# The Interdependence of Share Markets in the Developed Economies of East Asia

by

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**Abstract**

The interdependence of Japan's Nikkei (JN), Taiwan Weighted (TW), Singapore Strait Times (SST), Korea Composite (KC) and Hang Seng (HS) SPIs is tested on 2739 daily observations for July 8, 1990 to July 6, 2000. A preliminary correlation analysis reveals increased correlation in the post Asian currency crisis era. Standard multivariate cointegration tests reveal at most one cointegrating vector while smooth bivariate cointegration applies in the following pairwise comparisons: KC $\rightarrow$ TW, HS $\rightarrow$ TW, JN $\rightarrow$ TW. When structural breaks in cointegrating vectors are accommodated, there is the following evidence of cointegration with structural breaks: KC/SST (Jan 1998), SST/JN (Jan 1992), KC/JN (Jan 1996), HS/JN (Jan 1992). ECM reveals evidence of causal relationships, when there is smooth cointegration: $\Delta$SST $\rightarrow$ $\Delta$TW, $\Delta$KC $\rightarrow$ $\Delta$TW, $\Delta$HS $\rightarrow$ $\Delta$TW, $\Delta$JN $\rightarrow$ $\Delta$TW, $\Delta$HS $\rightarrow$ $\Delta$JN. Short run Granger also holds in these cases except for $\Delta$KC$\rightarrow$ $\Delta$TW. ECM estimation either side of break points reveals the following long lags and leads: $\Delta$KC $\rightarrow$ $\Delta$SST (pre not post), $\Delta$SST $\rightarrow$ $\Delta$KC (post not pre), $\Delta$SST $\rightarrow$ $\Delta$JN (post not pre), $\Delta$KC $\rightarrow$ $\Delta$JN (pre and post), $\Delta$HS $\rightarrow$ $\Delta$JN (pre and post). Granger short run causality applies in all these cases except for $\Delta$KC $\rightarrow$ $\Delta$SST.

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## 1. Introduction

This study is designed to assess the degree of interdependence among the share markets of the five more developed economies in the East Asian region. These nations are the most developed in this growing region of the world economy and are of strategic significance to the further development of Asia. Japan is the world's second largest economy while Singapore, Hong Kong, South Korea and Taiwan are newly industrialised countries (NICs) enjoying per capita incomes comparable to nations such as Australia which sit in the middle of the OECD rankings of per capita incomes.

These five developed economy share markets provide a significant channel for FDI and portfolio investment flowing from the developed nations to the emerging countries of the region. For example, an independent Hong Kong, Taiwan, Japan and Singapore accounted for 83 per cent of China's inward FDI over the period 1979-1993.[1] This East Asian group of equity markets is significant also in an international context representing 20 per cent of global market capitalisation in May 2000.[2] They also provide a vehicle for mutual fund investment in Asia, in particular for investment companies seeking the benefits flowing from international diversification. The performance of these five share markets collectively is an issue of considerable interest to professional international investors. Cheung and Mak (1992) for example indicate that several of the world's leading fund managers have established financial vehicles concentrated only in this region as media for international risk diversification.

Following Bekaert and Harvey (1995) two polar views of interdependence are proposed in which case, there are two extremities defined in the following way: complete or perfect independence arises if in a multi country cointegrative sense, the five East Asian share price indices follow a single, common stochastic trend. This definition is also suggested by Siklos and Ng (2001) in an analysis of Asia Pacific stock market integration. For the bivariate case, complete

[1] Islam and Chowdhury (1991, table 17.7, p. 274)
[2] *Asia Week*, May 26, 2000, p. 92.

interdependence is suggested if all twenty bivariate studies of share price movements among the East Asian five suggest the existence of a long run relationship. Following Bekaert and Harvey, at the opposite pole are segmented markets and for the current purpose, it is assumed that complete segmentation arises if in a bivariate context, there are no studies revealing long run relationships between the individual share price series, nor is there any evidence of a single, common stochastic trend in a multivariate framework. In between these polarities lie varying degrees of interdependence indicated by the presence of long run bivariate relationships in some, but not all individual country pairings and by the presence of more than one common stochastic trend in the multivariate case.

The significance of the interdependence of the 5 major East Asian SPIs is founded on some conventional financial principles. First, investors in East Asian share markets, such as the mutual funds investors mentioned above, will have less opportunity for risk diversification within the East Asian group, the greater is the degree of interdependence among these markets. Interdependence will also influence the nature of the regional response to common shocks affecting share markets in the region, an important consequence for the debate about contagion caused in this fashion. Further, share market interdependence may be manifest in the presence of causal effects within this East Asian group of share markets leading to the inference that the individual markets are inefficient. In summary, economic theory suggests cointegration or causal links are unlikely to be observed in efficient markets.

The methodological basis for the study is derived from the review of the literature appearing as Section 2, while the data set and methodology are detailed in Section 3. The results of the study are discussed in Section 4 and conclusions drawn in a closing Section.

## 2. Literature Review

The literature relating to stock market integration has its origins in the benefits derived from international portfolio diversification. This is established in the earliest, mean-variance analyses of international share price integration by Greubel (1968), Levy and Sarnat (1970), Solnick (1988) and Lessard (1973) and in the extended studies by Ripley (1973), Panton, Lessig and Joy (1976), Eun and Beswick (1984). Some of these analyses depend on correlation methods which are subject to the limitations identified by Forbes and Rigobon (1998), in particular, that heteroskedasticity of stock market returns produces an upward bias in the correlation of these indices. Correlation analysis can only represent a useful preliminary methodology as a consequence. Granger and Mortensen (1970) applied frequency domain techniques (spectral analysis) to assess the coherence of the power spectra in individual share price time series, but produce no evidence of international price integration among the major developed economies share markets. These spectral results are challenged by Hilliard (1979) using the same multivariate spectral techniques. Hiliard supports the consensus, that there is evidence of some international share price integration in the seventies and eighties among developed nations.

The cointegration revolution has provided further evidence supporting the presence of interdependence among international share prices. Studies by Kasa (1992), Corhay, et.al. (1993), Blackman et.al. (1994), Chung and Liu (1994) and Masih and Masih (1997 a,b) find some evidence of cointegration among US, European and Asian share price indices or returns. Error correction modelling and tests for short run Granger causality identify lags and leads between international share prices in studies conducted by Cheung and Mak (1992), Mallaris and Urrutia (1992), Arshanapalli and Doukas (1993), Brocato (1994) and Masih and Masih (1997 a,b) again are studies finding significant long and short run causality between share price indices.

The argument that the presence of share market interdependence implies market inefficiency is analysed by Richards (1995) using 16 individual share price indices. Richards finds little evidence of cointegration, but does find evidence of a permanent, common world component in the movement of each index. However, there is also evidence of a transitory, country specific mean reverting component of each index which accounts for differences in each series. If Richards results hold universally, then it is unlikely that a single common trend will be found. Kwan et al (1995) find that eight of 36 bivariate studies based on nine individual indices (USA, UK, Japan, West Germany, Australia, Taiwan, South Korea, Singapore and Hong Kong) are cointegrated presenting “mild evidence” for the rejection of a pairwise stock market efficiency hypothesis.

