and finds significant evidence of long-term dependence in all currencies. These findings are not supportive of the EMH. However, Barkoulas et al. (2003), applying the Gaussian semiparametric estimator developed by Robinson (1995), find little evidence of long-term dependence in weekly nominal exchange rates for eighteen industrial countries for the period 1974-1995. They conclude that exchange rates are best characterised as a martingale, rather than a fractionally integrated process. Therefore, existing studies on long-term dependence of exchange rates are inconclusive.

The aim of this study is to examine long-term dependence in the nominal exchange rate returns for nine Pacific Rim countries, namely, Australia, Japan, Hong Kong, Malaysia, New Zealand, Singapore, South Korea, Taiwan and Thailand. The focus on Pacific Rim exchange rates is appropriate for two reasons. First, most existing literature has considered major exchange markets, while only a few papers have studied the Pacific Rim exchange markets because of different currency management regimes. However, the Pacific Rim exchange markets have been liberalised since the Asian currency crisis. In October 1997 some of the most important Pacific Rim exchange markets faced a severe currency crisis, when a Russian default in its national debt drove the domestic currency market into turmoil. This caused most of exchange markets around the world to experience severe loses resulting small investors as well as financial institutions to experience some very negative yields in their market portfolios. These markets are stated to recover after several months of negative records

and heroic efforts by International Institution (IMF) and National Governments to re-establish reliability in exchange rate markets. Thus, the results of the Pacific Rim exchange rate markets may be different from those of the major exchange markets. Second, the same region markets allow comparison of the degree of market efficiency depending on the absence or presence of long-term dependence. For example, both Japanese and Australian exchange rate markets are relatively considered as relatively efficient market, while other exchange rate markets are classified as inefficient. Thus, understanding the exchange rate dynamics of these markets is an important undertaking.

Two different methods are applied to investigate long-term dependence in foreign exchange markets. The first method, Lo's modified rescaled range (R/S) analysis, 'is not sensitive to non-normality and conditional heteroskedasticity in the data' (Cheung and Lai 1995, p598). This method is also not sensitive to short-term dependence and provides a non-periodic or irregular cycle length. In addition to the modified R/S analysis, the fractional differencing test for long-term dependence proposed by Geweke and Porter-Hudak (1983) is employed in this study. The fractional differencing test allows for distinction between short-term dependence and long-term dependence. The details of the methodology framework are outlined in the following section.

The rest of this study is organised as follows. The second section presents two research methodologies, such as the modified rescaled range analysis and the Geweke and Porter-Hudak estimator. The third section details the data and presents the results. The final section provides some conclusions.

II. Methodology

This section provides a review of two independent methods used in this study. Both the modified R/S analysis and the fractional different test examine the null hypothesis of short-term dependence against long-term dependence as the alternative hypothesis.

1. Modified R/S analysis

The R/S analysis was first developed by Hurst (1951). The aim of the R/S analysis is to detect evidence of long-term dependence in a time series. The R/S analysis is useful to detect long-term dependence in highly non-Gaussian distribution which exhibits heavy tail with infinite variance (Willinger et al. 1999). However, Lo (1991) argues that the R/S analysis is sensitive to short-term dependence, which can give a bias toward accepting the null hypothesis of long-term dependence. Thus, the R/S analysis is not a reliable tool to test long-term dependence in a time series if the series exhibits short -term dependence.

To overcome this problem, Lo (1991) proposed a modified R/S analysis which examines the null hypothesis of short-term dependence. Let $\bar{X}$ represent the sample mean of a foreign exchange rate return series $\{X_1, X_2, \ldots, X_N\}$. The modified R/S statistic, denoted as Q_N, is given by the range of partial sums of deviations from the mean rescaled by its standard deviation.

$$Q_n = \frac{1}{(\sigma_n(q))}\left\{\max_{1\le k\le N}\sum_{n=1}^{k}\left(X_n - \bar{X}\right) - \min_{1\le k\le N}\sum_{n=1}^{k}\left(X_n - \bar{X}\right)\right\} \tag{1}$$

where

$$\sigma_n^2(q) = \frac{1}{n}\sum_{n=1}^{n}\left(X_n - \bar{X}\right)^2 + \frac{2}{n}\sum_{n=1}^{q} w_n(q)\left\{\sum_{t=n+1}^{n}\left(X_t - \bar{X}\right)\left(X_{t-n} - \bar{X}\right)\right\} \tag{2}$$

