

LONG-TERM MEMORY IN STOCK MARKET RETURNS: INTERNATIONAL EVIDENCE

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ABSTRACT

A lot of recent work has addressed the issue of the presence of long memory components in stock prices because of the controversial implications of such a finding for market efficiency and for martingale models of asset prices used in financial economics and technical trading rules used for forecasting. This paper examines the presence of long memory in the stock returns of seven countries, namely Japan, Korea, New Zealand, Malaysia, Singapore, the USA and Australia. The classical and modified rescaled range tests, the semiparametric test proposed by Geweke and Porter-Hudak, the frequency domain score test proposed by Robinson and its time-domain counterpart derived by Silvapulle, are applied to these returns in order to detect the long memory property. Evidence suggests that the Korean, Malaysian, Singapore and New Zealand stock returns are long-term dependent, indicating that these two markets are not efficient. The results of this study should be useful to regulators, practitioners and derivative market participants, whose success precariously depends on the ability to forecast stock price movements. Copyright © 2001 John Wiley & Sons, Ltd.

KEY WORDS: efficient markets; long memory; stock returns

1. INTRODUCTION

The efficient market, random walk model of stock prices requires that the arrival of new information be promptly arbitrated away. A necessary and sufficient condition for the existence of arbitrage price is that statistical dependence between distance observations of a price series must decrease very rapidly. If persistent statistical dependence is present, the arbitrage price changes do not follow the martingale process, which characterizes the efficient market, and should have an infinite variance. Therefore, the question of whether or not financial markets are efficient is directly related to whether or not long-term dependence is present in returns. Contrary to what random walk hypothesis suggests, several studies report positive autocorrelation for stock returns in the short-run and negative autocorrelation in the long-run, for example, see Lo and MacKinlay (1988), Poterba and Summers (1988) and Cutler *et al.* (1990). Negative correlation means that stock prices are mean-reverting, indicating the presence of long-term dependence and, if true, this will have important implications for modern financial economics. For example, optimal consumption/savings and portfolio decisions may become sensitive to the investment horizon. Problems also arise in the pricing of derivatives such as options and futures with martingale methods, since the class of continuous time stochastic processes most commonly employed is inconsistent with long-term memory. Traditional tests of the capital asset pricing model and arbitrage pricing theory are no longer valid since the usual forms of statistical inference do not apply to time series exhibiting such persistent statistical dependence. Moreover, the conclusions of tests of efficient market hypotheses or stock market rationality also depend precariously on the presence or absence of long-term memory in the stock return series.

The objective of this paper is to test for the presence of long memory in the stock returns of seven countries, namely Japan, Korea, New Zealand, Malaysia, Singapore, the USA and Australia. Several studies have used various procedures for testing for long-term dependence in composite and common stock returns. The conclusions of these studies are mixed depending on the testing procedure, sample

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period, frequencies of the series, composite stock returns or common stock returns and others used in these studies.

There are several procedures available in the literature for testing for the presence of long memory in time series. The primary aim of this paper is to use a number of testing procedures and, according to Banik and Silvapulle (1998), some of these tests have varying size and power properties indicating the reliability of these tests. Further, having found the presence of long memory, measuring the size of the long memory parameter is also an important issue. There are a number of estimation procedures developed over the years; application of these procedures to measure the size of the long memory in stock returns is a topic for future research.

Diebold and Rudebusch (1989) and Shea (1991), for example, have found that economic fundamentals have long memory property using the procedure proposed by Geweke and Porter-Hudak (1983; GPH hereafter). Since the stock market is one of the best leading economic indicators, considerable attention has been given to testing for the presence of long memory in stock returns. For example, see Greene and Fielitz (1977), Kaen and Rosenman (1986), Lo (1991), Ding *et al.* (1993), Mills (1993), Lee and Robinson (1996), Hiemstra and Jones (1997) and others for testing for long memory in stock returns. Also, see Cheung (1993) for testing for long memory property in nominal exchange rates, and the recent survey by Baillie (1996). Since the work of Baillie *et al.* (1996) on testing for generalized conditional heteroscedasticity (GARCH), several studies have investigated the presence of long memory in variance processes. For example, Hiemstra and Jones (1997) found that the mean returns of a number of US common stocks do not possess long-term dependence, whereas the squared returns do. This issue will be investigated in future research.

