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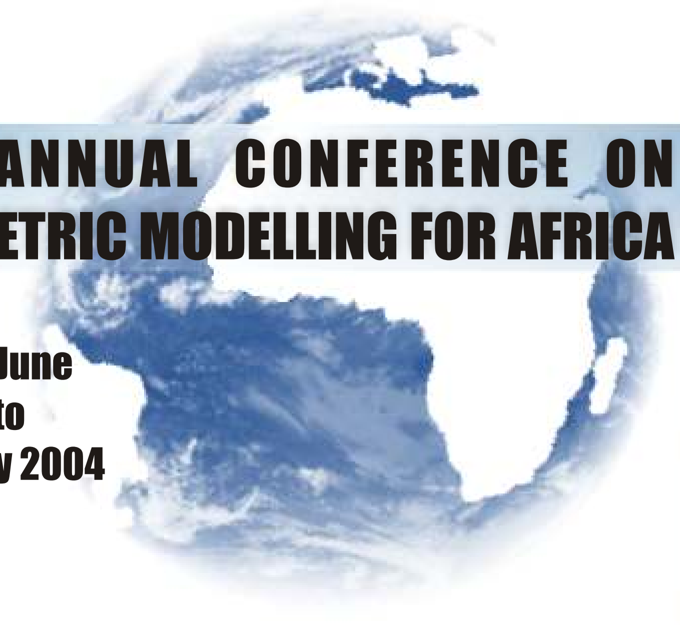

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The Causal Links between Equity Market Prices: The Case of Australia and its Major Trading Partners

Abdulnasser Hatemi-J
and
Eduardo D. Roca

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Abdulnasser Hatemi-J* and Eduardo D. Roca **

* Department of Economics and Political Sciences, University of Skovde, P-O Box 408, 521 28 Skovde, Sweden; tel: 46 500 44 87 31, fax: 46 500 44 87 99, email: abdulnasser.hatemi-j@ish.his.se.
Home page: http://www.his.se/ish/personal/personliga_sidor/abdulnasser_eng.htm

** School of Accounting, Banking and Finance, Griffith University, Nathan, Queensland, Australia 4111; tel: 61 7 3875 7583; fax: 61 7 3875 7760; email: E.Roca@griffith.edu.au.
Home page: http://www.gu.edu.au/school/abf/staff/eroca.html

Abstract

We examine the equity market price interdependence between Australia, on one hand, and Japan, US, UK, Hong Kong, Singapore, Taiwan and Korea, on the other hand, based on bootstrap causality tests with leveraged adjustments. We cover the period January 1, 1993 to September 10, 2001 taking into account the Asian financial crisis in 1997. We find that for both periods - before and after the Asian crisis, no causal linkages existed between Australia and these markets. This therefore indicates that this group of markets can serve as possible venues for international diversification. This also signifies that the transmission of information between the equity market Australia and those of its trading partners is informationally efficient.

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1. Introduction

This paper re-examines the issue of international equity market linkages – an issue that has importance in terms of international portfolio diversification and policy formulation. There is a significant body of literature on equity market integration but there is no conclusive evidence yet on the issue. Depending on the data, methodology, and theoretical models used, some studies have found that markets are integrated (see Agmon, 1972; Ripley, 1973; Hillard, 1979; Ibbotson, et. al., 1982; Jaffe and Westerfield, 1985; Schollhammer and Sand, 1987; Wheatley, 1988; Hamao, et. al., 1990; Espitia and Santamaria, 1994, among others) while others have found the opposite (see, for example, Grubel, 1968; Makridakis and Wheelwright, 1974; Adler and Dumas, 1983; Jorion and Schwartz, 1986; Levy and Lerman, 1988; Dwyer and Hafer, 1988; Jorion, 1989; Smith, et. al., 1995, among others).

The re-examination of international equity market integration is undertaken focusing on the equity markets of Australia and its major trading partners (Japan, US, UK, Hong Kong, Singapore, Taiwan and Korea) and using an Australian perspective. This group of markets provide a logical basis for such re-examination. First, Australia and its major trading partners represent markets that are relatively performing well and therefore are possible venues for international portfolio diversification. As is well-established in portfolio theory, the benefits arising from diversification depends on the degree of linkages between assets or markets. Secondly, in terms of policy implication, if the Australian equity market is found to be significantly interdependent with the equity markets of its major trading partners, then there is the danger that shocks in other markets may spill-over to Australia (the so-called "contagion effect" - see King and Wadwhani, 1990), and in this situation, there may be a need for closer cooperation between the prudential and monetary regulators of Australia and the other markets. Finally, there is a dearth of studies on equity market integration that take a non-US perspective; thus, the use of an Australian perspective addresses this particular need.

