Testing for covariance stationarity of stock returns in the presence of structural breaks: an intervention analysis
Ada K. F. Ho; Alan T. K. Wan

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This paper investigates whether the stock return series of Australia, Hong Kong, Singapore and the US are covariance stationary using Omran and McKenzie’s (*The Statistician*, 48, 361–69, 1999) testing procedure which comprises the Loretan and Phillips (1994) test and an intervention analysis. The main objective of the procedure is to ascertain the role of structural breaks on the stochastic properties of the stock return series. It is found that the intervention due to the 1997 Asian financial crisis is significant in the case of Hong Kong and Singapore, for which the hypothesis of covariance stationarity cannot be rejected after the effects of the financial crisis have been properly filtered. On the other hand, the evidence suggests that neither the Asian crisis nor the 1998 currency crisis of Russia and Latin America has any significant impact on the stock return series of Australia and the US, which are found to be covariance stationary and covariance nonstationary, respectively.

I. INTRODUCTION

The issue of covariance stationarity has been studied at some length by economists and statisticians in the past decade. Econometric models used in finance and commodity markets typically require the assumption that the data are covariance stationary. In the generalized autoregressive conditional heteroscedastic (GARCH) model, for example, there is the assumption that the unconditional second moments of a time series are constant, though it has been argued that this assumption is made more for statistical convenience than for the representation of reality. Pagan and Schwert (1990) propose three nonparametric tests for covariance stationarity and apply them to US common stock returns data. It is found that the unconditional variance of the US stock market returns cannot be assumed to be constant over 1834 to 1987. One drawback of Pagan and Schwert’s (1990) methodology, however, is that it relies on an auxiliary maintained assumption that the fourth unconditional moments of the data are finite. In a subsequent article, Loretan and Phillips (1994) show that stock market returns data used by Pagan and Schwert (1990) cannot assume finite fourth moments through a direct estimation of the time series’ maximal moment exponent. Loretan and Phillips (1994) also discuss formal moment based tests of variance constancy and apply them to data series of stock returns and exchange rates. It is found that the US stock market returns data used by Pagan and Schwert (1990) are not covariance stationary even under a general maintained hypothesis.

In a recent article, Omran and McKenzie (1999) argue that the rejection of covariance stationarity may be attributable to possible variance shifts caused by unusual events (e.g., oil crisis, stock market crash). Omran and McKenzie (1999) propose a two-step procedure which first filters the effects of the unusual events using an intervention model.
along the lines of Box and Tiao (1975), then applies the Lorentan and Phillips (1994) test to the intervention model’s residuals. Omran and McKenzie (1999) find that the hypothesis of variance constancy for the UK’s FTSE All-share index’s daily returns from 1970 to 1997 cannot be rejected after the effects of the 1973 oil crisis and the 1987 stock market crash have been properly filtered using an intervention model. On the other hand, if the Lorentan and Phillips test is applied directly on the raw data, then the hypothesis of constant variance is rejected. These results suggest that the FTSE All-share index’s returns have abrupt patterns of volatility during some exceptional periods. Outside these periods, however, the returns data can be reasonably described as covariance stationary.

In this paper is used a similar approach to that of Omran and McKenzie (1999) to investigate whether the daily stock returns of the Hong Kong, Singapore, Australian and US markets are covariance stationary over the past eight to eleven years. Intervention analysis is used to filter the effects of the 1997 Asian financial crisis and the Russian and Latin American currency crisis of 1998 on the variance of the returns. The empirical findings suggest that there have been significant variance shifts around the 1997 Asian financial crisis for the Hong Kong and Singapore stock returns data, for which the hypothesis of variance constancy cannot be rejected after the effects of the financial crisis have been properly filtered. On the other hand, the evidence suggests that neither the Asian crisis nor the 1998 currency crisis of Russia and Latin America has any significant impact on the stock return series of Australia and the US, which are found to be covariance stationary and covariance nonstationary, respectively.

The rest of this paper is organized as follows. Section II describes the data used for the analysis. Section III outlines the Omran and McKenzie (1999) testing procedure. Some empirical results are presented and discussed in Section IV. Section V concludes the paper.