One common component of technical trading rules is the moving average rule, which has been frequently used for forecasting and recommending buy and hold strategies by technical trading analysts with the objective of maximizing profit and minimizing risk of loss. Technical trading analysis is based on an assumption that the market's behaviour patterns do not change over time, particularly the long-term trend. While future events can be very different from any past events, the way market participants respond to new information or uncertainties is usually similar to the way they handled them in the past. The patterns in the past security price series are assumed to recur in the future, and thus these patterns can be used for predictive purposes. If the long memory component is present in security returns, it is very useful to recommend a high-order moving average as a trading rule. On the other hand, if only the short memory component is present, a low-order moving average rule can be recommended. Further, it has been argued that an abnormal profit can be made if the long memory is present in the stock return, see Kurz (1990), Vaga (1990) and Lux (1995) for details.

One statistical method extensively used in the early literature for detecting the presence of long-term dependence in time series is called 'rescaled range' or R/S analysis, proposed by Mandelbrot (1972). A study by Aydogan and Booth (1988) has revealed that the R/S test depends, to some extent, on short-term dependence and nonhomogeneity in the data series. To overcome these problems, Lo (1991) developed a modified R/S test that is robust to both weak dependence and heteroscedasticity. GPH have used the low-frequency periodogram and proposed a regression-based test for the long memory property. Recently, Robinson (1994) developed an efficient frequency domain score test for testing long memory in a wide class of models, whereas Silvapulle (2000) has derived its time-domain counterpart. In the recent past, many studies have used R/S and GPH procedures to detect long memory in stock returns. In this study we include these two tests and the score tests proposed by Robinson (1994) and Silvapulle (2000). Since these procedures, except the modified R/S test, are very sensitive to the presence of short-term dependence and dynamic heteroscedasticity, we have applied these tests to original, autoregressive moving average (ARMA)-filtered and ARMA-ARCH-filtered returns.

This paper is organized as follows: the next section briefly discusses the long-term dependence of a time series and its properties. The third section outlines the testing procedures. The fourth section discusses the data series used in this study, and reports and analyses the empirical results of the tests. The last section concludes the paper.

2. FRACTIONAL STATISTICAL ANALYSIS

Since the work of Mandelbrot (1972) on fractional processes, the interesting properties of these processes have been investigated in the literature, for example, see Granger and Joyeux (1980), Hosking (1981, 1982), Sowell (1990) and others. These studies have mainly examined the members of a family of fractional processes known as fractionally differenced processes. Because of their flexibility and simplicity in their specification, these fractional processes are useful for modelling low-frequency dynamics.

A general class of fractional processes ARFIMA(p, d, q) is described as

$$\Phi(B)(1 - B)^d x_t = \Theta(B)e_t \quad (1)$$

where $\{x_1, \dots, x_T\}$ is a set of time series data, $\Phi(B) = 1 - \varphi_1 B - \dots - \varphi_p B^p$ and $\Theta(B) = 1 + \theta_1 B + \dots + \theta_q B^q$ are polynomials in the lag operator B with all roots of $\Phi(B)$ and $\Theta(B)$ being stable, and e_t is a white noise term that will be relaxed in the data analysis. When $p = q = 0$, x becomes a simple fractional noise process. The fractional parameter d can take noninteger values. The variance of the process is finite when $d < 1/2$, but infinite when $d \geq 1/2$. The process is stationary for $d < 1/2$ and invertible for $d > -1/2$. Long memory is associated with $1/2 > d > 0$. If $d < 0$, the process is said to exhibit intermediate memory.