Unfortunately, the linkage between the equity markets of Australia and its major trading partners is far from fully understood. Very few studies have been done on Australia's equity market linkage with other countries and the bulk of these have concentrated on Australia's linkage with the developed markets of the US, UK, Japan, Canada, and Germany (see, for instance, Allen and MacDonald, 1995 and McNelis, 1993). Except for one study, i.e. Roca (1999), there is no other study that has investigated the exact set of markets that are included in this study. Moreover, many of the previous studies investigated linkages in terms of correlation. However, it is well known that correlation has a number of deficiency including instability and the existence of lags. Hence, even if there are low contemporaneous correlations, this can be deceptive if this is changing. Thus, the use of causality can provide a better understanding of how this correlation can change.

Furthermore, methodologically, most of these studies, including that of Roca (1999) are based on asymptotic methods of inference. It has been shown that when data possess non-normalities and ARCH effects, asymptotic methods can lead to bias statistical inferences. Financial data, particularly equity market prices, are well-known to possess these qualities (see, for example, Chunchachinda, et. al., 1997 for more recent evidence). It is also well-accepted in the finance literature that heteroskedasticity is a characteristic of financial time series data. For a certain period of time, volatilities can be successively large and in another period of time, these can be successively low. Diebold (1988) and Baillie and Bollerslev (1989) have shown that ARCH effects are often present in high frequency financial time series data such as daily and weekly stock market data. A number of studies have already shown that each of the markets covered in the present study are characterised by ARCH effects (see, for instance, Nicholls and Tonuri, 1995; Bae and Karolyi, 1994; among others).

The presence of ARCH characteristics affects the correlation structure among the time series. Often, the calculation of pairwise correlations is based on the assumption that the time series involved are mean and variance stationary in which the conditional and unconditional estimates are the same. If, however, this assumption is not true as in the case where ARCH effects are present, the computed correlations will be greater than what they should actually be. In the area of finance, this has far-reaching significance as hedging and diversification strategies are based on knowledge of the correlation between different markets or assets. If correlations are overestimated than what they should be, this leads to underestimation of diversification benefits and will in turn lead to erroneous portfolio construction strategies.

In order to avoid this problem, we therefore make use of a different econometric methodology. We apply the Toda-Yamamoto (1995) test of causality which we bootstrap. In order to make sure that the presence of ARCH effects does not render biased results, we apply the leveraged bootstrap as suggested by Davison and Hinkley (1999) and Hacker and Hatemi-J (2002). This method has been shown by Hacker and Hatemi-J (2002) to produce more reliable results compared to asymptotic methods. Thus, our study can produce more robust results and will therefore contribute new evidence on the issue of financial market integration.

The results show that there are no causal linkages between the equity market prices of Australia and its major trading partners. This can imply that this group of markets provide a good avenue for international portfolio diversification. It can also imply that the relationship between these markets is one that is informationally efficient.

The remaining parts of the paper are organised in this manner. Section 2 presents a brief review of the literature on equity market integration. Section 3 discusses the characteristics of the data and presents the results of some diagnostic tests. Section 4 provides a discussion of the methodology. The empirical results are presented in Section 5 while Section 6 concludes the study.

2. Literature Review

As stated earlier, the issue of financial market, particularly equity market, price linkages has been examined quite extensively in the literature. However, there is no clear conclusion as to the extent of these linkages. There is also no clear agreement as to which markets are significantly linked although the results seem to point to the existence of a linkage between certain groups of equity markets based on some unifying or common factor, such as close regional, economic, and geographical relationships. To, et. al. (1994) found the following clusters: Japan and Asian emerging markets, and the UK and African emerging markets. Hillard (1979) discovered a close association among intra-continental markets during the oil crisis of 1973 while Jorion (1989) reported a high degree of linkage among European continental markets. An Anglo-Saxon cluster was also reported by Jorion (1989). Thus, there is a need to examine whether other possible market clusters can be identified. The group consisting of Australia and its major trading partners is a logical candidate for this investigation.