$$= \sigma_x^2 + 2\sum_{n=1}^{q} w_n(q)\gamma_n \tag{3}$$

and σ_x^2 and γ_n represent the sample variance and autocovariances respectively. The weights $w_n(q)$ are given by

$$w_n(q) = 1 - \frac{n}{q+1}, \quad q < N \tag{4}$$

Under the null hypothesis of short memory, Lo (1991) defined the modified R/S statistic, $V_n(q)$, by setting

$$V_n(q) = \frac{1}{\sqrt{n}} Q_n \overset{a}{\sim} V, \tag{5}$$

in which the distribution function F_V of V is given by

$$F_V(\upsilon) = 1 + 2\sum_{k=1}^{\infty}\left(1 - 4k^2\upsilon^2\right)\exp^{\left(-2(k\upsilon)^2\right)} \tag{6}$$

The critical values of significant levels are computed from Equation (6) and tabulated by Lo (1991) for the purpose of the hypothesis test under the null hypothesis of short-term dependence against long-term dependence alternatives. Table 1 is reproduced from Lo (1991, p. 1288).

Table 1. Critical value of the distribution $F_V(\upsilon)$

$P(V < \upsilon)$	.005	.025	.050	.100	.200	.300	.400	.500
υ	0.721	0.809	0.861	0.927	1.018	1.090	1.157	1.223
$P(V < \upsilon)$	0.543	0.600	.700	.800	.900	.950	.975	.995
υ	$\sqrt{\frac{\pi}{2}}$	1.294	1.374	1.473	1.620	1.747	1.862	2.098

Lo tested the null hypothesis of short memory at the 95% level of confidence using the interval [0.809, 1,862] in Table 1. The null hypothesis is

$$H_0 = \{\text{short memory, i.e., } H = 0.5\}$$

against the alternative

$$H_1 = \{\text{there is long memory. i.e., } 1/2 < H < 1\}$$

The only difference between the classical and modified R/S analysis is that 'the denominator in Equation (1) normalises the range measure not only by the sample variance $(q = 0)$, as considered in the classical R/S analysis, but also by weighted sum of sample autocovariances for $q > 0$' (Cheung and Lai, 1995 p601).

2. Fractional differencing test

An alternative method of testing long-term dependence used in this study is that a fractionally integrated process is based on a fractionally differencing parameter d. The model of an ARFIMA process has been widely studied in literature (Granger and Joyeux, 1980; Hosking 1981), and is described as:

$$B(L)(1-L)^d y_t = D(L)\varepsilon_t, \quad \varepsilon_t \sim i.i.d.(0, \sigma_u^2) \tag{7}$$

where L is the backward-shift operator; $B(L) = 1 - \beta_1 L - \ldots - \beta_p L^p$ and $D(L) = 1 + \delta_1 L + \ldots + \delta q L^q$ are polynomials with stable roots; ε_t is white noise; d is the fractional order parameter and $(1-L)^d$ is a fractional differencing operator defined as:

$$(1-L)^d = \sum_{r=0}^{\infty} \frac{\Gamma(r-d)L^r}{\Gamma(-d)\Gamma(r+1)} \tag{8}$$

$\Gamma(\cdot)$ denotes the gamma function. The parameter d is allowed to take on any real value. The fractionally integrated process is stationary and invertible with $-0.5 < d < 0.5$. More precisely, the autocorrelation function of ARFIMA process slowly decays at a hyperbolic rate to zero rather than the exponential decay of the ARMA process (Hosking 1981). While $|d| > 0.5$, this process is non stationary as it has infinite variance. For $0 < d < 0.5$, the ARFIMA process displays persistence. In contrast, the fractionally integrated process with $-0.5 < d < 0$ has anti-persistence.

To obtain an estimate of the fractional differencing parameter d, let $I(\xi)$ be the periodogram of X at angular frequency ξ defined by

$$I(\xi) = \frac{1}{2\pi T}\left|\sum_{t=1}^{T} e^{it\xi}\left(X_t - \bar{X}\right)\right|^2 \tag{9}$$