In a similar fashion, some more recent literature about North American and European experiences suggests increasing stock market interdependence. Chelley-Steeley and Steeley (1999) attribute greater integration in Europe to the removal of exchange controls while Asimakopoulos et al (2000) apply a multivariate spectral analysis to determine the coherence in the movement of the US Standard and Poor’s index, the FTSE 100, Dax 30 and CDE 40 indices. The three European series cohere strongly while the US tends to lead the European series at lower frequencies. This evidence for growing stock market interdependence in Europe and North America is not universal. Kanas (1998), for example finds that the US and six major European indices are not cointegrated using a monthly time series spanning the period January 1983 to 1996. However, on balance, recent North American and European studies reveal increasing stock market independence. Some researchers have attempted to explain this outcome in terms of common fundamentals. Bracker, Docking and Koth (1999) apply a panel of nine equity markets over twenty two years to model the co-movement of returns to find that interest rate differentials, bilateral trade links, market size and geographic distance influence the observed co-movement. Similarly, Dickson (2000) analyses the co-movement of US and three leading European series and find a connection between real interest rates and the interdependence of US and European SPIs.

In an Asian regional context, Kwan et al (1995) examines the relationship between nine equity markets (Australia, HK, Japan, Singapore, South Korea, Taiwan, UK, USA and West Germany) for monthly data – January 1982 to February 1991. These researchers find that these share markets are not weak form efficient, individually and collectively. Hung and Cheng (1995) study the interdependence of five East Asian markets on weekly data for two sub periods. These authors find some evidence of cointegration in the second of these periods and only limited opportunities for diversification. Roca (1998) investigates the extent of SPI linkages among five

ASEAN countries using weekly data covering the period 1988 to 1995. No cointegration is evident, however, Granger tests establish causality links and also indicate that impulses flowing from one index to another impact immediately and extend in effect beyond two months. Roca (1999) and Roca and Selvanathan (2001) find no SPI linkage between Australia and Japan, HK, Singapore Taiwan and Korea, using weekly data over the period 1974 to 1995, although Australian share prices respond to US and UK shocks. Cha and Oh (2000) find that the emerging Asian markets (HK, Taiwan, Korea, Singapore) have strengthening links with US and Japanese markets. This conclusion is supported by Siklos and Ng (2001) who find that links between emerging markets and developed world share prices are a recent phenomenon.

This review of the recent literature about share market integration reveals the importance of selecting a potentially coherent sample of indices. In some studies, little if any explanation is offered for the selection of indices included. In the present study, the reasons for selecting the five East Asian nation SPIs is provided in the introduction. A second characteristic of the recent literature is its dependence on smooth cointegration techniques. It is appropriate to apply some recent techniques which allow for structural breaks in cointegrating vectors given the occurrence of recent events such as the Asian currency crisis in 1997. This has not been applied to the East Asian group of markets identified above.

## 3. Data and Research Design

### *3.1 Data*

The data set is developed from a daily time series of the following share price indices: the Singapore Strait Times (SST), Korea composite (KC), Taiwan Weight (TW), Hong Kong’s Hang Seng (HS) and Japan’s Nikkei (JN) index. The daily closing value of each index for the period January 8th 1990 to July 6th 2000 is sourced from the Datastream data base. The five East Asian SPIs comprising the data base are denominated in local currencies and are not converted to a common currency such as the US dollar. This warrants a brief explanation. The preference for local currency denomination of individual share price indices is governed by the objectives of the study which are focussed on domestic causes of share price interdependence. By converting these indices to a common currency there is the possibility that the impact of local economic conditions and domestic economic policy maybe distorted. This is particularly relevant if the spot exchange rate used to convert to common currency is also influenced by local conditions and policy[3].

The argument for conversion is based on quite valid reasoning: foreign investors base foreign investment decisions on returns denominated in their own currencies: Japanese investors think in terms of yen denominated returns, Americans in US dollar returns and European in Euro denominated returns. The difficulty with converting local SPIs to a common currency leaves the following question unresolved: Which currency should be selected as the common one? The answer is not as clear as it was in the eighties and nineties when a stronger case could be made for the US dollar as a world common currency benchmark. This problem is eschewed by observing the local currency value of the five SPIs included in this study.

The intuition for time period 1990 to 2000 is that it begins early enough to reflect the effects of the 1990-92 global recession and the bursting of the Japanese bubble. It also encompasses the occurrence of the Asian crisis and the recovery phase. The raw daily share price indexes are expressed in natural logarithmic form, so that share price links can be expressed in terms of elasticities. The outcome of this data transformation is to produce a time series comprised of 2,739 observations for each of the five indices.

Certain stylised facts about the dynamic behaviour of each index are evident on the graphs of the untransformed share price indices enclosed as charts 1 to 5. The following behaviour patterns are noted. The SST (Chart 1) is comparatively stable during the 1990-1992 recession hovering around the 1250 mark throughout. The KC (Chart 2) falls from its starting value of 900 to its lowest value of 425 on May 7 1992, however, this fall in the KC index does not match the crash of the TW

[3] By domestic interest, growth and inflation rates relative to their foreign counterparts, or by domestically held exchange rate expectations.

(Chart 3) which tumbles from a peak of 12,000 to 2,600 on August 8 1990. Japan’s Nikkei (JN) reflects the impact of the bubble economy’s demise falling from 31,500 in October 1990 to a minimum of 15,000. Thereafter, the Nikkei traded in the range 15,000 to 22,000. The 1990-92 recession does not impact on HK’s Hang Seng (Chart 4), which maintains a strong upward trend throughout this period.

The Asian currency crisis influences these four NIC share markets at similar times. The SST, KC, TW and HS indices feel the chill of the crisis in August 1997 and bottom out in the period June to October 1998. The loss in the value of each index in this period is also similar, the SST losing 122 percent of its value between peak and trough, the TW, KC and HS lost 83, 167 and 138 percent of their value respectively. It is noteworthy also that the four NIC indices have recovered to at least the same level in Taiwan and to levels above the pre Asian crisis peak in Singapore, Korea and Hong Kong. However, the crisis did not push the post bubble value of the Nikkei off course. It continued to trade in the range 14,500 to 20,000 in the post crisis era.

Empirical tests for cointegration can only proceed if tests of the time series involved are I(1). To test for this property, the ADF (Augmented Dickey Fuller) and PP (Phillip Perron) tests are conducted on the level and first-differenced stock index series for the whole sample. The lag order required for the ADF and PP was automatically selected by SHAZAM. The results of the tests conducted on the stock indices levels and first-differences are reported in Table 1.