II. DATA

The data used in this study are from the DataStream database and comprise the following stock return series: (i) daily returns of Australia’s All Ordinaries Index (AOI) from 2 June 1992 to 30 December 1999; (ii) daily returns of Hong Kong’s Hang Seng Index (HSI) from 2 January 1990 to 15 June 2000; (iii) daily returns of the Singapore Straits Time Index (STI) from 2 January 1990 to 15 June 2000; and (iv) daily returns of the Dow Jones Industrial Average index (DJIA) from 2 January 1992 to 31 December 1999. These constitute 1920, 2589, 2623 and 2016 observations respectively, excluding public holidays. As in Omran and McKenzie (1999), the word ‘return’ refers to the change in the natural logarithms of the series, and to eliminate potential calendar (weekdays and month of the year) and seasonal effects, the data are ‘demeaned’ through a regression on five dummy variables representing the days of the week and 12 dummies representing the months of the year. An autoregressive (AR) filter is then used to remove any linear serial dependence in the residuals of the regression. The lag length of the AR process is determined by the Schwartz Bayesian criterion (SBC). The residuals from these transformations, denoted by $u_t$, are used in the subsequent analysis.

III. METHODOLOGY

As in Omran and McKenzie (1999), the testing procedure is as follows: (i) begin with a test of variance homogeneity of $u_t$ using the Lorentan and Phillips (1994) test; (ii) construct a dynamic model along the lines of Box and Tiao (1975) capable of representing the effects of the interventions; (iii) re-apply the Lorentan and Phillips (1994) test to the standardized residuals from the intervention model.

Lorentan and Phillips (LP) (1994) test (i)

In conducting the test, the sample is split into two parts such that $n = n_1 + n_2$ with $n_1 = k_n n_2$, where $n$ is the total number of observations, $n_1$ and $n_2$ are the lengths of the first and second sub-samples, respectively, and $k_n$ is the constant for determining the relative lengths of $n_1$ and $n_2$. The null hypothesis

$$H_0 : E(u_2^{(1)}) = E(u_2^{(2)})$$

states that the unconditional variances in the two sub-samples are equal, where

$$u_2^{(1)} = n_1^{-1} \sum_{i=1}^{n_1} u_i^2$$

and

$$u_2^{(2)} = n_2^{-1} \sum_{i=n_1+1}^{n} u_i^2$$

are respectively the conditional variances of $u_t$ of the first and second sub-samples. Now, writing $d = u_2^{(1)} - u_2^{(2)}$, the null hypothesis can be stated equivalently as $H_0 : E(d) = 0$, and a test of $H_0$ is based on the statistic,

$$V_k(d) = \left\{ (1 + k_n)\gamma^2 \right\}^{-1/2} n_1^{1/2} d$$

where

$$\gamma^2 = \gamma_0 + 2 \sum_{j=1}^{L} \{ 1 - j/(L + 1) \} \gamma_j$$

are non-central chi-square distributed. A.K.F. Ho and A.T.K. Wan

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is the kernel-based estimator of the long-run variance of \( u_t^2 \), \( \gamma_j \) is the \( j \)th serial covariance of \( u_t^2 \), \( L \) is an appropriate lag truncation number,\(^1\) and \( k \) is the limiting value of \( k_n \) as \( n \to \infty \). It can be shown that \( n^{1/2}d \to_d N(0, (1 + k)\gamma_0) \) so long as \( \alpha \), the maximal moment exponent of the data, is larger than four. If \( \alpha \leq 2 \), then the test is inconsistent. In the intermediate case for which \( 2 < \alpha \leq 4 \) the test is consistent, but the test statistic’s limiting distribution is a ratio of correlated stable variates and depends on the value of \( k \). Critical values of the test are to be obtained through Monte Carlo simulation. Table 1 shows the estimated percentiles of \( V_k \) for a range of \( \alpha \) and \( k \) values. Note that the standard normal percentiles are used for the case of \( \alpha > 4 \).\(^2\)