Then the least squares regression is defined by

$$\ln\{I(\xi_\lambda)\} = \beta_0 + \beta_1 \ln\left\{4\sin^2\left(\frac{\xi_\lambda}{2}\right)\right\} + \eta_\lambda, \quad \lambda = 1,\ldots,\nu \tag{10}$$

where $\xi_\lambda = \frac{2\pi\lambda}{T}(\lambda = 0,\ldots,T-1)$ denotes the Fourier frequencies of the sample, T is the number of observations, and $\nu = g(T) << T$ is the number of low frequency peridogram ordinates included in the spectral regression. Here, β_0 is a constant, and η_λ is independently, identically distributed. Thus, the slope coefficient d is provided by the OLS estimator of $-\beta_1$. Theoretically, the variance of η_λ is equal to $\pi^2/6$ and $g(T)$ is commonly expressed as $T^{0.5}$. Geweke and Porter-Hudak (1983) suggested that the theoretical variance of ε_j in spectral regression is known to be equal to $\pi^2/6$, which enhances the efficiency of estimation. They also showed that the hypothesis test with regard to the value of d can be based on the t statistics of the regression coefficient. Since Geweke and Porter-Hudak (1983), several researchers have

modified the GPH estimator due to its poor finite sample size (Sowell 1992; Agiakloglou et al.1992). Agiakloglou et al. (1992) argued that when either the autoregressive or moving average possesses non-zero value of parameters in the ARFIMA model, the periodiogram regression estimator of d presents bias and inefficiency. Nevertheless, the GPH test continues to be widely used in applied work (Cheung and Lai 1993; Barkoulas et al. 2000) because of its computational simplicity, i.e. without the autoregressive and moving average parameters.

III. DATA AND EMPIRICAL RESULUTS

1. Data

The data set consists of U.S dollar nominal exchange rates for nine Pacific-rim countries: Australian dollar (AUD/USD), Hong Kong dollar (HKD/USD), Japanese Yen (JPY/USD), New Zealand dollar (NZD/USD), Malaysian ringgit (MYR/USD), Singapore dollar (SGD/USD), South Korean won (KRW/USD), Taiwan dollar (TWD/USD), and Thai baht (THB/USD). All sample foreign exchange rates are obtained from the database of the Federal Reserve Bank of St. Louis for a selection of countries.[1] The frequencies of observation are examined in two groups: weekly and monthly data. These weekly exchange rates represent Friday noon-time bid prices in New York City for cable transfers payable in foreign currencies. When Friday prices are not available, Thursday prices are used (The construction of the data set of weekly observations follows Barkoulas et al (2003)). Furthermore, these monthly exchange rates represent average daily figures, provided by the Federal Reserve Bank of St.

[1] Federal Reserve Economic Data (or FRED®) can be found on the internet http://research.stlouisfed.org/fred2/search/exchange+rate/1

Louis. Both weekly and monthly nominal exchange rates used in this study provide more reliable results for testing the nature of long-term dependence rather than daily nominal ones because daily basis data generally exhibit significant autoregressive tendencies. Thus, these tendencies can bias the estimates of long memory tests.

The sample period covers the post-Bretton Woods period, beginning in January 1974 through December 2004, 30 years of data. For some countries, data is only available for a sub-sample. This study considers a longer, broader recent sample than previous studies. All weekly and monthly price series are converted into the first logarithmic differences (returns) of the exchange rate series,

$$X_t = \ln\left(P_t \big/ P_{t-1} \right) \tag{11}$$

where X_t is the returns for foreign currency at time t, and P_t is the current price and P_{t-1} is the previous day's price. Table 2 presents the descriptive statistics for both weekly and monthly nominal exchange rate returns. Specially, the skewness and kurtosis of the returns are computed. Skewness measures the extent to which a distribution is not symmetric about its mean value and kurtosis measures how fat the tails of the distribution are. For a normal distribution, skewness and kurtosis coefficient are zero and 3, respectively.

As shown in Table 2, both the weekly and monthly nominal exchange rate returns display a similar pattern of results. In the case of skewness, the distributions of the weekly nominal returns for Japan and New Zealand are negatively skewed, others are positively skewed. The distributions of the monthly nominal returns for Japan, New Zealand, and Singapore are negatively skewed, others are positively skewed. In the case of kurtosis, kurtosis for all returns is highly greater than that of the normal distribution. Thus, the returns are extremely abnormal. Such skewness and kurtosis are

common characteristics in return distribution, which appear to be leptokurtoic. The statistics of Jarque-Bera test indicate that the null hypothesis of normality should be rejected at the 5 % level.