**Table 1 here**

There is striking evidence of the non-stationarity of the time series in levels, which means a single hypothesis of a unit root and joint hypothesis of trend and unit root cannot be rejected at the 10% significance level. The only exceptions are the TW and the JN stock price series. In relation to TW, both the ADF test and PP test statistics with constant and trend are more than their critical values indicating that we can reject the null hypothesis. However, both the ADF test and PP test statistics “with constant but no trend” indicate that both the joint and the single hypotheses cannot be rejected at the 10% level. In the case of JN, all the tests indicate stationarity in levels except for the test statistics of ADF with constant and trend. In light of this contradictory evidence, we will proceed assuming non-stationarity of all time series in levels. Similar tests are then conducted on the first differenced stock indices series. All tests reject the null hypothesis of a unit root at the 10% significance level. Each index can be treated as an I(1) series and tests for cointegration can be undertaken.

### *3.2 Research Design*

The following formal analysis contains four procedures: the first involves a simple correlation test of the strength of association of the selected share price indices; the second involves both a bivariate and multivariate cointegration analysis followed by the third, error correction modelling and finally, if appropriate, Granger tests for short run causality. These procedures are required to satisfy various facets of our research goals.

The correlation matrix involving correlation coefficients for pairs of stock prices is estimated for two sub samples which are described as the pre Asian crisis period dating from January 8 1990 to July 1 1997 and the post Asian crisis period dating from July 2 1997 to December 31 1999. The significance or otherwise of the correlation coefficient for each potential pair of share price indices provides a preliminary indication about the strength of association of share price movements in East Asia, before and after, the Asian currency crisis. Further, formal tests for significant differences between equivalent correlation coefficients pre and post Asian crisis will indicate if interdependence has increased in recent times, in particular since the crisis first became evident in Thailand on July 2 1997.

The phrase “preliminary” is used advisedly in describing the findings of the correlation analysis. This caveat is entered because correlation coefficients are known to be biased upward if share price indices are heteroskedastic. So correlation analysis does not provide a sound basis for studies of interdependence. However, correlation analysis provides a commonly used preliminary technique which is applied also to this study. Apart from the issue about the validity of correlation analysis, there is also the restrictive nature of the correlation method in relation to studies of interdependence. It provides little if any insight into the dynamic behaviour of stock market links and no information about causality. The results of the correlation analysis appear on Table 2.

An essential feature of the dynamics of stock market links is to know if different stock price indices follow a common trend and if there is a long run relationship between them. Insights into these issues can be determined from cointegration analyses. Two bivariate cointegration tests are applied. The first is the familiar residual based test proposed by Engle and Granger (EG) (1987). The results are included on Table 3. The EG procedure only provides evidence in favour of time invariant cointegrating vectors between the selected stock indices. There is the prospect that a linear combination of the non stationary stock indices is stationary but the cointegrating vector has shifted at a certain point in the sample period, for example, during the Asian Crisis. In such a case, the conventional test for cointegration may not be appropriate as the power of the conventional tests falls sharply in the presence of a structural break. Gregory-Hansen (GH) (1996) techniques are used as these methods take into account a regime shift in the time series data. At the same time, these methods also allow analysis of the time-series property of the data without prior knowledge of the date of a shift. To illustrate Gregory and Hansen's (GH) (1996) method of testing for cointegration subject to structural breaks, consider the model involving two share price indices $S_t^D$ and $S_t^I$ where $S^D$ and $S^I$ refer to the dependent and independent country series respectively.

$$S_t^{\ D} = \mu + aS_t^{\ I} + e_t \qquad (1)$$

The model captures a long run relationship, thus $\mu$ and $\alpha$ must be constant over the test period. To model for structural change, the timing of this shift is treated as unknown so it is useful to define the dummy variable $\Phi_{1\tau}$ such that:

$$\Phi_{1\tau} \qquad \begin{cases} 0 \ \text{ if } \ t \le [nt] \\ 0 \ \text{ if } \ t > [nt] \end{cases}$$

Where the unknown parameter $t \in (0,1)$ denotes the timing of the change point and [nτ] denotes the absolute integer part with n = 2739, the number of observations. The current study investigates one model of structural change, namely a "level shift with trend". The level shift model can be modelled as a change in the intercept $\mu$ while the slope coefficients $\alpha$ are held constant implying that the equilibrium equation has shifted in a parallel fashion. Consider the second model:

$$S_t^{\,D} = \mu_1 + \mu_2 \varphi_{1t\tau} + \alpha S_t^{\,I} + e_t \qquad (2)$$

t = 1,…,n

where $\mu_1$ represents the intercept before the shift and $\mu_2$ represents the change in the intercept at the time of the shift. In addition, a time trend can be introduced into the level shift model as follows:

$$S_t^{\,D} = \mu_1 + \mu_2 \varphi_{1t\tau} + \beta t + \alpha_1 S_t^{\,I} + e_t \qquad (3)$$

t = 1,…,n

The regression of equation (3) is estimated for each potential break point in each pair of stock indices examined. The cointegration test statistic is computed for several possible regime shifts where $t \varepsilon T$. The set T can be any compact subset of (0, 1) but following GH, $t$ is set over the range (0.15, 0.85). The GH test for the cointegration of $S_t^{\,D}$ and $S_t^{\,I}$ is conducted by applying modifications of the ADF and PP test statistics which are standard tools for the analysis of cointegrating regressions without regime shifts. The modification is that the smallest values of the ADF and PP test statistics are applied from the estimates of (3) at each $t$. The new test statistics are denoted as follows:

$$ADF' = \inf\left(t \in ?\right) ADF(t) \qquad (4)$$

$$PP' = \inf\left(te?\right) Z_t(t) \qquad (5)$$

The smallest values of the $ADF'$ and $PP'$ statistic for all $t$ apply here because the more negative are the $ADF(t)$ and $Z_t(t)$, the greater is the probability of rejecting the null hypothesis of no cointegration with a structural break. The critical values Crt (n, p, m) for up to four regressors are calculated by Gregory and Hansen (1996) using simulation methods. The choice of critical value is dependent upon the number of observations in the sample; n, the percent quantile; p and the number of regressors; m which exclude constant and trend.

The GH technique does not require potential breakpoints to be predetermined, a characteristic evident in the correlation analysis where July 2 1997 has been chosen as the likely breakpoint. However, there is no guarantee that breaks in the relationship between these East Asian share price indices occur during the Asian currency crisis. Shifts could occur at any point, so the GH technique will add to knowledge about the behaviour of share price indices in East Asia by identifying the timing of breaks. The results of this GH analysis, particularly, the relevant breakpoints appears on Table 3. Standard multivariate cointegration tests for common trends among the five stock market indices complete the battery of multivariate cointegration tests. The results are included in Table 5.

Causality is an issue not addressed adequately by correlation tests. Error correction modelling provides the most comprehensive approach to this issue. In the presence of cointegration, there always exists a corresponding error-correction representation. This implies that changes in the dependent stock indices are a function of the level of disequilibrium in the cointegrating relationship captured by the error-correction term as well as changes in other independent or explanatory stock indices. In other words, there may exist co-movements between pairs of time series stock indices and the possibility that they will trend together in establishing a long-run stable equilibrium. If we exploit this idea using the Granger representation theorem we may posit a test relationship which constitutes a vector error correction model (VECM).