So, the maximal moment exponent \( \alpha \) is crucial in determining the distribution of \( V_k \) and hence the critical values to be used in the test. One way to proceed is to estimate \( \alpha \) by the following estimators:

\[
\alpha_L(s) = 1 / \left\{ s^{-1} \sum_{j=1}^{s} \ln(-u_{(j)}) - \ln(-u_{(s+1)}) \right\} \tag{6}
\]

and

\[
\alpha_R(s) = 1 / \left\{ s^{-1} \sum_{j=1}^{s} \ln(u_{(a-j+1)}) - \ln(u_{(a-s)}) \right\} \tag{7}
\]

which are, respectively, the left-hand and right-hand tail estimators of \( \alpha \). Note that in Equation (6) and (7), the \( u \)'s are ordered such that \( u_{(1)} < u_{(2)} < u_{(3)} < \cdots < u_{(s)} \) and \( s \) is an arbitrary integer. To ensure that \( \alpha_L(s) \) and \( \alpha_R(s) \) are real quantities, it is assumed that \( n \) is large enough and \( s/n \) is small enough so that \( -u_{(s+1)} > 0 \) and \( u_{(a-s)} > 0 \). By virtue of Theorem 2 of Hall (1982), it can be shown that,

\[
s^{1/2} \{ \alpha_L(s) - \alpha \} \to_d N(0, \alpha^2) \tag{8}
\]

Table 1. Estimated percentiles of the variance constancy test statistic \( V_k \)

<table>
<thead>
<tr>
<th>( \alpha )</th>
<th>( k )</th>
<th>0.5%</th>
<th>2.5%</th>
<th>97.5%</th>
<th>99.5%</th>
</tr>
</thead>
<tbody>
<tr>
<td>2.89</td>
<td>0.5</td>
<td>-2.52</td>
<td>-1.95</td>
<td>1.69</td>
<td>2.17</td>
</tr>
<tr>
<td>2.89</td>
<td>1.0</td>
<td>-2.30</td>
<td>-1.79</td>
<td>1.82</td>
<td>2.35</td>
</tr>
<tr>
<td>2.89</td>
<td>1.5</td>
<td>-2.21</td>
<td>-1.72</td>
<td>1.90</td>
<td>2.44</td>
</tr>
<tr>
<td>3.04</td>
<td>0.5</td>
<td>-2.54</td>
<td>-1.96</td>
<td>1.71</td>
<td>2.18</td>
</tr>
<tr>
<td>3.04</td>
<td>1.0</td>
<td>-2.34</td>
<td>-1.81</td>
<td>1.82</td>
<td>2.34</td>
</tr>
<tr>
<td>3.04</td>
<td>1.5</td>
<td>-2.27</td>
<td>-1.76</td>
<td>1.92</td>
<td>2.47</td>
</tr>
<tr>
<td>3.58</td>
<td>0.5</td>
<td>-2.58</td>
<td>-1.97</td>
<td>1.84</td>
<td>2.41</td>
</tr>
<tr>
<td>3.58</td>
<td>1.0</td>
<td>-2.47</td>
<td>-1.87</td>
<td>1.91</td>
<td>2.48</td>
</tr>
<tr>
<td>3.58</td>
<td>1.5</td>
<td>-2.40</td>
<td>-1.85</td>
<td>1.96</td>
<td>2.54</td>
</tr>
<tr>
<td>Normal</td>
<td></td>
<td>-2.58</td>
<td>-1.96</td>
<td>1.96</td>
<td>2.58</td>
</tr>
</tbody>
</table>

where \( \alpha(s) \) is either \( \alpha_L(s) \) or \( \alpha_R(s) \), and \( s \) is the number of ordered statistics included in the computation and obtained by using,

\[
s = n^{2/3} / \ln(\ln(n)) \tag{9}
\]

Box and Tiao’s (1975) intervention model (ii)

An intervention model along the lines of Box and Tiao (1975) is used to take account of the variance shifts of the returns \( u_t \) around the time of the 1997 Asian financial crisis and during the 1998 Russian and Latin American currency crisis. Following Omran and McKenzie (1999),

\[
u_t = \sigma_t \varepsilon_t \tag{10}
\]

where \( \varepsilon_t \) is a Student’s \( t \) variate. Note that,

\[
\sigma_t^2 = \omega_0 + (\omega_1 + \delta_1)Z_{t1} + \{\delta_1 + b_1 (\sigma_t^2 - \omega_0 - \delta_1)\}X_{t1}
\]