In addition, the Ljung-Box test is to check for serial correlation in both weekly and monthly nominal returns. The Ljung-Box test statistics provide evidence of possible correlation in the first and higher moments of the return distributions. Under the null hypothesis of no serial correlation, the $Q(5)$ and $Q(10)$ statistics are distributed asymptotically as a x^2 distribution (chi-square) with 5 and 10 degrees of freedoms respectively. According to the $Q(5)$ and $Q(10)$ statistics, there is significant evidence of serial correlation for seven (Hong Kong, Japan, Malaysia, Singapore, South Korea, Taiwan and Thailand) weekly nominal returns, while all monthly nominal returns show strong evidence of serial correlation. Thus, such a serial correlation indicates that returns are correlated with each other and show non-linear dependence.

Table 2. Descriptive statistics for all nominal exchange rate returns

Country	Mean	Std. Dev.	Skew	Kurt	J-B	$Q(5)$	$Q(10)$
(a) Weekly nominal exchange rate returns							
Australia	0.0004	0.014	3.204	39.542	92732.75 (0.000)	6.02	7.77
Hong Kong	0.0003	0.006	1.832	80.566	314562.2 (0.000)	18.79**	25.20**
Japan	-0.0006	0.015	-0.910	10.497	4009.87 (0.000)	13.89**	17.62
Malaysia	0.0003	0.013	0.792	102.830	671634.9 (0.000)	143.48**	158.94**
New Zealand	-0.0004	0.015	-3.350	42.052	105774.4 (0.000)	5.57	6.79
Singapore	-0.0002	0.007	0.952	22.274	19551.68 (0.000)	18.56**	31.03**
South Korea	0.0003	0.015	8.660	214.73	2325989 (0.000)	62.81**	205.96**
Taiwan	-0.0002	0.006	0.714	17.424	9856.06 (0.000)	71.71**	90.54**
Thailand	0.0006	0.012	3.501	56.309	153937.4 (0.000)	38.98**	52.96**
(b) Monthly nominal exchange rate returns							
Australia	0.0018	0.024	1.120	7.081	335.06 (0.000)	32.20**	41.13**
Hong Kong	0.0014	0.010	3.135	30.641	9606.37 (0.000)	36.45**	54.21**
Japan	-0.0028	0.029	-0.540	3.874	29.81 (0.000)	48.04**	60.26**
Malaysia	0.0011	0.019	0.992	29.267	10726.61 (0.000)	29.35**	36.98**
New Zealand	-0.0018	0.026	-0.846	7.904	415.96 (0.000)	48.22**	66.88**
Singapore	-0.0008	0.013	-0.055	6.305	130.75 (0.000)	31.38**	38.56**
South Korea	0.0015	0.029	7.574	97.839	109148.8 (0.000)	75.29**	86.25**
Taiwan	-0.0008	0.013	0.416	7.604	231.63 (0.000)	55.06**	82.80**
Thailand	0.0022	0.027	2.963	28.520	8208.45 (0.000)	23.17**	37.09**

Notes: This table describes several descriptive statistics, including mean, standard deviation, skewness, kurtosis, Jarque-Bera (J.B.) test and Ljung-Box test for individual stock returns. P-values are in brackets. Under the null hypothesis for normality, the Jarque-Bera statistic is distributed as $x^2(2)$. In the columns for $Q(n)$, the Ljung-Box statistic for returns up to n-th order of serial correlation. Critical values are 11.1 and 18.3 for n=5 and 10, respectively, at 5% significance. ** indicates statistical significance at the 5% level.

To test for stationarity, all exchange rate returns are subjected to three unit root tests, namely the ADF (augmented-Dickey-Fuller), PP (Phillips-Person) and KPSS

(Kwiatkowski, Phillips, Schmidt and Shin). The tests differ in the null hypothesis. The null hypothesis of the ADF and PP tests is that a time series contains a unit root, I (1) process, while the KPSS test has the null hypothesis of stationarity, I (0) process. The latter serves as a complement to the ADF and PP tests. Comparing the results of unit root tests under the different null hypothesis is characterised by four possible outcomes (Barkoulas et al. 1997): (1) when the null hypothesis of the ADF and PP tests is rejected and that of the KPSS test is not rejected, a time series is stationary. (2) Conversely, failure to reject a unit root by the ADF and PP tests and rejection stationary by the KPSS test supports that a time series is non-stationary. (3) Failure to reject a unit root and stationary null hypotheses shows that the series are not sufficiently informative with respect to the low-frequency properties. (4) Rejection of null hypotheses indicates that a series is not well represented as either I (0) or I (1), and fractional integration may be more appropriate.