For the bivariate case, consider a pair of variables which are I(0) after applying the differencing filter once. Provided that the variables are also cointegrated of order r, we may impose this constraint upon our unrestricted VAR to enable a VECM formulation as follows:

$$\Delta S_t^{\ D} = \xi + \sum_{i=1}^{k1} \eta_I \Delta S_{t-i}^{\ D} + \sum_{i=1}^{k2} \lambda_i \Delta S_{t-i}^{\ I} + \gamma ECT_{t-1} + v_t \qquad (6)$$

Where the ECT is the residual from the long run cointegrating regression and $\gamma$ is the impulse which represents the unanticipated movement in $S_t^D$. From this model, two channels of causation may be noted. The first one is the standard Granger test, which evaluates the significance of the coefficient, $\lambda_i$ of the lagged independent variable. The null hypothesis of no causation from $S_{t-i}^{I}$ to $S_{t-i}^{D}$ can be tested using standard F tests as mentioned above. Through the ECT, the ECM opens up the second channel for Granger-causality which is exposed through the statistical significance of the lagged ECTs $(\gamma's)$ by a separate t-test. The non-significance of both the t and F tests in the VECM indicates econometric exogeneity of the dependent stock indices.

Cointegration and error-correction models hold several intuitive implications. When the pair of stock indices are cointegrated, then in the short-term, deviations from this long-run equilibrium will feed back on changes in the dependent stock index in order to force its movement towards a long-run equilibrium. Should the dependent stock index be driven directly by this long-run equilibrium I(0) error, then it is responding to this feedback otherwise it is responding only to short-term shocks to the stochastic environment. The F-tests of the differenced independent stock index give us a sign of the short-term causal effects. On the other hand, the long run relationship is implied through the significance of the t test of the lagged error-correction term, which contains the long-term information as it is derived from the long run cointegrating relationship. However, the coefficient of the lagged error-correction term $\gamma$, is a short-term adjustment coefficient and represents the proportion by which the long-run disequilibrium in the dependent stock indices is being corrected in each short period. Therefore, the non-significance or elimination of any of the lagged error-correction terms affects the implied long-run relationship and may be a violation of theory. Nevertheless, the non-significance of any of the independent stock indices which reflects only short-run relationships does not involve such violation because theory normally has little to say about a short-term relationship. The results of this bivariate VECM analysis are included on Table 4 and indicate the nature of lead and lag relationships between pairs of stock market indices.

The multivariate case is also of importance in identifying lead, lag relationships between the five stock market indices. The multivariate VECM formulation is similar to the VECM in a bivariate context except that it includes all the five markets in each equation. In the case, five multivariate VECMs will be examined as follows:

$$\Delta X_t = \sum_{i=1}^{n} A_i \Delta X_{t-i} + \sum_{i=1}^{r} \xi_i \phi_{t-i} + v_t \tag{7}$$

where $X_t$ is an n x 1 vector of variables (stock indices), A's are estimable parameters including the constant term, $\Delta$ is a difference operator, $\xi_t$ is a vector of impulses representing the unanticipated movements in $X_t$ and $\phi$ contains the r individual error correction terms derived from the r long run cointegrating vectors.

## 4. Results and Interpretation

The following questions are addressed in the interpretation of the results of this analysis:

- Are the five time series stationary?
- How closely associated are the East Asian SPIs
- Are there long run relationships in a bivariate and multivariate context?
- Are there structural breaks in bivariate relationships?
- Are there causal relationships between East Asian SPI's.

These questions are addressed in sequence by referring to the results given by the application of specific procedures[4], noting that the I(1) property of each time series is established in the discussion of the data set.

### *4.1 Correlation: How Closely Related are the Selected East Asian SPI Before and After the Asian Currency Crisis?*

Table 2 shows the estimated correlation matrix for the 5 East Asian SPIs pre and post crisis and also the value of Anderson's (1958) Z statistic used as a basis to test for significant differences between these correlation coefficients.

**Table 2 here**

In general, the correlation between the selected indices is higher following the Asian crisis. This is supported by Anderson's (1958) tests for significant differences between correlation

[4] All the results obtained result from estimation using SHAZAM 7.0 edited for the GH procedure and to increase data dimensionality.

coefficients. The Z statistics appearing on Table 1(c) are all significant at the 0.01 level of significance with the exception of the HS and SST indices where Z is significant at the 10 percent level only. A feature of these results is the increase in the correlation between Japan's Nikkei and the four NIC indices in more recent times. In summary, it can be argued that links between the five share markets have strengthened since the occurrence of the Asian crisis, but this conclusion remains tentative given the explained limitations of correlation analysis. A deeper understanding depends on the application of more sophisticated techniques. Our discussion of these follows.

### *4.2 Are There Long Run Relationships in a Bivariate Context?*

The standard Engle Granger (EG) (1987) procedure is applied to the relationship between a dependent stock price series labelled $S_{it}^{D}$ and $S_{jt}^{I}$ a single independent series: $i \neq j$. Following EG, the OLS equation takes the form:

$$\Delta S_{it}^{D} = \alpha + \beta \Delta S_{jt}^{I} + \varepsilon_t \tag{8}$$

where $\Delta S_{jt}$ is I(1) and $\varepsilon_t$ is I(0).

The outcomes of ADF and PP tests for a standard EG model, namely, a model with constant and trend are contained in columns (2) and (3) of Table 3.

**Table 3 here**

When the cointegrating vector is smooth, there are only ten possible long run equilibrium relationships between the selected East Asian share price indices. From columns (2) and (3) of Table 2 Korea composite (KC) and Singapore Straits Times (SST) indices are cointegrated at the 10 percent level of significance. This outcome applies in both directions from SST to KC and from KC to SST. There is also evidence of a long run relationship between the Hang Seng (HS) and Taiwan weighted (TW) indices again in both directions at the 10% level. An identical interpretation applies to the SST as the independent variable and the TW index. The KC – TW indices have a long run relationship while Japan's Nikkei has a long run co-movement with the KC, TW, SST and HS indices according to the PP but not the ADF test. Our conclusion is that long run relationships can exist in only half the bivariate studies reported on columns (2) and (3) of Table 3. However, there is evidence of a long run relationship in all ten cases if the directional issue is put to one side. The results for SST/KC, SST/TW, KC/TW and HS/TW are stronger in this respect being supported by both Phillips Peron and ADF tests. However, the PP tests alone suggest a long run equilibrium for the Nikkei with the SST, KC and HS.