\[
+ (\omega_2 + \delta_2)Z_{t2} + \{\delta_2 + b_2 (\sigma_t^2 - \omega_0 - \delta_2)\}X_{t2} \tag{11}
\]

is the return’s variance, where

\[
Z_{t1} = \begin{cases} 
0, & t \neq T_1 \\
1, & t = T_1
\end{cases}
\]

\[X_{t1} = \begin{cases}
0, & t \leq T_1 \\
1, & t > T_1 \text{ and } t < T_2
\end{cases}
\]

\[
Z_{t2} = \begin{cases} 
0, & t \neq T_2 \\
1, & t = T_2
\end{cases}
\]

\[X_{t2} = \begin{cases}
0, & t \leq T_2 \\
1, & t > T_2
\end{cases}
\]

\( T_1 \) represents 28 October 1997 (or 27 October 1997 for the DJIA returns series), on which the world’s major stock markets crashed, and \( T_2 \) represents 31 August 1998, on which the US and European stock prices fell as a result of the Russian and Latin American currency crisis.

Briefly, Equation (11) is interpreted as follows. Prior to the first intervention, volatility level is equal to \( \omega_0 \). On the crash day of 1997, volatility level jumps to \( \omega_0 + \omega_1 + \delta_1 \), before dying down at a rate \( b_1 \) to settle at \( \omega_0 + \delta_1 \). On the day of the second intervention due to the Russian and Latin American currency crisis, volatility level moves to \( \omega_0 + \omega_2 + \delta_2 \), then dies out exponentially at a rate \( b_2 \) and settles at \( \omega_0 + \delta_2 \). A graphical depiction of the responses is given in Fig. 1.

Note that the conditional density function for \( u_t \) is given by,

\[
f(u_t) = \frac{\Gamma\{w + 1/2\}}{\pi^{1/2} \Gamma\{w/2\}} w^{-1/2} \sigma_t^{-1} \left(1 + \frac{u_t^2}{\sigma_t^2}\right)^{-\left(w+1\right)/2} \tag{12}
\]

where \( w \) is the degrees of freedom parameter for the Student’s \( t \) distribution. Following Omran and McKenzie

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\(^1\) Loretan and Phillips (1994) set \( L = 8 \) for the monthly stock return series (also used by Pagan and Schwert (1990)), \( L = 12 \) for the daily stock return series and \( L = 20 \) for the exchange rate series.

\(^2\) Table 1 is obtained from Omran and McKenzie (1999).

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(1999), the parameters in Equation (11) are estimated by the Berndt–Hall–Hausman (BHHH) routine (Berndt et al., 1974).

Loretan and Phillips (1994) test on the residuals of the intervention models (iii)

The procedure is per step (i), except that the data for testing are the standardized residuals from the intervention model of step (ii). The idea is to investigate if the returns are covariance stationary after the effects of the Asian financial crisis and the Russian and Latin American currency crisis have been adjusted for. If the findings are different from those in step (i), then the differences are attributed to the effects of the interventions.

IV. EMPIRICAL RESULTS

The empirical results have been obtained with the RATS time series package, version 4 (Doan, 1995). Table 2 gives the estimates of $\alpha$, the maximal moment exponents of the four series obtained using the method of Hill (1975). Almost without exception, the estimates of $\alpha$ range between 2 and 4, suggesting that the Loretan and Phillips (1994) test is consistent, but modified critical values (as given in Table 1) are to be used for the test. Table 3 gives the values of $V_k(d)$, the LP test statistics for $k = 0.5, 1$ and 1.5. It is found that the null hypothesis of covariance stationarity as given by $H_0$ in Equation (1) can be rejected at both the 1% and 5% level of significance the for HSI and DJIA return series for $k = 0.5, 1$ and 1.5; for Singapore’s STI, $H_0$ can be rejected at both levels of significance only for $k = 1$ and 1.5, but for $k = 0.5, H_0$ is rejected only at the 5% level but not at the 1% level. Note that in the case of Australia’s AOI return series, the right-hand tail estimate of $\alpha$ is greater than four while it is the opposite for the left-hand tail estimate. Accordingly, to test $H_0$, we apply both the modified critical values of Table 1 and the critical values of the standard normal distribution. The results suggest that with the modified critical values, $H_0$ cannot be rejected for all values of $k$; with the standard normal critical values, $H_0$ can be rejected, though only at a larger than usual level of significance (say, > 10%) for $k = 1$ and 1.5.