There is, however, consensus in the literature that the US equity market is the most influential stock market in the world (see, for instance, Chowdhury, 1994 and Masih and Masih, 1999, among others). There is overwhelming evidence that the

United States leads other markets, with the exception of such markets as Korea, Taiwan and Thailand (see, for example, Eun and Shim, 1989; Fischer and Palasvirta, 1990). But there are some studies such as those of Granger and Morgenstern (1970) and Hillard (1979) that reported no lead/lag relationships among markets.

In terms of studies dealing with Australia's interaction with other markets, the results are also mixed, most especially in relation to long-term linkages. For example, Roca (1999), applying cointegration analyses on weekly MSCI data covering the period 1974 to1995 found no significant relationship between Australia and its major trading partners. Allen and MacDonald (1995), using pairwise cointegration analysis on monthly stock price data from 1970 to 1992, found a significant relationship between Australia and the markets of the UK and Hong Kong but not with the US and Japan. Blackman, et. al. (1994), also using cointegration procedures on monthly stock price data for the period 1970 to 1989, found significant linkages between Australia, the US, UK, Japan, and Hong Kong. Other studies (for example, Corhay, et. al., 1995; Kwan, et. al., 1995 and Chan, et. al, 1997) also provide the same conflicting findings. Thus, there is a need to further examine the issue of long-term linkages between Australia and other markets. The present study seeks to address this knowledge gap.

3. Data Properties

The stock markets of the US, Japan and the UK are the first, second and third largest equity markets in the whole world. These markets have been highly deregulated and liberalised for already a long period of time, particularly the US and the UK, and Japan starting in the late 1970s. Australia, Hong Kong, Singapore, Taiwan and Korea, have equity markets which are comparable in terms of size. The equity markets of Australia, Hong Kong and Singapore are markets that have been deregulated and liberalised for quite a period of time - Hong and Singapore starting in the early 1970s and Australia commencing in the early 1980s. The markets of Taiwan and Korea, however, are still deregulating and liberalising their markets, a process which began in the late 1980s. Thus, given the relatively small size of the Australian market, we expect it not to be influential on the US, UK and Japanese markets. However, we expect that the opposite will take place – i.e., the US, UK and Japanese markets will influence Australia. With respect to Australia's relation with the other markets, we do not expect any significant linkage, most especially with Taiwan and Korea.

The study utilises daily MSCI price index data, expressed in US dollars, covering the period January 1, 1993 to September 1, 2001. The sample period is broken into two sub-periods to take into account the on-set of the Asian financial crisis of 1997. The first sub-period is from January 1, 1993 to July 1, 1997 while the second sub-period is from January 1, 1998 to September 1, 2001. The MSCI data are well-known for their reliability and absence of double-listing.

Table 1 presents the contemporaneous correlation between the markets. For both periods, it can be seen that Australia is least correlated with Japan and most correlated with the UK. Australia became more correlated with its trading partner, with the exception of the UK during the second period. The correlations during both periods are less than one; therefore, based on these, there seems to be scope for international portfolio diversification gains for Australian investors in these markets. Correlations, however, cannot capture the long-term interactions between the markets in a reliable way as they are time varying. Hence, these results regarding the linkages between the equity markets of Australia and its trading partners based on the correlations results should be taken with caution. What is needed is a long-term causality analysis between the markets. This is undertaken in this study with the use of the Toda-Yamamoto (1995) test which we bootstrap with leveraged adjustments. The next section explains this procedure.

(Insert Table 1 here)

Tests of normality were conducted. All variables are non-normal in both sub-periods (both individually and systemwise). There are also significant multivariate ARCH effects. Therefore, given this situation, in order to obtain more accurate results, we bootstrap with leveraged adjustments the Toda-Yamamoto (1995) causality test.

4. Methodology

In this section we explain the methodology that is utilised to carry out the empirical analyses of our paper. First, we describe the Perron (1989) test for unit roots. Next, we represent the Toda and Yamamoto (1995) test statistics for Granger (1969) causality between non-stationary (integrated) variables. Given that the choice of lag order is a crucial issue in this aspect we will also define a new information criterion that performs well for non-stationary data. Finally, we describe the bootstrap simulation technique that is employed to produce more precise critical values for tests of Granger causality.