### *4.3 Are there Breaks in Bivariate Relationships? – Comment*

It is possible that the null hypothesis of no cointegration is rejected in the smooth (no break) case when cointegration is present provided structural change is evident. The results of GH tests for each of twenty bivariate comparisons are contained in columns (4), (5), (6), (7) of Table 3. These outcomes do not add to the list of bivariate pairs, which are cointegrated in the smooth case. The Korea composite and Singapore Strait Times indices are cointegrated with a break occurring in June 1998. This break point coincides with differing dates for the recovery of each index: the SST on June 19 1998 and the KC on October 5 1998. The JN index and SST display evidence of a break in links between them in January 1992 which coincides with a small rally in the Nikkei while the SST continued through a period of stability. This occurrence may have temporarily altered the pattern of the relationship between these. The KC and Nikkei link breaks in January-February 1996 when the Nikkei was in the midst of an upswing and the KC was in a trough. These contrary movements coincide with the presence of a break. The final breakpoint evident on Table 4 occurs in January 1992 and involves the HS and JN indices. This break occurs as the HS accelerates briefly in an upswing phase of the business cycle possibly explaining a break in the HS→JN relationship.

### *4.4 Are there Causal Relationships in a Bivariate Context?*

Table 4 summarises the F stats and t ratios in the ten instances were bivariate cointegration reveals the existence of a long run bivariate equilibrium. In developing this table, a distinction between these bivariate studies in which cointegration was evident with and without break (no break on Table 4) is drawn. In the no break case the error correction term is significant at the 5% level in all five cases suggesting that the TW SPI bears the burden of adjusting back to long run equilibrium with each of the four indices SST, KC, HS and Japan's Nikkei (JN). Further, the JN adjusts to any long run deviation from its equilibrium with the HS index. The significance of the F stat on lagged values of $\Delta S^{I}_{t-i}$ suggests that short run causality also runs from the HS and JN indices to the TW index at the 5% level and from the HS to the JN index. The SST also leads the TW at 10% while there is no causal link from the KC to the TW index.

Table 4 also reports the results of estimating the ECM model (6) with trends, when these structural breaks identified above are accommodated under the heading "pre and post break". In the pre break eras, the significance of relevant ECTs suggests that the following adjustments of deviations from long run equilibria take place: the SST to its equilibrium with the KC, the JN to equilibrium with the KC and the JN to equilibrium with the HS. In relevant post break eras the following long adjustments occur: the KC to the SST, the JN to the SST otherwise the pattern of long run adjustment is identical to pre break patterns.

There is universal evidence of short run as distinct from long run adjustments in the pre break era, all five F-stats on lagged values of $\Delta S^{I}_{t-i}$ in (6) are significant at the 5% level. A similar result holds in post break periods, although there is no short run adjustment in one case, namely, the relationship between SST and KC indices. The directions of causation are indicated by the arrows in the "combination" column of Table 4.

The results disclosed in Table 4 relate to studies in which there is some evidence for the presence of pairwise long run equilibria. However, the absence of cointegration does not preempt the presence of short run causal links in the Granger sense. These results are not tabulated[5]. The results of standard Granger tests of causality for these non cointegrated pairs of indices reveal that

[5] The F-stat for causality tests are as follows for a preferred lag of 2 periods: ST $\rightarrow$ HS (3.12), HS $\rightarrow$ ST (12.59), KC $\rightarrow$ HS (5.19), HS $\rightarrow$ KC (9.80) and JN $\rightarrow$ KC (3.42). These bracketed scores all exceed the 5% cut off score of 3.00 from the F-distribution.

causality is bi-directional in the cases of the HS and SST indices, and the KC and HS indices, while the JN Granger causes the KC. There is no evidence of short run lag or lead relationships in the following bivariate studies: the TW and the SST, the JN and SST, the TW and the KC, the TW and the HS and the JN and HS indices.

**Table 4 here**

### *4.5 Is there a Causal Direction Relationship in a Multivariate Context?*

Ten pairs of stock indices are not cointegrated in both the bivariate conventional cointegration tests and the bivariate modified cointegration with breaks; this applies in both directions for ST/HS and KC/HS and in one direction for ST→TW, ST→JN, KC→JN, HS→TW and HS→JN. However, our failure to find significant bivariate cointegration does not mean necessarily the absence of no cointegration. There is the possibility of a multivariate common trend. JJ tests for simultaneous cointegration are conducted. Results of cointegration ranked by the JJ procedure appear in Table 5. For completeness, two statistics, the trace and maximal eigenvalue are calculated to test for the presence of r cointegrating vectors. Note that, Cheung and Lai (1993) favour the trace test because it shows little bias in the presence of either skewness of excess kurtosis and is found to be more robust to both skewness and kurtosis than the maximal eigenvalue test. Evidence from both trace and maximal eigenvalue tests suggests that there is at most a single cointegrating vector. This indicates a degree of interdependence among the five East Asian SPIs although there are at least four common stochastic trends and as a consequence, there is a problem of identification. In this context, there is no evidence supporting the existence of a single common trend which may suggest perfect integration of the East Asian SPIs. The findings from these multivariate study confirm the results of the bivariate analysis where exactly 50 percent of the bivariate studies reveal the existence of long run equilibria.

**Table 5 here**

Given the presence of a unique cointegrating vector in the five-dimensional VAR used in the Johansen and Juselius (1990) cointegration tests, this then provides us with one error-correction term for constructing VEC models specified by equation (7). The first channel, the standard Granger-causality tests can be examined by evaluating the joint significance of the sum of the lags of each explanatory stock index in turn for equation (7) using joint F-tests. The second channel of causation which is represented by the ECT can be evaluated by separate t-test on $\zeta_1, \zeta_2, \zeta_3, \zeta_4$ and $\zeta_5$ in (7). The sequential F test is applied in order to determine the appropriate number of lags in the multivariate VECM. Two lags are indicated. The results of multivariate VECM estimation are reported on Table 6.

**Table 6 here**

According to Table 6, at least one channel of Granger causality is active: either the short-run one through joint tests of lagged-differences or a statistically significant ECT. The latter channel is a novelty of the VECM formulation but it is significant only in the equations for SST, TW and JN. The economic intuition arising from this finding implies that when there is a deviation from equilibrium cointegrating relationships as measured by the ECTs, it is mainly changes in these three markets that adjust to clear the disequilibrium.

Although the ECT is only statistically significant in three equations, one cannot assume that all other markets are non-causal since the short-run channels are still active. For example, fluctuations in the South Korean market seem to explain movements in all the regional markets apart from the Taiwan market. This intuitively implies that the Korean stock market is a potential leading stock market in spreading disturbances to these regional stock markets. The Singapore stock market also seems to play a leading role as it causally affects two national stock markets in the region, namely those in Korea and Taiwan. These findings are consistent with the bivariate case where KC and ST are the stock indices, which have causal effect (either short term or long term or both) on the national stock markets. Interestingly, although Japan and Taiwan stock markets are massive and influential in a global context, they do not seem to have short-term causal effects to any of the regional markets included here.