Approximating $(1 - B)^d$ expansion, one obtains

$$(1 - B)^d = \sum_{k=0}^{\infty} \Gamma(k - d) B^k / \Gamma(k + 1) \Gamma(-d) \quad (2)$$

where $\Gamma(\cdot)$ is the standard gamma function. In both cases the autocorrelation function obeys a hyperbolic pattern of decay contrary to the exponential decline for the commonly used stationary ARMA process (1) with $d = 0$. An extension to have noninteger values of d increases the flexibility in modelling long-term dynamics by allowing for a richer class of spectral behaviour at low frequencies than that implied by ARMA models. This can be seen from the spectral density of x_t . Granger and Joyeux (1980) and Hosking (1981) show that the spectral density function of x_t , denoted by $f_x(\lambda)$ and defined in Equation (3) below is proportional to λ^{-2d} as $\lambda \rightarrow 0^+$. Thus, the fractional parameter d crucially determines the low-frequency dynamics of the process. For $d > 0$, $f_x(\lambda)$ is unbounded at frequency $\lambda = 0$, but bounded for ARMA processes with $d = 0$. It becomes clear that when $d > 0$, x_t is a long memory process showing a slow autocorrelation decay at a hyperbolic rate (Hosking, 1981), which is contrary to an exponential decay of a short memory process. Note that the larger the value of d , the stronger the long-term dependence will be, see Rosenbaltt (1956) for details.

The general form of the spectral density of x_t in Equation (1) is given by

$$f_x(\lambda) = |1 - \exp(-i\lambda)|^{-2d} g_u(\lambda) \quad (3)$$

where $u_t = \Phi^{-1}(B)\Theta(B)e_t$ is a stationary process and $g_u(\lambda)$ is its spectral density at frequency λ . The spectral density $g_u(\lambda)$ is defined as

$$g_u(\lambda) = (\sigma^2/2\pi) \frac{|\Theta e^{-2i\lambda}|^2}{|\Phi e^{-2i\lambda}|^2}$$

3. TESTING PROCEDURES

In this section, (i) the classical and modified rescaled range R/S test statistics, (ii) the GPH procedure and (iii) frequency and time-domain versions of the score statistic are outlined.

3.1. Rescaled Range Analysis

One approach to detect evidence of strong dependence in time series x_t is to use the 'range over standard deviation' or 'rescaled range' statistic, originally developed by Hurst (1951) and popularized by

Mandelbrot (1972) in the economic context. This R/S statistic is the range of partial sums of deviations of a time series from its mean, rescaled by its standard deviation. For $\{x_1, \dots, x_T\}$, it may be defined as

$$Q_n = S_n^{-1} \left[\max_{1 \leq k \leq n} \sum_{j=1}^k (x_j - \bar{x}_n) - \min_{1 \leq k \leq n} \sum_{j=1}^k (x_j - \bar{x}_n) \right] \quad (4)$$

where $\bar{x}_n = (1/n) \sum_{j=1}^n x_j$ and $S_n^2 = (1/n) \sum (x_i - \bar{x}_n)^2$ are the sample mean and variance, respectively.

The first term in brackets in Equation (4) is the maximum (over k) of the partial sums of the first k deviations of x_i from the sample mean. Since the sum of all T deviations of the x_i from their mean is zero, this maximum is always non-negative. The second term is the minimum (over k) of the same sequence of partial sums; hence it is always nonpositive. Therefore, the difference between the two quantities, known as the 'range', is always non-negative and hence $Q_n \geq 0$. Although it has long been established that the Q_n statistic has the ability to detect long-range dependence, later studies have found that it is sensitive to short-range dependence. Therefore, Lo (1991) considers a modified R/S statistic in which short-run dependence is incorporated into its denomination, which becomes a consistent estimate of the variance of the partial sum in Equation (4), defined as

$$Q_{n,q} = S_q^{-1} \left[\max_{1 \leq k \leq T} \sum_{j=1}^k (x_j - \bar{x}_n) - \min_{1 \leq k \leq T} \sum_{j=1}^k (x_j - \bar{x}_n) \right] \quad (5)$$

where

$$S_q^2 = S_n^2 + 2T \sum_{j=1}^q w_{qj} r_j, \quad w_{qj} = 1 - \frac{j}{q+1}, \quad q < T$$

r_j is the j th sample autocorrelation, $T^{-1} \sum_{j+1} (x_t - \bar{x})(x_{t-j} - \bar{x}_n) S_n^{-1}$. Lo (1991) provides the assumptions and technical details to allow the asymptotic distribution of $Q_{n,q}$ to be obtained. Under the null hypothesis that $\{x_1, \dots, x_T\}$ is short-range dependent, $V - T^{-1/2} Q_{n,q}$ converges in distribution to a well-defined random variable, whose distribution function is given. Its associated critical values at various significance levels have been reported in Lo (1991). The V statistic is consistent against a class of long-range dependent alternatives that, for example, includes all ARFIMA models with $|d| \leq 1/2$. In the presence of positive strong dependence $0 \leq d \leq 1/2$, V diverges in probability to infinity, and in the presence of negative strong dependency $-1/2 \leq d \leq 0$, it converges in probability zero.