Table 3 presents the empirical results of stationary tests for all nominal exchange rate returns. On the one hand, the ADF and PP tests have a unit root under the null hypothesis. If the test statistics are more negative than the critical value, the null hypothesis of a unit root should be rejected. The results of the ADF and PP tests, computing the statistics with and without trend, indicate that the non-stationarity under the null hypothesis is rejected at the 1% significance level. Thus, all nominal exchange rate returns are stationary. On the other hand, the results of the KPSS test, calculating the statistics with μ and τ respectively, are reported in Table 3. In the case of weekly nominal returns, the null hypothesis of stationarity is rejected for Hong Kong and Taiwan exchange rate returns at the 1% significance level. However, the monthly nominal exchange rate returns for Hong Kong, New Zealand, Singapore and Taiwan reject the null hypothesis at the 1% significance level. Therefore, the results of

stationary tests for weekly Hong Kong and Taiwan nominal exchange rate returns and monthly Hong Kong, New Zealand, Singapore and Taiwan correspond to the fourth case, implying that they are neither I (0) processes nor I (1) processes. They might be a half way of both I (0) processes and I (1) processes, indicating long memory processes.

Table 3. The results of stationary test for all nominal exchange rate returns.

Country	ADF		PP		KPSS	
	Without Trend	With Trend	Without Trend	With Trend	η_μ	η_τ
(a) Weekly nominal exchange rate returns						
Australia	-17.34***	-17.43***	-41.36***	-41.42***	0.303	0.056
Hong Kong	-16.33***	-16.66***	-36.58***	-36.86***	1.114***	0.273***
Japan	-16.02***	-16.02***	-40.13***	-40.12***	0.089	0.070
Malaysia	-17.03***	-17.05***	-52.08***	-52.10***	0.097	0.033
New Zealand	-17.03***	-17.19***	-40.91***	-41.02***	0.567	0.072
Singapore	-14.86***	-14.86***	-37.58***	-37.57***	0.178	0.146
South Korea	-13.56***	-13.56***	-36.57***	-36.56***	0.104	0.106
Taiwan	-12.65***	-12.76***	-28.86***	-28.94***	1.103***	0.332***
Thailand	-15.01***	-15.00***	-32.29***	-32.28***	0.087	0.086
(b) Monthly nominal exchange rate returns						
Australia	-8.46***	-8.67***	-14.18***	-14.26***	0.412	0.077
Hong Kong	-6.91***	-7.46***	-12.19***	-12.49***	1.543***	0.396***
Japan	-8.46***	-8.46***	-13.39***	-13.38***	0.128	0.080
Malaysia	-7.88***	-7.91***	-14.80***	-14.81***	0.206	0.066
New Zealand	-7.48***	-7.83***	-13.17***	-13.32***	0.781***	0.105
Singapore	-7.92***	-7.94***	-12.29***	-12.27***	0.242	0.203**
South Korea	-6.96***	-6.95***	-9.27***	-9.25***	0.113	0.117
Taiwan	-5.06***	-5.24***	-10.55***	-10.64***	1.022***	0.277***
Thailand	-6.29***	-6.28***	-12.90***	-12.88***	0.084	0.088

Notes: (1) The ADF and PP critical values without trend: -3.43 and -2.86 at the 1% and 5% significance levels, respectively; the ADF and PP critical values with trend: -3.97 and -3.41 at the 1% and 5% significance levels, respectively. All of the estimated statistics above are smaller than the critical value, indicating that the weekly and monthly nominal exchange rate returns are stationary. (2) KPSS critical values are: (a) η_μ is 0.739 and 0.574 at the 1%, and 5% significance levels, respectively; (b) η_τ is 0.216, 0.176 at the 1% and 5% significance levels, respectively. *** indicates significance at 1% level. ** indicates significance at the 5% level.

2. Empirical results

To investigate long-term dependence in nominal exchange rate returns, this study considers two techniques. The first approach, Lo's modified rescaled range (R/S) analysis, is used to minimise short-term dependence which may bias the estimates of long-term dependence in the returns. The second approach is the GPH test which estimates the fractionally integrated parameter (d) in returns. Both techniques test short-term dependence under the null hypothesis against long-term dependence as alternatives.