From Table 6, all the stock markets are significantly Granger caused by only one foreign market with the exception of Singapore which is simultaneously caused by two markets, namely Korea and Hong Kong. One can also observe that Singapore and Korea have a bi directional short-term dynamic causal relationship. Overall, one can observe that there are linkages between Singapore, South Korea and Hong Kong stock markets. Our interpretation of interdependence here focuses on economic similarities or common factors. In particular, Singapore, South Korea, and Hong Kong were fast growing economies and these East Asian "miracle" countries shared a export-oriented development strategy which is increasingly reliant on the markets of South East Asia (the ASEAN countries).

## 5. Conclusion

The premier share price indices of the East Asian region, namely, Japan's Nikkei (JN), Singapore Straits Times (SST), Korea composite (KC), Hang Seng (HS) and the Taiwan weighed (TW) exhibit a degree of interdependence which runs counter to the market segmentation hypothesis and market efficiency, thereby limiting the opportunities for the diversification of risk.

Collectively, the evidence from this study sums to the following: the correlation of these five indices strengthens following the Asian crisis beginning in June 1997. An increasing correlation between the JN and the four NICs is particularly evident post June 1997. In ten of the twenty potential pairings of East Asian indices, there is some evidence of bivariate cointegration. In five instances a pairwise long run equilibrium occurs without structural breaks in the cointegrating vector and in the five remaining cointegration subject to a structural break is evident. This bivariate evidence is supported by multivariate tests which reveal at most one cointegrating and at least four common stochastic trends which might be followed by a combination of the five SPIs. Both long and short run causal links between the five series are evident and in those cases where cointegration is not evident, there is still some evidence of short run Granger causality. The evidence suggested by bivariate error correction modelling is supported by the results of vector error correction modelling in the multivariate case. At the completion of this study, there are only five of twenty pairings of these East Asian SPIs which appear to be unrelated either in terms of the presence of a long run equilibrium or in terms of long and short lags and leads. In conclusion, the evidence for the interdependence of the five developed country East Asian share price indices appears to be quite strong.

Further research is suggested by this analysis. This should involve a formal study of common factors influencing share price movements in the East Asian region and simultaneous modelling of share prices and other financial variables.

## Table 1: Stationarity Tests Levels and First Differences

| Stock Indices | Levels | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | ADF test | | | | PP test | | | |
| | t-stat | | F-stat | | t-stat | | F-stat | |
| | Constant, no trend | Constant, trend | Constant, no trend | Constant, trend | Constant, No trend | Constant, trend | Constant, no trend | Constant, trend |
| ST | -1.8090 | -2.3908 | 1.7802 | 2.0108 | -1.6367 | -2.1496 | 1.5140 | 1.6614 |
| KC | -2.3652 | -2.3372 | 2.7972 | 1.8996 | -2.1903 | -2.1455 | 2.4023 | 1.6660 |
| TW | -2.5057 | -4.8862* | 3.1661 | 8.5834* | -2.3429 | -3.6735* | 2.7548 | 4.9332* |
| HS | -1.4915 | -2.308 | 2.7512 | 2.8548 | -1.4967 | -2.2195 | 2.9545 | 2.9274 |
| N | -2.8621* | -2.9598 | 4.4813* | 3.5084 | -3.4516* | -3.5203* | 6.5268* | 5.0203* |

| Stock Indices | First Differences | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | ADF test | | | | PP test | | | |
| | t-stat | | F-stat | | t-stat | | F-stat | |
| | Constant, no trend | Constant, trend | Constant, no trend | Constant, trend | Constant, No trend | Constant, trend | Constant, no trend | Constant, trend |
| ST | -6.6612* | -6.6614* | 22.186* | 14.795* | -46.676* | -46.667* | 1089.3* | 725.98* |
| KC | -6.3558* | -6.3724* | 20.205* | 13.544* | -48.764* | -48.748* | 1189.0* | 792.18* |
| TW | -7.0994* | -7.2096* | 25.201* | 17.336* | -50.803* | -50.692* | 1290.6* | 856.67* |
| HS | -6.7713* | -6.7891* | 22.926* | 15.372* | -50.693* | -50.683* | 1284.9* | 856.27* |
| N | -7.5076* | -7.5776* | 28.188* | 19.148* | -53.276* | -53.330* | 1419.1* | 947.98* |

The critical values for t-test and F-test with constant, no trend at 10% level are –2.57 and 3.78 respectively.
The critical values for t-test and F-test with constant. Trend at 10% level are –3.13 and 4.03 respectively.
*indicates significance at 10% level.

**Table 2: Correlation Matrices**

(a) Correlation Matrix of stock indices for non-crisis period: 08/01/1990 – 01/07/1997

| | | | | | |
|---|---|---|---|---|---|
| SST | 1.00000 | | | | |
| KC | 0.68763 | 1.00000 | | | |
| TW | 0.46524 | 0.45631 | 1.00000 | | |
| HS | .9166 | 0.46704 | 0.24175 | 1.00000 | |
| JN | -0.37626 | -0.03569 | 0.32283 | -0.62683 | 1.00000 |
| | ST | KC | TW | HS | JN |

(b) Correlation Matrix of stock indices for crisis period: 02/07/1997 – 31/12/1998\

| | | | | | |
|---|---|---|---|---|---|
| SST | 1.00000 | | | | |
| KC | 0.90408 | 1.00000 | | | |
| TW | 0.73854 | 0.66784 | 1.00000 | | |
| HS | 0.94530 | 0.90328 | 0.76406 | 1.00000 | |
| JN | 0.75616 | 0.74513 | 0.86825 | 0.77441 | 1.00000 |
| | ST | KC | TW | HS | JN |

(c) Results for Anderson's Z-test for significance of changes in correlation coefficients

| | | | | | |
|---|---|---|---|---|---|
| SST | N/A | | | | |
| KC | 5.08869*** | N/A | | | |
| TW | 3.46675*** | 2.45719** | N/A | | |
| HS | 1.70735* | 7.69144*** | 5.93795 | N/A | |
| JN | 10.81503*** | 7.80204*** | 7.75122*** | 13.82244*** | N/A |
| | ST | KC | TW | HS | JN |

The critical values at 1%, 5% and 10% significance level are 2.575, 1.960 and 1.645 respectively.
*indicates significance at 10% level.
**indicates significance at 5% level.
***indicates significance at 1% level.
(1) Legend: SST (Singapore Strait Times), KC (Korea Composite), TW (Taiwan Weighted), HS (Heng Seng), JN (Japan Nikkei).