3.2. The GPH Test Statistic

We use a spectral procedure developed by GPH (1983) to estimate and test the fractional parameter d in Equation (1). The GPH test makes use of the fact that the low frequency dynamics of a process are parameterized by the fractional parameter. The spectral density function $f_x(\lambda)$ defined in Equation (3) is equivalent to

$$f_x(\lambda) = (2 \sin(\lambda/2))^{-2d} g_u(\lambda) \quad (6)$$

Taking the logarithm of Equation (6) and evaluating at harmonic frequencies $\lambda_j = 2\pi j/T$, $j = 1, \dots, T-1$, we obtain

$$\ell n(f_x(\lambda_j)) = \ell n(g_u(0)) - d \ell n(4 \sin^2(\lambda_j/2)) + \ell n(g_u(\lambda_j)/g_u(0)) \quad (7)$$

For low frequency ordinates λ_j near zero, the last term is negligible compared with the other terms. Adding $\log(I(\lambda_j))$ —the periodogram at ordinate j —to both sides of Equation (7) yields

$$\ell n(I(\lambda_j)) = \ell n(g_u(0)) - d \ell n(4 \sin^2(\lambda_j/2)) + \ell n(I(\lambda_j)/f_x(\lambda_j))$$

This suggests that d can be estimated using a simple regression equation

$$\ell n(I(\lambda_j)) = \alpha_0 - \alpha_1 (4 \sin^2(\lambda_j/2)) + \varepsilon_t \quad (8)$$

where $\varepsilon_t = \ell n(I(\lambda_j)/f_x(\lambda_j))$ is asymptotically independent identically distributed (i.i.d.) across harmonic frequencies. The periodogram $I(\lambda_j)$ is compared as $(2\pi n)^{-1} |\sum_{t=1}^T x_t e^{it\lambda_j}|^2$. The number of low frequency ordinates n used for the regression is an increasing function of T . It is shown in GPH (1983) that for $n = T^\alpha$ with $0 < \alpha < 1$, the least squares estimate of α_1 in Equation (8) provides a consistent estimate of d and testing the null hypothesis that $d=0$ can be based on the t -statistic of the slope coefficient. The theoretical asymptotic variance of ε_t is known to be $\pi^2/6$, which can be imposed in estimation to raise efficiency.

3.3. Frequency Domain Score Test

In this section, we briefly outline the score test statistic for testing the null hypothesis that $d = 0$ against $d \neq 0$, proposed by Robinson (1994). The test statistic is defined as

$$S_{\text{freq}} = (\lambda/\tilde{\sigma}^4)\tilde{a}'\tilde{A}^{-1}\tilde{a} \tag{9}$$

where $\tilde{a} = -(2\pi/\lambda) \sum_j \psi(\lambda_j)I_x(\lambda_j)$ and $\tilde{A} = (2/n) \sum_{j=1} \psi(\lambda_j)\psi(\lambda_j)'$,

$$\psi(\lambda_j) = \log \left| 2 \sin \left(\frac{1}{2} \lambda_j \right) \right|,$$

$\lambda_j = 2\pi j/T$ and primed sums are over $\lambda_j \in M$, where

$$M = \{ \lambda : -\pi < \lambda < \pi, \lambda \notin (\rho_\ell - \lambda_\ell, \rho_\ell + \lambda_\ell), \ell = 1, \dots, s \},$$

such that $\rho_\ell, \ell = 1, \dots, s < \infty$ are the distinct poles of $\psi(\lambda)$ on $(-\pi, \pi)$ and $\tilde{\sigma}^4$ is the square of the estimated variance of u_t in Equation (3). The S_{freq} statistic has a chi-squared distribution with one degree of freedom; see Robinson (1994) for more details.