It is well accepted in the literature that it is important to check the time series properties of the underlying data in order to avoid false and spurious inference. It is also well known that standard tests for unit roots have low power if structural breaks have occurred during the period of study. In order to take into account the effect of a potential structural break arising from the Asian crisis of 1997 we make use of a test suggested by Perron (1989) to test for the integration order of the variables. This test permits for a structural break in both the mean value and the deterministic trend of the variable under investigation for unit root. The Perron (1989) test for unit roots of variable z is based on the following regression:

$$z_t = c_1 + c_2 D_t + d_1 t + d_2 D_t t + gJ_t + \gamma z_{t-1} + \sum_{i=1}^{m} b_i \Delta z_{t-i} + \varepsilon_t \tag{1}$$

where

t = the time period (the linear trend term),

D_t = a dummy variable that takes value zero for the time period before break and one for the rest of the period,

J_t = a dummy variable that is equal to one if the time period t is the first period after that of the structural break, and is zero otherwise,

Δ = the first difference operator and

ε_t = a white noise error term.

The null hypothesis of a unit root is $\gamma = 1$, and the alternative hypothesis of stationarity is $\gamma < 1$. If necessary, we will include lag values of Δz in equation (1) to make sure that the error term is white noise. The optimal number of lagged differences (m) is determined by including more lags until the null hypothesis of independence for ε_t is not rejected by the Ljung-Box test at the 5% significance level.

The next step in our empirical inquiry is to examine the causal nexus of the variables. By causality, we mean causality in the Granger (1969) sense. That is, we aim to find out whether one variable precedes another variable or not. For this purpose, the following vector autoregressive model of order p, VAR(p), is utilised:

$$y_t = \nu + A_1 y_{t-1} + \ldots + A_p y_{t-p} + \varepsilon_t , \tag{2}$$

Here:

y_t = the number of variables in the VAR model, which is six in this particular case,

v = an eight-dimensional vector of intercepts,

ε_t = an eight-dimensional vector of error terms, and

A_r = a 8×8 matrix of parameters for lag r (r = 1, …, p).

An important matter in this regard to pay extra attention to is the choice of optimal lag order (p) because all inference in the VAR is naturally based on the chosen lag order. To accomplish this, we make use of a new information criterion introduced by Hatemi-J (2003). This information criterion is shown to perform well for choosing the optimal lag order, especially if the variables in the VAR model are integrated. The Hatemi-J (2003) criterion is defined as the following:

$$HJC = \ln\left(det\ \hat{\Omega}_j\right) + j\left(\frac{n^2 \ln T + 2n^2 \ln(\ln T)}{2T}\right), \qquad j = 0, \cdots, p, \quad (3)$$

Here:

ln = the natural logarithm,

$det\,\hat{\Omega}_j$ = the determinant of the estimated variance-covariance matrix of ε_t for lag order j,

n = the number of variables in the VAR model, and

T = the sample size used to estimate the VAR model.

The optimal lag order is chosen based on the minimization of equation (3). It is well known in the econometric literature that critical values based on standard asymptotical distributions can not be used to test for restrictions in the VAR model if the variables are integrated (see Sims et. al. (1990)). To circumvent this problem Toda and Yamamoto (1995) suggested the following augmented VAR(p+d) model to be used for tests of causality if the variables are integrated:

$$y_t = \hat{v} + \hat{A}_1 y_{t-1} + \ldots + \hat{A}_p y_{t-p} + \ldots + \hat{A}_{p+d} y_{t-p-d} + \hat{\varepsilon}_t\,, \quad (4)$$

where the circumflex above a variable represents its estimated value. Note that d is equal to the integration order of the variables. The null hypothesis of non-Granger causality can be defined as the following:

H_0: the row j, column k element in A_r equals zero for r = 1,…, p. (5)