## Table 3: Bivariate Cointegration with and without Break

| Without Break | | | With Break | | | |
|---|---|---|---|---|---|---|
| ① **S ® $S^D$** | ② **ADF** | ③ **PP** | ④ **ADF$(t)$** | ⑤ **Break** | ⑥ **Z$(t)$** | ⑦ **Break** |
| KC→SST | -3.92* | -4.41* | -5.43* | June 98 | -63.64** | June 98 |
| SST→KC | -3.85* | -4.41* | 05.33* | “ | -60.46** | “ |
| TW→SST | -2.31 | -2.06 | -3.37 | n.a | -19.97 | n.a |
| SST→TW | -4.55* | -3.67* | -4.34 | n.a | -37.12 | n.a |
| HS→SST | -2.95 | -2.85 | -3.73 | n.a | -28.96 | n.a |
| SST→HS | -3.10 | -2.96 | -4.48 | n.a | -38.98 | n.a |
| JN→SST | -2.47 | -2.18 | -3.76 | n.a | -26.78 | n.a |
| SST→JN | -2.87 | -3.63* | -5.17** | March 92 | –47.38* | Jan 92 |
| TW→KC | -2.16 | -2.04 | -3.12 | n.a | -18.59 | n.a |
| KC→TW | -4.52* | -3.65* | -4.35 | n.a | -36.56 | n.a |
| HS→KC | -2.57 | -2.52 | -3.63 | n.a | -25.25 | n.a |
| KC→HS | -2.64 | -2.64 | -4.30 | n.a | -36.53 | n.a |
| JN→KC | -2.28 | -2.20 | -4.01 | n.a | -31.83 | n.a |
| KC→JN | -2.93 | -3.61* | -4.82* | Feb 96 | -44.59* | Jan 96 |
| HS→TW | -4.57* | -3.77* | -4.33 | n.a | -37.12 | n.a |
| TW→HS | -2.28 | -2.26 | -4.30 | n.a | -36.87 | n.a |
| JN→TW | -4.58* | -3.61* | -4.32 | n.a | -38.40 | n.a |
| TW→JN | -2.80 | -3.56 | -4.00 | n.a | -31.42 | n.a |
| JN→HS | -2.22 | -2.23 | -4.34 | n.a | -38.19 | n.a |
| HS→JN | -2.89 | -3.55* | -4.99* | Jan 92 | –44.09* | Jan 92 |
| *0.10 critical value for ADF -3.04 and PP = -3.50 | | | *10% critical value ADF $(t)$ = -4.72 and Z$(t)$ = -43.22<br>**5% critical value ADF$(t)$ = -4.99 and Z = -47.96 | | | |

## Table 4: Bivariate ECMs for Cointegrated Series: Results*
## No Breaks[1]

| Combination $\Delta S^P_t - \Delta S^I_{t-i}$ | F Stat of $\Delta^I_{t-i}$ | $ECT_{t-1}$ (t – ratios) |
|---|---|---|
| $\Delta TW \leftarrow \Delta SST$ | 12.009* | -0.004** (-2.841) |
| $\Delta TW \leftarrow \Delta KC$ | 1.369 | -0.003** (-4.042) |
| $\Delta TW \leftarrow \Delta HS$ | 8.142** | -0.005** (-3.293) |
| $\Delta TW \leftarrow \Delta JN$ | 5.058** | -0.003** (-2.095) |
| $\Delta JN \leftarrow \Delta HS$ | 3.566** | -0.011** (-3.903) |

## Pre and Post Break[2]

| Combination Break[3] | | Pre Break | | | Post Break | | |
|---|---|---|---|---|---|---|---|
| $\Delta S^D_t - \Delta S^I_{t-i}$ | | F-stat | $ECT_{t-1}$ (t ratios) | | F-stat | $ECT_{t-1}$ (t ratios) | |
| $\Delta SST \leftarrow \Delta KC$ | June 98 | 9.157** | -0.003** | (-2.583) | 1.155 | -0.007 | (-1.496) |
| $\Delta KC \leftarrow \Delta SST$ | June 98 | 6.438** | -0.002 | (-0.084) | 3.904** | -0.016** | (-2.358) |
| $\Delta JN \leftarrow \Delta SST$ | Jan 92 | 6.319** | -0.007 | (-1.186) | 7.655** | -0.006** | (-3.216) |
| $\Delta JN \leftarrow \Delta KC$ | Jan 96 | 4.566** | -0.005** | (-2.710) | 3.522** | -0.006** | (-2.620) |
| $\Delta JN \leftarrow \Delta HS$ | Jan 92 | 4.827** | -0.021** | (-2.768) | 2.955* | -0.006** | (-3.384) |

Critical Values:
F-test: 3.00 (5.5%), 2.30 (10%).
t-test: 1.96 (5.0%), 1.645 (10%).
** denotes significant at 5% level.
* denotes significant at 10% level.

Notes:
1. These are the five bi-variate studies in which no breaks were evident in cointegration.
2. Where breaks in the cointegrating vector were detected, ECMs were conducted pre and post break to determine if structural change occurs post break.
3. Sourced from Table 3.

**Table 5: Results for Multivariate Cointegration Tests**

| $H_0$ | Johansen-Juselius Test | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | 5% Critical values | | Lags = 2 | | Lags = 4 | | Lags = 6 | |
| | Trace test | Eigen test | Trace test | Eigen test | Trace test | Eigen test | Trace test | Eigen test |
| r = 0 | 70.0 | 33.3 | 88.301* | 40.863* | 87.324* | 40.735* | 86.561* | 40.400* |
| r ≤ 1 | 48.4 | 27.3 | 47.438 | 46.590 | 46.590 | 21.609 | 46.161 | 22.080 |
| r ≤ 2 | 31.3 | 21.3 | 25.556 | 24.980 | 24.980 | 16.618 | 24.080 | 16.710 |
| r ≤ 3 | 17.8 | 14.6 | 8.431 | 8.362 | 8.362 | 5.105 | 7.371 | 4.385 |
| r ≤ 4 | 8.1 | 8.1 | 2.982 | 3.257 | 3.257 | 3.257 | 2.986 | 2.986 |