3.4. Time Domain-Score Test

The time-domain version of the score tests developed by Robinson (1994) is derived by Silvapulle (2000). It is briefly defined as follows:

$$S_{\text{tim}} = \left(\frac{(\partial L(\cdot)/\partial d)^2}{(T-m) \sum_{j=1}^m j^{-2}} \right) \tag{10}$$

where $L(\cdot)$ is the log likelihood function of $\{x_1, \dots, x_T\}$, m is the number of terms chosen from the expression of $\log(1 - B)$. The S_{tim} statistic has a chi-squared distribution with one degree of freedom; see Silvapulle (2000) for details.

In what follows we shall apply all four statistics defined in this section to stock returns of all seven countries studied in this paper. Since there is evidence of the presence of short-term dependence and dynamic heteroscedasticity, we have applied these tests to the original returns, ARMA-filtered returns and ARMA-GARCH-filtered returns in all cases.

4. DATA SERIES AND EMPIRICAL ANALYSIS

We have used the weekly aggregate stock price series of seven countries, namely, Japan, Korea, New Zealand, Malaysia, Singapore, the USA and Australia. The series spans the period January 1983–December 1998, consisting of 1040 observations. In a preliminary analysis, we find that all price series are nonstationary, $I(1)$, and that short-term dependence and dynamic heteroscedasticity are present in all seven series.

We applied all four statistics, namely (i) the classical R/S test and its modified versions, (ii) the semiparametric GPH test, (iii) the frequency domain score test and (iv) the time-domain score test, to

seven return series for testing for the presence fractional integration—long memory property—in the mean of the return series. Under H_0 that the fractional parameter is zero, we estimated the order of autoregressive and moving average processes using the sample autocorrelation and partial autocorrelation functions. Note that the serial correlation tests are biased in the presence of ARCH/GARCH and therefore we used the ARCH-adjusted tests proposed by Silvapulle and Evans (1997). Since mounting evidence (e.g. see Akgiray, 1989) suggests the presence of dynamic heteroscedasticity in stock returns we have used Engle (1982) and Silvapulle and Silvapulle (1995) procedures for testing for the presence of ARCH-type dependence. Various information criteria were used to estimate the values of p and q of ARMA and the order of GARCH processes.

Having estimated the order of ARMA and GARCH processes, we filtered the series appropriately and then applied the tests for the presence of long memory. The classical R/S test and its modified series are applied to the original data series, ARMA-filtered series and ARMA-ARCH-filtered series. The results are reported in Table 1. Taking all the results in this table together, there is strong evidence that only the Korean and the New Zealand series have long-term dependence in the mean, and weak evidence (only in Panel C, Table 1) of the presence of long memory in the Malaysian and Singapore stock markets, suggesting that these four markets are inefficient.

The GPH procedure is used to estimate the fractional parameter d in the general model (1) with and without an assumption that e_t follows an ARCH-type process. The d parameter is estimated in three different specifications of (1): (i) $(1 - B)^d y_t = e_t$, (ii) $\Theta(B)(1 - B)^d = \Theta(B)e_t$, where e_t is a white noise, and (iii) the model in (ii) with e_t following a GARCH process. Table 2 exhibits the estimated parameters of d for three ordinates n by setting $\alpha = 0.4, 0.5, 0.55$, where α is defined in the section 'The GPH Test Statistic', and corresponding p -values. This test also identifies that the Korean, New Zealand and Singaporean stock returns possess the long memory property with weak evidence for the Malaysian market.

The efficient score test statistic of Robinson (1994) based on frequency domain is applied to all series. Table 3 reveals the results. This test's results indicate that almost all the series possess the long memory property. Given that this score test over-rejects the null far more frequently under some conditions (see Banik and Silvapulle, 1998), it is difficult to draw any conclusion from the results, as they are unreliable. An application of the time-domain score test derived by Silvapulle (2000), however, indicates that the same four series possess the long memory property. We note that the time domain score tests lose power when the returns are ARMA-ARCH filtered; none of the test statistics for ARMA-ARCH-filtered returns is significant at the 5% level (see Table 4).