It should be pointed out that the parameters for the extra lag(s), i.e. d, are unrestricted in testing for Granger causality. Toda and Yamamoto (1995) show analytically that these unrestricted parameters ensure that the asymptotical distribution theory can be applied. Assuming initial values are given, the estimated VAR(p+d) model can be written compactly as:

$$Y = \hat{D}Z + \hat{\delta}\,. \tag{6}$$

where

$Y := (y_1, \cdots, y_T)$ $(n \times T)$ matrix,

$\hat{D} := (\hat{v}, \hat{A}_1, \cdots, \hat{A}_p, \cdots, \hat{A}_{p+d})$ $(n \times (1 + n(p+d)))$ matrix,

$$Z_t := \begin{bmatrix} 1 \\ y_t \\ y_{t-1} \\ \vdots \\ y_{t-p-d+1} \end{bmatrix} \quad ((1 + n(p+d)) \times 1) \text{ matrix, for } t = 1, \ldots, T,$$

$Z := (Z_0, \cdots, Z_{T-1})$ $((1 + n(p+d)) \times T)$ matrix, and

$\hat{\delta} := (\hat{\varepsilon}_1, \cdots, \hat{\varepsilon}_T)$ $(n \times T)$ matrix.

The modified Wald (MWALD) test statistic, introduced by Toda-Yamamoto, for testing the null hypothesis of non-Granger causality is defined as

$$MWALD = (C\hat{\beta})' \left[C\left((Z'Z)^{-1} \otimes S_U\right)C' \right]^{-1} (C\hat{\beta}) \sim \chi^2_p\,, \tag{7}$$

Here:

$\otimes$ = element by element multiplication operator (the Kronecker product).

C = a $p \times n(1+n(p+d))$ matrix. Each of the p rows of C is associated with the restriction to zero of one parameter in β. The elements in each row of C acquire the value of one if the related parameter in β is zero under the null hypothesis, and they get the value of zero if there is no such restriction under the null.

S_U = the estimated variance-covariance matrix of residuals in equation (6) when the restrictions implied by the null hypothesis of non-Granger causality is not imposed (unrestricted model).

$\hat{\beta} = vec(\hat{D})$, where *vec* signifies the column-stacking operator.

The MWALD test statistic is asymptotically χ^2 distributed, conditional on the assumption that the error terms are normally distributed, with the number of degrees of freedom equal to the number of restrictions to be tested. The number of restrictions is equal to p in this case. However, Hacker and Hatemi-J (2002) show through Monte

Carlo experiments that the MWALD test statistic overrejects the null hypothesis, especially if the data generating process for the error terms is characterised by non-normality and autoregressive conditional heteroscedasticity (ARCH). To improve on the inference based on tests for causality under such circumstances, the authors suggest making use of leveraged bootstrap simulations. The bootstrap method, which was originally introduced by Efron (1979), is based on resampling the underlying data to estimate the distribution of a test statistic. Using this distribution can decrease bias in inference by providing more precise critical values.

To perform the bootstrap simulations we first estimate regression (6) with the restriction for the null hypothesis of no Granger causality imposed. For each bootstrap simulation we generate the simulated data, Y^*, based on the coefficient estimates from this regression, $\hat{D}$; the original Z, data; and δ^* (the bootstrapped residuals). That is:

$$Y^* = \hat{D}Z + \delta^* \tag{8}$$

Note that the bootstrap residuals are based on T random draws with replacement from the regression's modified residuals, each with equal probability of $1/T$. The mean of the resulting set of drawn modified residuals is subtracted from each of the modified residuals in that set. This adjustment is done to make sure that the mean value of the bootstrapped residuals is zero. The modified residuals are the regression's raw residuals modified to have constant variance, through the use of *leverages*.[i] It should be mentioned that $\hat{D}$ is the estimated value of the parameters in equation (6). That is:

$$\hat{D} = YZ'(ZZ')^{-1}. \tag{9}$$

We run the bootstrap simulation 1000 times and then we calculate the MWALD test statistic each time. By this way we are in a position to generate the empirical distribution for the MWALD test statistic. Subsequent to these 1000 estimations we find the $(\alpha)th$ upper quantile of the distribution of bootstrapped MWALD statistics and attain the α-level "bootstrap critical values" (c_α^*). We generate the bootstrap critical values for 1%, 5% and 10% significance levels. The next step is to calculate the MWALD statistic using the original data (not the bootstrapped simulated data). Then, the null hypothesis of no Granger causality is rejected based on bootstrapping if the actual MWALD is greater than c_α^*. The simulations are performed by making use of a programme procedure written in GAUSS.[ii]