So, the results show that there is strong evidence of covariance nonstationarity in the case of the HSI, STI and DJIA’s return series. For the AOI’s returns, the evidence is more in favour of covariance stationary. Next, the intervention model of Equation (11) is estimated with a proper allowance for the interventions of 1997 and 1998. A visual inspection of the series reveals quite definite structural breaks as a result of Asian financial crisis of October 1997. Typically, a sudden plunge in the return series after a period of relative stability characterizes these breaks, as shown in Fig. 2. For the DJIA’s return series, there is apparently one more distinct break due to the currency crisis of Russia and Latin America in 1998.

Table 4 shows that the parameter estimates of the intervention model (1) for the four series obtained using the BHHH routine. The results of intervention models are quite different for the four series. For the AOI’s and STI’s return series, only the parameters $\omega_0$ and $\delta_1$ differ significantly from zero at the 5% level of significance. Accordingly, the level of volatility increased from 0.463 ($\omega_0$) before the crash to 0.686 ($\omega_0 + \delta_1$) after the crash in the case of the AOI, and from 0.520 ($\omega_0$) to 2.040 ($\omega_0 + \delta_1$)

![Volatility response to interventions](image_url)

**Note:** Standard errors given in parentheses.

<table>
<thead>
<tr>
<th>Series</th>
<th>$s$</th>
<th>$\alpha_s(s)$</th>
<th>$\alpha_k(s)$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Daily stock returns:</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(1) All Ordinaries Index (AOI)</td>
<td>76</td>
<td>3.65</td>
<td>4.05</td>
</tr>
<tr>
<td>(2) Hang Seng Index (HSI)</td>
<td>91</td>
<td>2.85</td>
<td>2.90</td>
</tr>
<tr>
<td>(3) Straits Time Index (STI)</td>
<td>92</td>
<td>3.36</td>
<td>2.54</td>
</tr>
<tr>
<td>(4) Dow Jones Industrial Average (DJIA)</td>
<td>79</td>
<td>2.85</td>
<td>3.40</td>
</tr>
</tbody>
</table>

**Note:** ***significant at the 1% level. **significant at the 5% level.
Figure 2. Plots of the returns series
in the case of the STI. On the other hand, for the HSI’s returns, the variance surged to 7.455 (ω₀ + ω₁ + δ₁) immediately after the crash from 0.917 (ω₀) before the crash. Then ω₁ died down exponentially at the rate 0.991 (b₁) to the new variance level of 2.169 (ω₀ + δ₁). In other words, an increase in volatility was observed after the dying out of the market crash jump. Finally, for the DJIA’s return series, it is found that volatility level increased from 0.319 (ω₀) to 0.749 (ω₀ + δ₁) after the Asian financial crisis. Then, the variance of returns stayed at this level until the Russian and Latin American currency crisis, after which the variance increased to 0.903 (ω₀ + δ₂).

The final task is to reapply the LP test on the standardized residuals ut/σ, from the intervention models to check if the rejection/acceptance of the hypothesis of constant variance is attributable to the Asian financial crisis and the Russian and Latin American currency crisis. Tables 5 and 6 give the point estimates of the maximal moment exponents and the test results based on standardized residuals respectively. Clearly, for the HSI’s and STI’s return series, regardless of the value of k, the results indicate the failure to reject the null hypothesis of covariance stationarity. Notice that this finding differs from that when the same test is applied on the raw data, for which the hypothesis of covariance stationarity is rejected. On the other hand, for the AOI’s and the DJIA’s returns, the results are the same as those using the raw data, i.e., the two series are found to be covariance stationary and covariance nonstationary, respectively.