In summary, application of the four testing procedures discussed in this paper provides convincing evidence the Korean, the New Zealand, Malaysian and Singaporean data series possess the long memory property. Possible implications of these observations are the following: (i) this evidence is not supportive of efficient market hypothesis in the four countries' financial markets; (ii) these markets are not well diversified; and (iii) the predictive power can be improved by using the presence of long memory. Further research is required to examine the source of inefficiency in these four markets.

5. CONCLUSION

This paper has examined the presence of the long memory property in weekly stock returns of seven countries, namely Japan, Korea, New Zealand, Malaysia, Singapore, the USA and Australia, and has found the mean returns of Korea, Malaysia, Singapore and New Zealand have slowly decaying components. We note that measuring the strength of the memory is also an important issue. A number of estimation procedures for measuring the size of the long memory parameter are proposed in the literature; the next project will utilize these procedures to estimate the size of d in all stock returns studied in the paper, and will compare the reliability of the estimates.

The contribution of this paper is to provide additional insight into the debate by focusing on the time series behaviour of the weekly aggregate returns of seven countries. Previous studies have used the

Table 1. Rescaled range tests results of weekly stock return index

| Data series | \tilde{V}_n | | $\tilde{V}_n(q)$ | | $q=0$ | | $q=4$ | | $q=6$ | | $q=10$ | | $q=20$ | |
|--|---------------|-------|------------------|-------|-------|--------|-------|--------|--------|--------|--------|--------|--------|--|
| | $q=0$ | $q=4$ | Bias | $q=6$ | Bias | $q=10$ | Bias | $q=20$ | Bias | $q=10$ | Bias | $q=20$ | Bias | |
| A. Results of the tests applied to original returns | | | | | | | | | | | | | | |
| Japan | 1.61 | 1.51 | 5.10 | 1.57 | 5.92 | 1.53 | 8.44 | 1.56 | 6.16 | | | | | |
| Korea | 1.81* | 1.51 | 1.39 | 1.79* | 0.94 | 1.75 | 3.59 | 1.69 | 7.13 | | | | | |
| NZ | 2.44* | 1.78* | 31.12 | 1.77* | 37.90 | 1.67 | 45.97 | 1.59 | 53.61 | | | | | |
| Malay | 1.57 | 1.32 | 18.98 | 1.28 | 22.56 | 1.23 | 27.60 | 1.23 | 28.08 | | | | | |
| Sing | 1.24 | 1.07 | 16.49 | 1.03 | 20.53 | 0.977 | 27.07 | 0.99 | 25.02 | | | | | |
| USA | 1.31 | 1.29 | 1.90 | 1.25 | 1.10 | 1.27 | 2.95 | 1.29 | 1.78 | | | | | |
| Aust | 1.34 | 1.38 | 19.20 | 1.11 | 20.91 | 1.13 | 19.12 | 1.18 | 13.32 | | | | | |
| B. Results of the test applied to ARCH-filtered returns | | | | | | | | | | | | | | |
| Japan | 1.62 | 1.56 | 3.29 | 1.55 | 4.040 | 1.52 | 6.50 | 1.54 | 4.77 | | | | | |
| Korea | 1.95* | 1.81* | 7.64 | 1.81* | 7.66 | 2.05* | 10.83 | 1.70 | 14.96 | | | | | |
| NZ | 2.08* | 1.82* | 14.4 | 1.74 | 19.14 | 1.66 | 25.35 | 1.59 | 31.14 | | | | | |
| Malay | 1.38 | 1.29 | 6.85 | 1.27 | 9.14 | 1.22 | 13.01 | 1.23 | 12.86 | | | | | |
| Sing | 1.06 | 1.04 | 2.06 | 1.02 | 4.36 | 0.97 | 9.40 | 0.99 | 6.92 | | | | | |
| USA | 1.35 | 1.29 | 4.41 | 1.30 | 3.81 | 1.27 | 5.82 | 1.29 | 4.73 | | | | | |
| Aust | 1.44 | 1.36 | 5.91 | 1.35 | 6.65 | 1.37 | 4.74 | 1.45 | -0.765 | | | | | |
| C. Results of the tests applied to ARMA-filtered returns | | | | | | | | | | | | | | |
| Japan | 1.68 | 1.57 | 7.26 | 1.55 | 8.40 | 1.51 | 11.34 | 1.530 | 10.04 | | | | | |
| Korea | 1.90* | 1.81* | 5.29 | 1.91* | 5.07 | 2.16* | 8.04 | 1.697 | 11.99 | | | | | |
| NZ | 2.05 | 1.81* | 12.90 | 1.74 | 17.55 | 1.66 | 23.39 | 1.585 | 29.01 | | | | | |
| Malay | 1.34 | 1.29 | 3.75 | 1.27 | 5.51 | 1.23 | 8.83 | 1.235 | 8.37 | | | | | |
| Sing | 0.96 | 1.02 | -6.47 | 1.01 | -5.48 | 0.972 | -1.69 | 1.004 | -4.78 | | | | | |
| USA | 1.33 | 1.29 | 2.86 | 1.30 | 2.06 | 1.28 | 3.82 | 1.297 | 2.39 | | | | | |
| Aust | 1.55 | 1.43 | 8.61 | 1.42 | 9.59 | 1.44 | 7.77 | 1.516 | 2.32 | | | | | |
| D. Results of the test applied to ARMA-ARCH-filtered returns | | | | | | | | | | | | | | |
| Japan | 1.42 | 1.44 | -1.96 | 1.45 | -2.63 | 1.43 | -1.30 | 1.47 | -3.96 | | | | | |
| Korea | 1.68 | 1.78° | 2.69 | 1.74 | -3.45 | 1.79* | -1.57 | 1.66 | 1.26 | | | | | |
| NZ | 1.63 | 1.76° | 2.09 | 1.57 | 3.66 | 1.54 | 6.10 | 1.50 | 8.79 | | | | | |
| Malay | 1.38 | 1.43 | -3.25 | 1.78* | -3.40 | 1.41 | -1.63 | 1.43 | -3.20 | | | | | |
| Sing | 1.18 | 1.22 | -3.68 | 1.20 | -2.01 | 1.76* | 2.52 | 1.77* | 0.08 | | | | | |
| USA | 1.37 | 1.32 | 3.55 | 1.33 | 3.06 | 1.31 | 4.83 | 1.32 | 3.68 | | | | | |
| Aust | 1.03 | 1.02 | 0.88 | 1.04 | -0.88 | 1.08 | -4.69 | 1.17 | -11.67 | | | | | |