5. Empirical Results

Before undertaking the causality analysis, we perform the Perron (1989) test for unit roots on the data. These tests have better power properties because they allow for a shift in the mean value as well as a shift in the trend for the underlying variable. The estimation results from these tests are reported in Table 2. For the null hypotheses of I(1), i. e. integration of the first order, the estimated test statistics are found to be less than the critical values at the conventional significance levels. Hence, the null hypothesis that each variable is I(1) cannot be rejected. The next step is to investigate whether each variable becomes stationary after taking the first difference. Nevertheless, the null hypothesis that each variable is I(2) is rejected at the one percent significance level. Hence, we can conclude that each variable contains one unit root. This implies that we have to pay attention to the integration properties of the data in order to avoid spurious and false inference.

(Insert Table 2 here)

We also tested for parameter stability by applying the Chow (1960) multivariate test. The period for a structural break was the second week in July 1997 arising from the Asian crisis of 1997. The results showed that the null hypothesis of no structural break could be rejected even at the one percent significance level. To allow for the possibility that the economic process has changed during the sample period we make use of sub-periods.

The results of the Toda-Yamamoto (1995) causality analysis based on leveraged bootstrap methods are presented in Tables 3 (for sub-period 1) and 4 (for sub-period 2). None of the null hypotheses can be rejected at different levels of significance. Thus, no significant long-term causal linkages were found between Australia and any of its trading partners during the first sub-period. The same results were obtained during the second sub-period. Thus, the Asian financial crisis did not lead to a change in the long-term relationship between Australia and its trading partners. These results are in line with those of the existing literature (see, for instance, Roca, 1999). These could indicate the absence of long-term equity market integration between Australia and its trading partners. These results could also imply that the transmission of information between Australia and its trading partners is efficient. Thus, there are benefits to be obtained for long-term Australian investors by diversifying into the equity markets of its trading partners.

Insert Tables 3 and 4 here.

6. Conclusion

We examined the equity market interaction between Australia and its major trading partners (Japan, US, UK, Hong Kong, Singapore, Taiwan and Korea) using the Toda-Yamamoto (1995) causality test which we bootstrap with leveraged adjustment. This method is more efficient compared to alternative methods, especially if ARCH effects exist and/or if the underlying distribution is not characterized by normal distribution. Formal tests for multivariate normality and ARCH showed that the underlying daily data in this study is non-normal and ARCH effects exist. Thus, applying bootstrap simulations with leveraged adjustments is a necessary condition to draw valid inference. We also paid extra attention to choosing the appropriate dynamics.

Based on conducted Granger causality tests, we found no significant long-term causal linkages in both the period before and after the Asian crises. These could indicate the absence of long-term equity market integration between Australia and its trading partners. These could also imply that the transmission of information between Australia and its trading partners is efficient. From the point of view of long-term Australian investors, the equity markets of its trading partners offer diversification benefits.

Table 1

Contemporaneous Correlations Between Australia and Selected Major Trading Partners

Trading Partner	Correlation Coefficient	
	Sub-Period 1	Sub-Period 2
Australia	1.000	1.000
Japan	-0.174	0.041
US	0.036	0.114
UK	0.633	0.434
Hong Kong	0.043	0.331
Singapore	-0.007	0.247
Taiwan	0.046	0.113
Korea	0.301	0.333

Table 2

Test for Unit Roots Using the Perron Test

	H_0: I(1), H_1: I(0)	H_0: I(2), H_1: I(1)
Australia	-2.951 (0)	-45.680 (0) ***
Japan	-2.076 (0)	-42.000 (0) ***
US	-1.722 (7)	-19.180 (6) ***
UK	-2.701 (0)	-45.830 (0) ***
Hong Kong	-2.825 (50)	-5.450 (49) ***
Singapore	-2.104 (1)	-37.71 (0) ***
Taiwan	-2.538 (32)	-7.377 (31) ***
Korea	-2.062 (48)	-6.060(47) ***

Notes:

(a) The critical value is –4.90and –4.24 at the 1% and 5% significance level, respectively.

(b) The notation *** implies significance at the one percent significance level.

(c) The numbers in the parentheses indicate the number of lags required to remove potential autocorrelation in the Perron regression (equation 1) at the 5% significance level using the Breusch-Godfrey test.

(d) The break for the Asian crisis is from July 2-December 31, 1997.