2.1 Empirical results of the modified R/S analysis

The modified R/S analysis is performed on both weekly and monthly nominal exchange rate returns, and the results are reported in Table 4. That table displays the classical R/S statistic with $q = 0$ and the modified R/S statistic with optimal truncation lag for each nominal exchange rate returns (q). The classical R/S statistics are included for a comparative purpose. The results of classical R/S and modified R/S statistics show slight difference. In the case of classical R/S statistics, there is significant evidence of long-term dependence in two weekly nominal returns (Hong Kong and Taiwan) as well as in three monthly nominal returns (Hong Kong, New Zealand and Taiwan). This finding is consistent with that of the stationary tests.

In the case of modified R/S statistics, only one case (Hong Kong) can reject the null hypothesis of no long-term dependence in both weekly and monthly nominal exchange rate returns. Generally, little evidence of long-term dependence is found in the

Australian, Japanese, Malaysian, New Zealand, Singaporean, South Korean and Thailand currency returns. These findings are not incompatible with that of the classical R/S statistic which might be biased as the estimator of the classical R/S analysis is highly sensitive to short-term dependence. In short, nominal exchange rate returns are supportive of the EMH. This finding contradicts that of Mulligan (2000) who found significant evidence of long-term dependence in the monthly nominal exchange rate returns for Australia, Japan, Malaysia and New Zealand over the period January 1973 to December 1997.

Table 4. Results of modified R/S analysis for all nominal exchange rate returns

Country	Weekly nominal returns			Monthly nominal returns		
	Classical R/S	Modified R/S	q-lag selected	Classical R/S	Modified R/S	q-lag selected
Australia	1.1938	1.1747	8	1.3637	1.1729	5
Hong Kong	2.4191***	2.391***	7	2.4213***	1.9315**	5
Japan	1.3596	1.2314	8	1.4726	1.1407	5
Malaysia	1.1783	1.375	8	1.5432	1.2276	5
New Zealand	1.6467	1.5657	8	1.8822**	1.448	5
Singapore	1.5863	1.5229	7	1.7484	1.5702	5
South Korea	1.632	1.3815	7	1.6097	1.2766	5
Taiwan	2.3285***	1.7725	7	2.1303***	1.4951	5
Thailand	1.8366	1.499	7	1.6562	1.3667	5

Note: 1% significant level is in the interval [0.721, 2.098]
5% significant level is in the interval [0.861, 1.747]
*** indicates significant at 1%
** indicates significant at 5%

2.2 Empirical results of the GPH test

The results of the long-term dependence parameter d for the weekly and monthly nominal exchange rate returns are presented in Table 5. A choice has to be made with respect to the number of low-frequency ordinates, n in the estimation. Inclusion of

medium or high order periodogram ordinates will cause bias in the d estimate. On the other hand, 'a too small value of n will lead to imprecise estimates due to limited degrees of freedom in the regression' (Cheung and Lai 1993b, p.107). In order to ensure the robustness of the GPH test to the choice of the number of low-frequency ordinates, this paper follows Barkoulas et al. (2000) work which allows several choices of low-frequency ordinates. These choices vary with the sample size T and are established in terms of $n = T^{\alpha}$ with α = {0.5, 0.525, 0.55, 0.575, and 0.6}. Table 5 shows the d estimates corresponding to $d(0.50)$, $d(0.525)$, $d(0.55)$, $d(0.575)$ and $d(0.60)$ when sample size $n = T^{0.5}$, $n = T^{0.525}$, $n = T^{0.55}$, $n = T^{0.575}$ and $n = T^{0.6}$ respectively. The d estimates are reported together with their t-statistics.

To test for the statistical significance of the d estimates, the null hypothesis ($H_0 : d = 0$) and the alternative ($H_0 : d \neq 0$) are performed. As is shown in Table 5, the d estimates for weekly and monthly nominal returns have a similar pattern. Significant evidence of long-term dependence can be found in only three (Hong Kong, Malaysia, and Taiwan) cases of weekly nominal returns and two (Hong Kong and Taiwan) cases of monthly returns. The majority of the returns can not reject the null hypothesis of short-term dependence. Thus, there is little evidence of long-term dependence in the nominal exchange rate returns. This finding is consistent with that of Barkoulas et al. (2003) who found no convincing evidence of long-term dependence in the returns for currencies of eighteen industrial countries.

In comparison with the earlier results from the modified R/S analysis, the results of the GPH test provide slightly more favourable evidence for long-term dependence in the weekly and monthly nominal exchange rate returns. Nevertheless, both results suggest that there is the absence of long-term dependence in weekly and monthly nominal exchange rate returns.