\tilde{V}_n and $\tilde{V}_n(q)$ denote the classical and modified rescaled range statistics respectively. * Indicates significance at the 5% level. The 5% critical value is 1.75 (see Lo, 1991, p. 1288 for details).

The % bias is computed using the formula $[(\tilde{V}_n/\tilde{V}_n(q)) - 1] * 100$, indicating the bias of the classical rescaled range statistic in the presence of short-term dependence.

aggregate returns of the USA, the UK and Japan. We extend these studies by including four other countries, namely Korea, Malaysia, Singapore and New Zealand. The aim is to study whether the same results hold across all countries with different financial and economic structures and varying market sizes. Further, previous studies mainly applied the modified R/S test of Lo (1991) and the GPH procedure for testing for the presence of long memory. In addition to these two tests, we include frequency domain and time-domain versions of the score test. We have strong evidence to suggest that Korea, Malaysian, Singapore and New Zealand aggregate stock returns possess the long memory property. The evidence is not supportive of efficient market hypothesis; this result is normally interpreted to cast doubt on the ability of these markets to channel funds into the most productive sectors of the economy. It also indicates that these four markets are not well diversified. However, further research is required to examine the source of inefficiency in these markets. The results of this study should be useful to regulators, practitioners and derivative market participants whose success precariously depends on the ability to predict stock price movements.