Table 3

Causality Tests Based on Bootstrap Simulation Techniques for First Sub-Period

THE NULL HYPOTHESIS	THE ESTIMATED TEST VALUE (MWALD)	1% BOOTSTRAP CRITICAL VALUE	5% BOOTSTRAP CRITICAL VALUE	10% BOOTSTRAP CRITICAL VALUE
SP_{AU} ≠> SP_{HK}	0.695	7.253	3.677	2.514
SP_{HK} ≠> SP_{AU}	0.477	5.138	2.927	2.267
SP_{AU} ≠> SP_{JP}	0.005	9.758	4.049	2.747
SP_{JP} ≠> SP_{AU}	0.351	6.348	3.291	2.497
SP_{AU} ≠> SP_{KR}	0.001	4.213	2.592	2.042
SP_{KR} ≠> SP_{AU}	0.035	4.011	2.617	2.061
SP_{AU} ≠> SP_{SIN}	2.047	8.746	6.184	3.123
SP_{SIN} ≠> SP_{AU}	0.002	6.562	3.205	2.476
SP_{AU} ≠> SP_{TW}	0.002	5.677	2.969	2.578
SP_{TW} ≠> SP_{AU}	0.139	7.429	3.940	3.036
SP_{AU} ≠> SP_{UK}	0.021	8.046	3.509	2.874
SP_{UK}≠> SP_{AU}	0.280	5.116	2.707	2.116
SP_{AU} ≠> SP_{US}	0.715	7.446	4.617	3.119
SP_{US} ≠> SP_{AU}	0.821	8.616	5.049	3.647

Notes:

(a) The notation ≠> implies non-Granger causality.

(b) MWALD represents the modified Wald test statistic as described in equation (7).

(c) The lag order of the VAR model, *p*, was set to one for this sub-period. Also the augmentation lag, *d*, was set to one since each variable contains one unit root.

Table 4

Causality Tests Based on Bootstrap Simulation Techniques for Second Sub-Period

THE NULL HYPOTHESIS	THE ESTIMATED TEST VALUE (MWALD)	1% BOOTSTRAP CRITICAL VALUE	5% BOOTSTRAP CRITICAL VALUE	10% BOOTSTRAP CRITICAL VALUE
$SP_{AU} \not\Rightarrow SP_{HK}$	0.479	7.144	3.348	2.925
$SP_{HK} \not\Rightarrow SP_{AU}$	0.546	7.844	2.967	2.471
$SP_{AU} \not\Rightarrow SP_{JP}$	0.358	7.632	3.853	2.523
$SP_{JP} \not\Rightarrow SP_{AU}$	0.256	5.688	3.501	3.001
$SP_{AU} \not\Rightarrow SP_{KR}$	0.271	9.863	3.880	3.029
$SP_{KR} \not\Rightarrow SP_{AU}$	0.749	9.969	3.868	2.321
$SP_{AU} \not\Rightarrow SP_{SIN}$	0.517	6.193	4.764	2.409
$SP_{SIN} \not\Rightarrow SP_{AU}$	1.400	9.380	5.290	3.738
$SP_{AU} \not\Rightarrow SP_{TW}$	0.125	8.082	3.591	2.828
$SP_{TW} \not\Rightarrow SP_{AU}$	0.598	5.264	3.481	2.525
$SP_{AU} \not\Rightarrow SP_{UK}$	0.074	8.136	4.479	2.589
$SP_{UK} \not\Rightarrow SP_{AU}$	0.728	5.048	2.783	2.176
$SP_{AU} \not\Rightarrow SP_{US}$	0.191	6.977	4.601	2.919
$SP_{US} \not\Rightarrow SP_{AU}$	0.337	6.566	4.943	2.766

Notes:

(a) The notation $\not\Rightarrow$ implies non-Granger causality.

(b) MWALD represents the modified Wald test statistic as described in equation (7).

(c) The lag order of the VAR model, p, was set to two for this sub-period. Also the augmentation lag, d, was set to one since each variable contains one unit root.

Endnote

[i] For more details on leverage adjustment, see Davison, and Hinkley (1999) and Hacker and Hatemi-J (2002). The latter authors introduce this adjustment for multivariate equation cases.

[ii] A program procedure written in GAUSS to conduct leveraged bootstrap simulations is available on request from the authors.

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