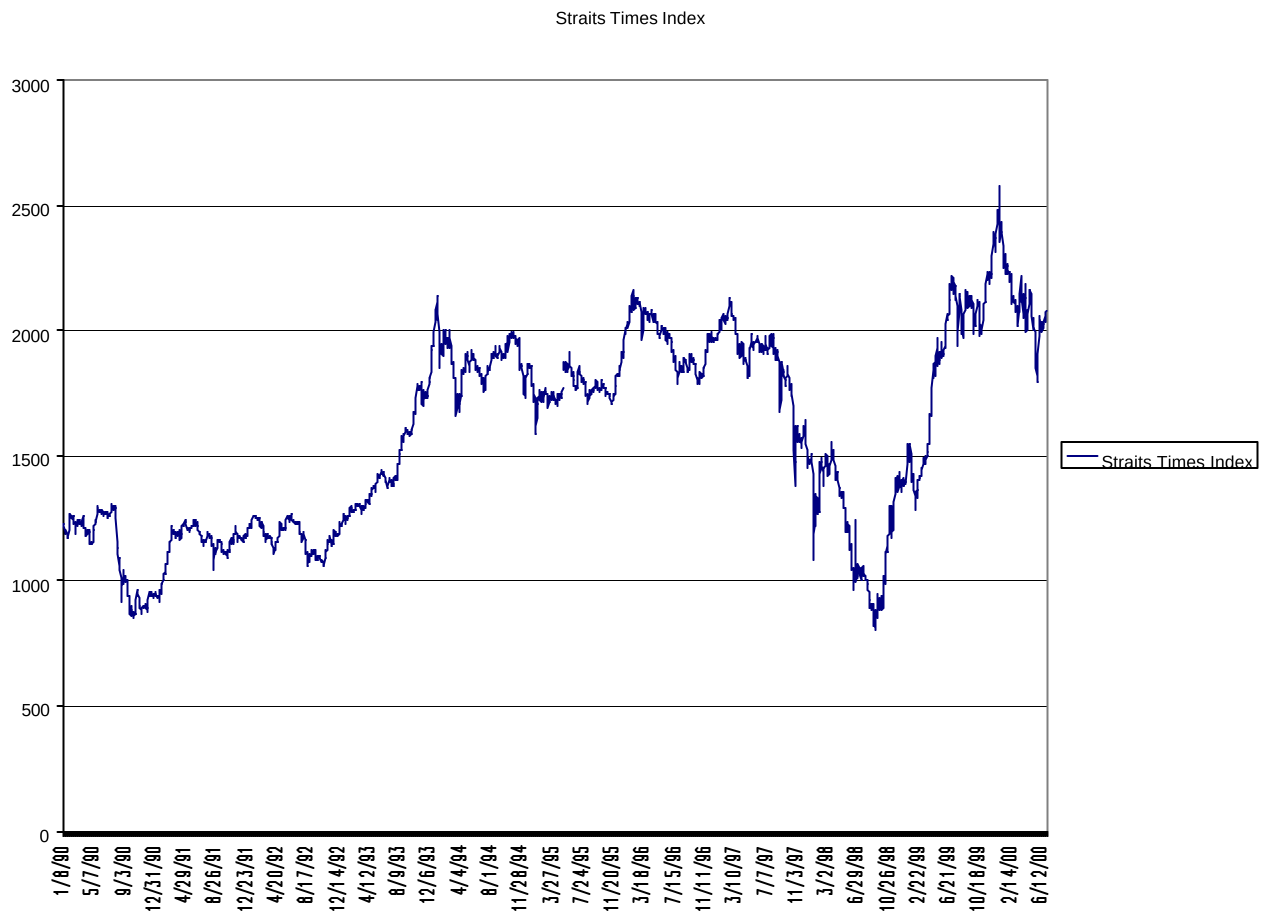
* indicates significant at 5% level

**Table 6: Results of Multivariate Vector Error-Correction Model**

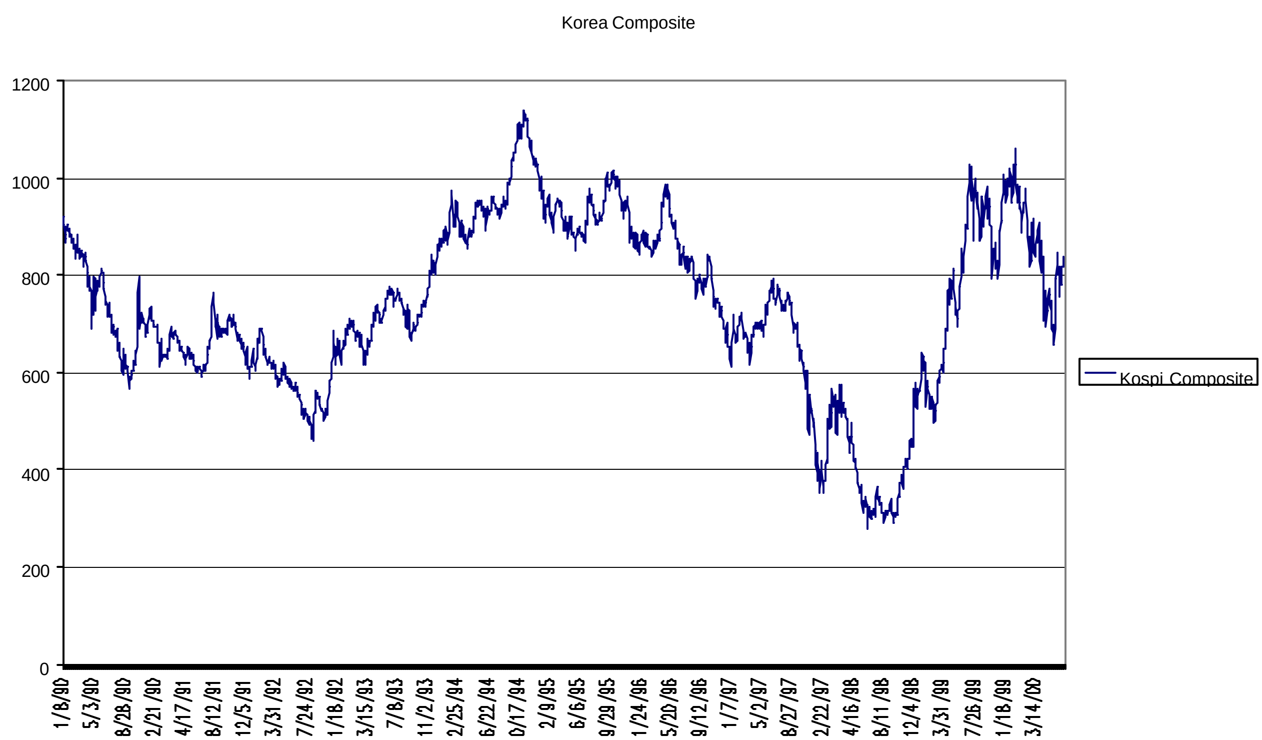
| Dep Var | Short-Run Lagged Differences F Statistics | | | | | Lagged ECT |
|---|---|---|---|---|---|---|
| | DST F-statistics | DKC | DTW | DHS | DN | $ECT_{t-1}$ |
| ΔST | - | 9.0360** | 0.0490 | 8.5863** | 0.3682 | -0.0221** (-4.729) |
| ΔKC | 7.7146** | - | 0.3963 | 0.6300 | 0.2600 | -0.0060 (-1.900) |
| ΔTW | 5.5235** | 0.3340 | - | 0.8934 | 1.7747 | -0.0056** (-3.262) |
| ΔHS | 1.6441 | 5.3769** | 0.2036 | - | 1.9221 | -0.0015 (-0.554) |
| ΔN | 1.5326 | 3.4216** | 0.4446 | 0.4240 | - | -0.0102** (-3.601) |

The numbers in parentheses are t-statistics
The critical values at 5% and 10% significance levels for F tests are 3.00 and 2.30 respectively.
The critical values at 5% and 10% significance levels for t-test are 1.960 and 1.645 respectively.
*indicates 10% significance level.
**indicates 5% significance level.

## Chart 1: Singapore Straits Times




## Chart 2: Korea Composite



## Chart 3: Taiwan Weighted


## Chart 4: Hong Kong Hang Seng

**Chart 5: Japan Nikkei**

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