Table 2. Calculated t -statistics of long memory parameter d in Equation (5) using the GPH procedure

| Data series | Original series | | | ARMA-filtered series | | | ARMA-ARCH-filtered series | | | |
|-------------|-----------------|-------|-------|----------------------|-------|--------|---------------------------|--------|--------|--------|
| | α | 0.55 | 0.50 | 0.45 | 0.55 | 0.50 | 0.45 | 0.55 | 0.50 | 0.45 |
| Japan | | -0.36 | -0.09 | 1.11 | 0.39 | -0.14 | 0.87 | 0.78 | -0.08 | 0.25 |
| Korea | | 0.64 | 1.01 | 1.05 | -1.62 | -2.09* | -1.83* | -1.98* | -1.62 | -2.34* |
| NZ | | 1.12 | 0.08 | 0.90 | -1.60 | 1.57 | -1.85* | -1.67 | -1.98* | -1.58 |
| Malay | | -0.36 | -0.09 | 1.08 | -2.89 | -3.80* | -3.89* | -1.91 | -3.00* | -1.58 |
| Sing | | -0.43 | -0.41 | -1.92* | -1.56 | -2.99* | -2.10* | -1.61 | -2.73* | -2.88* |
| USA | | 0.37 | 0.38 | 0.07 | 0.32 | -0.36 | 0.78 | 0.73 | 0.07 | 1.09 |
| Aust | | -0.39 | -0.45 | 0.08 | -0.44 | -0.99 | 0.32 | 0.09 | 0.11 | 0.45 |

The reported test statistics were computed assuming the error variance $\pi^2/6$.

α is the root of the sample size, used to compute the number of periodogram ordinates, which in turn were used in the GPH regression. $H_0: d = 0$ is tested against $H_1: d < 0$.

* Indicates significance at the 5% level.

Table 3. Results of the frequency domain score test statistics

| Data series | Original series | | | ARMA-filtered series | | | ARMA-ARCH-filtered series | | | |
|-------------|-----------------|-------|-------|----------------------|--------|--------|---------------------------|-------|-------|-------|
| | m | 5 | 10 | 15 | 5 | 10 | 15 | 5 | 10 | 15 |
| Japan | | 1.99 | 2.58 | 3.11 | 10.41* | 10.48* | 11.32* | 1.67 | 1.72 | 1.88 |
| Korea | | 3.74 | 3.92 | 3.99 | 19.89* | 19.85* | 19.85* | 3.70 | 3.98* | 4.25* |
| NZ | | 4.10* | 4.12* | 4.18* | 8.37* | 8.39* | 9.56* | 8.28* | 8.29* | 9.21* |
| Malay | | 1.43 | 1.89 | 2.51 | 6.08* | 6.19* | 6.21* | 0.56 | 1.20 | 1.48 |
| Sing | | 2.57 | 2.59 | 2.92 | 4.83 | 4.89 | 4.90 | 0.02 | 0.99 | 1.58 |
| USA | | 6.09* | 6.10* | 6.19* | 7.08* | 7.81* | 8.02* | 6.34* | 6.38* | 7.01* |
| Aust | | 1.19 | 1.72 | 1.77 | 3.93 | 3.99* | 4.56* | 3.90* | 3.92* | 4.08* |

* Indicates significance at the 5% level.

Table 4. Results of the time-domain score test statistics

| Data series | Original series | | | ARMA-filtered series | | | ARMA-ARCH-filtered series | | | |
|-------------|-----------------|------|------|----------------------|-------|-------|---------------------------|------|------|------|
| | m | 5 | 10 | 15 | 5 | 10 | 15 | 5 | 10 | 15 |
| Japan | | 0.02 | 0.04 | 0.00 | 2.64 | 2.71 | 2.06 | 0.13 | 0.05 | 0.16 |
| Korea | | 2.34 | 3.40 | 7.65* | 5.60* | 5.44* | 5.42* | 2.86 | 2.23 | 2.23 |
| NZ | | 4.01 | 4.15 | 4.20 | 8.11* | 8.31* | 8.48* | 2.83 | 3.38 | 2.49 |
| Malay | | 1.05 | 2.29 | 2.02 | 3.66 | 4.23 | 4.99 | 0.58 | 0.79 | 0.70 |
| Sing | | 3.45 | 3.60 | 3.88* | 3.26 | 4.09* | 4.38* | 1.59 | 2.28 | 2.54 |
| USA | | 0.96 | 1.36 | 1.29 | 0.21 | 0.47 | 0.36 | 0.55 | 0.76 | 0.60 |
| Aust | | 2.78 | 2.09 | 2.03 | 3.57 | 2.12 | 2.06 | 1.70 | 1.51 | 1.33 |

* Indicates significance at the 5% level.

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