Table 5. Estimates of the fractional differencing parameter (d)

Country	$d(0.50)$	$d(0.525)$	$d(0.55)$	$d(0.575)$	$d(0.60)$
(a) Weekly nominal exchange rate returns					
Australia	0.039 (0.20)	-0.012 (-0.07)	-0.020 (-0.13)	-0.016 (-0.11)	0.047 (0.37)
Hong Kong	0.216*** (3.04)	0.184*** (3.05)	0.107 (1.76)	0.096 (1.49)	0.112 (1.85)
Japan	0.192 (0.98)	0.189 (1.13)	0.181 (1.19)	0.072 (0.53)	0.030 (0.25)
Malaysia	0.088 (1.14)	0.090 (1.33)	0.070 (1.14)	0.127** (1.96)	0.104 (1.86)
New Zealand	0.112 (1.04)	0.009 (0.09)	0.049 (0.55)	0.037 (0.46)	0.057 (0.82)
Singapore	0.005 (0.05)	-0.013 (-0.13)	-0.038 (-0.43)	-0.043 (-0.55)	0.027 (0.35)
South Korea	-0.086 (-0.30)	-0.110 (-1.79)	-0.043 (-0.76)	-0.042 (-0.84)	-0.025 (-0.57)
Taiwan	0.347*** (2.83)	0.257** (2.13)	0.291*** (2.74)	0.238*** (2.58)	0.278*** (3.15)
Thailand	0.062 (0.55)	0.097 (0.90)	0.032 (-0.43)	-0.033 (-0.43)	-0.033 (-0.50)
(b) Monthly nominal exchange rate returns					
Australia	0.032 (0.27)	-0.096 (-0.87)	-0.097 (-1.07)	-0.105 (-1.34)	-0.066 (-0.99)
Hong Kong	0.499*** (5.48)	0.513*** (6.25)	0.349*** (3.30)	0.245** (2.32)	0.269*** (2.92)
Japan	-0.021 (-0.20)	0.018 (0.17)	0.056 (0.62)	0.108 (1.29)	0.101 (1.31)
Malaysia	-0.085 (-0.89)	-0.017 (-0.19)	-0.012 (-0.14)	0.420 (0.56)	0.030 (0.43)
New Zealand	0.105 (0.44)	0.161 (0.78)	0.122 (0.68)	0.115 (0.78)	0.100 (0.76)
Singapore	0.275 (1.77)	0.218 (1.67)	0.168 (1.44)	0.079 (0.72)	0.106 (0.74)
South Korea	-0.014 (-0.10)	-0.025 (-0.10)	0.005 (0.06)	-0.022 (-0.25)	-0.024 (-0.30)
Taiwan	0.281 (1.50)	0.248 (1.65)	0.336** (2.51)	0.365*** (3.14)	0.405*** (3.91)
Thailand	-0.187 (-1.58)	-0.166 (-1.71)	-0.104 (-1.15)	-0.068 (-0.79)	-0.039 (-0.47)

Note: Two-sided test $H_0 : d = 0$ and $H_1 : d \neq 0$, Two-sided critical values: 2.576(1%), and 1.960(5%). $d(0.50)$, $d(0.525)$, $d(0.55)$, $d(0.575)$, and $d(0.60)$ give the d estimates corresponding to the GPH spectral regression of sample $v = T^{0.50}, T^{0.525}, T^{0.55}, T^{0.575}$, and $T^{0.60}$ respectively. The t-statistics are given in brackets.

*** Implies rejection of the null of the two-sided test at the 1% level.

** Implies rejection of the null of the two-sided test at the 5% level.

IV. CONCULSION

This study examines long-term dependence in the weekly and monthly nominal exchange rate returns for nine Pacific-Rim countries during the post-Bretton Wood period, using Lo's modified R/S analysis and the GPH test. All nominal returns are not random or independent but correlated with others. Further, significant short-term dependence is presented, making it difficult to describe the nature of long-term dependence in nominal exchange rate returns. Therefore, the modified R/S analysis and the GPH test are used to minimise short-term dependence in this study.

Using stationary tests and the classical R/S analysis, significant evidence of long-term dependence is found in the nominal exchange rate returns. Unlike the results of conventional techniques (stationary tests and the classical R/S analysis), the results of modified R/S statistics can not reject the null hypothesis of short-term dependence in the weekly and monthly nominal returns except for Hong Kong. In addition to modified R/S statistics, the results of the GPH test suggest that there is significant evidence of long-term dependence in the two weekly nominal returns (Hong Kong, Malaysia and Taiwan) and the three monthly nominal returns (Hong Kong and Taiwan). However, the majority of nominal returns show little evidence of long-term dependence. Thus, the presence of long-term dependence is supportive of market efficiency.