V. CONCLUSIONS

This paper investigates the hypothesis of covariance stationarity of the returns for four daily stock series using the

Table 4. Parameter estimates of the intervention models

<table>
<thead>
<tr>
<th>Series</th>
<th>(1) AOI</th>
<th>(2) HSI</th>
<th>(3) STI</th>
<th>(4) DJIA</th>
</tr>
</thead>
<tbody>
<tr>
<td>ω₀</td>
<td>0.463***</td>
<td>0.917**</td>
<td>0.520**</td>
<td>0.319**</td>
</tr>
<tr>
<td>(18.932)</td>
<td>(17.868)</td>
<td>(18.662)</td>
<td>(17.101)</td>
<td></td>
</tr>
<tr>
<td>ω₁</td>
<td>–</td>
<td>5.286**</td>
<td>–</td>
<td>–</td>
</tr>
<tr>
<td>(2.670)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>δ₁</td>
<td>0.223**</td>
<td>1.252**</td>
<td>1.520**</td>
<td>0.430**</td>
</tr>
<tr>
<td>(4.031)</td>
<td>(3.617)</td>
<td>(9.298)</td>
<td>(4.371)</td>
<td></td>
</tr>
<tr>
<td>h₁</td>
<td>–</td>
<td>0.991**</td>
<td>–</td>
<td>–</td>
</tr>
<tr>
<td>(246.590)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>ω₂</td>
<td>–</td>
<td>–</td>
<td>–</td>
<td>–</td>
</tr>
<tr>
<td>δ₂</td>
<td>–</td>
<td>–</td>
<td>–</td>
<td>–</td>
</tr>
<tr>
<td>b₂</td>
<td>–</td>
<td>–</td>
<td>–</td>
<td>–</td>
</tr>
<tr>
<td>w</td>
<td>8.452</td>
<td>3.553</td>
<td>3.934</td>
<td>5.266</td>
</tr>
</tbody>
</table>

Note: ** Significant at the 5% level.

Table 5. The point estimates of α (based on residual series)

<table>
<thead>
<tr>
<th>Series</th>
<th>s</th>
<th>α₁(s)</th>
<th>α₂(s)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Daily stock returns:</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(1) All Ordinaries Index (AOI)</td>
<td>76</td>
<td>3.59</td>
<td>3.74</td>
</tr>
<tr>
<td>(2) Hang Seng Index (HSI)</td>
<td>91</td>
<td>2.63</td>
<td>3.14</td>
</tr>
<tr>
<td>(3) Straits Time Index (STI)</td>
<td>92</td>
<td>2.59</td>
<td>3.03</td>
</tr>
<tr>
<td>(4) Dow Jones Industrial Average (DJIA)</td>
<td>79</td>
<td>3.11</td>
<td>3.97</td>
</tr>
</tbody>
</table>

Note: Standard errors given in parentheses.

Table 6. The estimated values of V_k(d) (for residuals series)

<table>
<thead>
<tr>
<th>Series</th>
<th>Value of V_k(d)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Daily stock returns:</td>
<td></td>
</tr>
<tr>
<td>(1) All Ordinaries Index (AOI)</td>
<td>1.073 0.257 -0.108</td>
</tr>
<tr>
<td>(2) Hang Seng Index (HSI)</td>
<td>-1.509 0.972 0.258</td>
</tr>
<tr>
<td>(3) Straits Time Index (STI)</td>
<td>0.481 1.061 0.507</td>
</tr>
<tr>
<td>(4) Dow Jones Industrial Average (DJIA)</td>
<td>-2.380** -3.629*** -3.285***</td>
</tr>
</tbody>
</table>

*** Significant at the 1% level.
** Significant at the 5% level.

Loretan and Phillips (1994) test and Box and Tiao’s (1975) intervention analysis. The interventions of relevance are those corresponding to the Asian financial crisis of 1997 and the Russian and Latin American currency crisis of 1998. The results suggest that the Asian financial crisis is largely responsible for the rejection of the hypothesis of covariance stationarity for the return series of Hong Kong’s Hang Seng Index and Singapore’s Straits Time Index. It is found that the series can be reasonably described as covariance stationary when the impact of the financial crisis is properly filtered. On the other hand, while significant variance shifts are found around the breaks for the All Ordinary Index and the Dow Jones Industrial Index’s return series, the breaks have no significant effect on the stationarity properties of the series. Clearly, as is discussed in Omran and McKenzie (1999), more efforts are needed to develop time series techniques for investigating stock volatility in the presence of structural breaks and outliers, and this remains interesting departure for future research.

ACKNOWLEDGEMENTS

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REFERENCES