This finding is not consistent with previous literature, which found significant evidence of long-term dependence in the major foreign currency markets (Booth et al. 1982; Cheung 1993; Mulligan 2000). There are two possible explanations for different agreements with previous literature. First, this study uses a longer, broader and more recent sample. Second, Pacific-rim exchange rate markets have recently been

liberalised since Asian currency crisis. In contrast, this finding is advocated by Barkoulas et al. (2003) who found little evidence of long-term dependence in eighteen currency returns. Therefore, the findings of long-term dependence in foreign exchange markets have been inclusive.

REFERENCES

Agiaklogou, C., P. Newbold, and M. Wohar, "Bias in an Estimator of the Fractional Difference Parameter," *Journal of Time Series Analysis*, 14, 1992, pp. 235-246.

Baillie, R.T. and T. Bollerslev, "Common Stochastic Trend in a System of Exchange rates," *Journal of Finance*, 44(1), 1989, pp. 167-181.

Barkoulas, J., C. Baum, M. Caglayan, and A. Chakraborty, "Persistent dependence in foreign exchange rates? A reexamination," *Forthcoming in global financial markets: Issues and policies*. Westport, CT: Greenwood Press, 2003.

Barkoulas, J., W.C. Labys, and J. Onochie, "Fractional Dynamics in International Commodity Prices," *Journal of Futures Markets*, 17(2), 1997, pp. 161-189.

Booth, G.G., F.R. Kaen, and P.E. Koveos, "R/S Analysis of Foreign Exchange Rates Under Two International Monetary Regimes," *Journal of Monetary Economics*, 17(1), 1982, pp. 407-415.

Cheung, Y.W., "Long Memory in Foreign-Exchange Rates," *Journal of Business & Economic Statistics*, 11(1), 1993, pp. 93-101.

Cheung, Y.W. and K.S. Lai, "A Search for Long Memory in International Stock Market Returns," *Journal of International Money and Finance*, 14(4), 1995, pp. 597-615.

Geweke, J. and S. Porter-Hudak, "The Estimation and Application of Long Memory Time Series Models," *Journal of Time Series Analysis*, 4, 1983, pp. 221-238.

Granger, C.W.J. and R. Joyeux, "An Introduction to Long Memory Time Series Models and Fractional Differencing," *Journal of Time Series Analysis*, 1, 1980, pp. 5-39.

Hakkio, C.S., "Does the Exchange Rates Follow a Random Walk? A Monte Carlo Study of Four Tests for a Random Walk," *Journal of International Money and Finance*, 5, 1986, pp. 221-229.

Hosking, J.R.M., "Fractional Differencing," *Biometrika*, 68(1), 1981, pp. 165-176.

Hurst, H.E., "Long-Term Storage Capacity of Reservoirs," *Transactions of the American Society of Civil Engineers*, 116, 1951, pp. 770-799.

Lee, D. and P. Schmidt, "On the Power the KPSS Test of Stationarity against Fractionally-Integrated Alternatives," *Journal of Econometrics*, 73, 1996, pp. 285-302.

Lo, A.W., "Long-Term Memory in Stock Market Prices," *Econometrica*, 59, 1991, pp. 1279-1314.

Mulligan, R.F. and G.A. Lombardo, "A Fractal Analysis of Foreign Exchange Markets," *International Advances in Economic Research*, 6(1), 2000, pp. 33-50.

Robinson, P., "Gaussian Semiparametric Estimation of Long Range Dependence," *Annals of Statistics*, 13, 1995, pp. 1630-1661.

Pan, M.S., L.Y. Angela, and H. Bastin, "An Examination of the Short-Term and Long-Term Behavior of Foreign Exchange Rates," *The Financial Review*, 31(3), 1996, pp. 603-622.

Skjeltorp, J.A., "Scaling in the Norwegian Stock Market," *Physica A*, 283, 2000, pp. 486-528.

Sowell, F., "Maximum Likelihood Estimation of Stationary Univariate Fractionally Integrated Time Series Models," *Journal of Econometrics*, 53, 1992, pp. 165-188.

Willinger, W., M.S. Taqqu, and V. Teverovsky, "Stock Market Prices and Long-Range Dependence," *Finance and Stochastics*, 3, 1999, pp. 1-13.