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How close a relationship does a capital market have with other such markets? The case of Taiwan from the Asian financial crisis☆

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ABSTRACT

In assessing how far and how close the relationships are between the Taiwan capital market and other international capital markets in Asian financial case, we examine the co-movement patterns by developing the "unequal variance test". We find that a closer relationship exists between Taiwan and Hong Kong throughout the sample period than between Taiwan and other Asian countries and the US. It thus appears that adjacent regions with similar backgrounds in terms of their capital markets will reflect price patterns that are more similar to those of Taiwan than those of countries with which Taiwan frequently trades or cooperates.

1. Introduction

Owing to reasons attributed to close interaction in economic trade or the sudden dispersion of international information, capital markets in different countries may frequently affect each other. The past literature has provided extensive discussions and findings related to the issue of linkages among stock markets using the cointegration approach. Similar examples related to the co-movement patterns

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in international capital markets abound in the literature and the reasons for these relationships have been investigated (Malliaris and Urrutia (1992); Chan et al. (1992); Arshanapalli and Doukas (1993); Arshanapalli et al. (1995); Najand (1996); Uri et al. (1996); Wang et al. (2003); Atteberry and Swanson (1997); Ng (2002); and Yang et al. (2003)). What is implied in the cointegration of international capital markets is that there exist cross-country spillover effects and transmissions of shocks.

Some of the empirical results suggest that co-movements in the international stock markets could be observed following the occurrence of a financial crisis. Two famous financial crises have often been mentioned in the past: the October 1987 global stock market crash and the July 1997 Asian financial crisis that emerged in Thailand. The two cases appear to exhibit similar spillover effects. Arshanapalli and Doukas (1993) regarded the state of the stock markets from 1980 to 1990 and the crash of October 1987 as a watershed. By taking into consideration the Dow-Jones Industrial index, Frankfurt's Faz index, the London FT-SE 100 index, Japan's Nikkei index, and Paris's CAC index, they found that there was an increased tendency for co-movement among most international markets after the crash. Arshanapalli et al. (1995) examined the cointegration effect after the crash in October 1987 in spite of no relationship being found among these markets including the US and various Asian countries (Japan, Malaysia, Singapore, Thailand, the Philippines and Hong Kong) before the crash.

Several studies have suggested that a co-movement pattern exists following the Asian financial crisis, for instance, Arshanapalli, Doukas, and Lang (1995). Ng (2002) pointed out that the linkages among the stock market returns of Indonesia, the Philippines and Thailand exhibited a long-term relationship following the Asian financial crisis in 1997 although no such evidence was found before the crisis. Yang et al. (2003) found similar evidence regarding the transmission mechanism in the 1997 Asian financial crisis among stock markets that included those of the United States, Japan, and Asian emerging countries, showing that country-by-country stock market integration tends to change the leader–follower relationships around the time of the financial crisis period.

Although these past studies have yielded common results with regard to the cointegration effects, the evidence regarding the relationships is still somewhat contradictory. Chan, Gup, and Pan (1992) analyzed the degree of integration among securities markets in Hong Kong, South Korea, Singapore, Taiwan, Japan, and the United States and found no close relationships among any of these markets.

Some natural questions that arise are as follows. First, “how close” are the relationships between the stock markets? Although past studies point out whether or not the interconnections may exist by applying cointegration models, little is known about the how close the co-movement relationship is at present. Second, what is the main reason the spillover effect exists in the first place? Is it because of the adjacent region and the similarities of background of the capital markets, or the international trade and the business cooperation that take place among countries? While the closeness of the relationship could be observed, the answers to the factors behind the spillover effects are much clearer. Third, have the relationships changed because of the financial crisis? It is reasonable to consider that the cointegration relationships will change over time.

Regarding the first question, as to “how close” the relationship is between capital markets has become an interesting and important challenge to traditional thinking that is based on forming an international portfolio for investors. In assessing how far and close the relationships are between international capital markets, we examine the co-movement patterns by developing the “unequal variance test”. This paper builds a statistical test referred to as “the unequal variance test” which, as will be seen later, seeks to find which are the closest linkages.

As for the second issue which relates to why spillover effects exist, we specifically focus on the linkages between each pair of stock markets and include Taiwan in each pair, such as East Asia and Taiwan, and the US and Taiwan. The country of most interest to us in this paper is Taiwan. The other countries related to Taiwan include Australia, Hong Kong, Malaysia, the Philippines, Singapore, South Korea, Thailand, and the US. The main reason why we choose Taiwan is that much trade has taken place with Taiwan with countries other than those in the Asian region over the past several decades, for instance with the US. If the crisis, for example, emerges in Asia, the influence of the spillover effect will be slight for the US and, consequently, for Taiwan. Therefore, this paper attempts to use Taiwan as its central focus to examine the long-run relationships among stock markets including the US market and other Asian markets. The links in terms of stock price movements are closer between the Taiwan and US markets owing to their great trading activities, or between Taiwan and other Asian markets owing to their similar backgrounds.

We are able to observe the different linkages with Taiwan and to include the possible reasons for the spillover effect. While "how close" the relationship is can be observed, it is much more clear with regard to the answer to the factors driving the spillover effect.

With regard to the last question and whether the relationships have changed, we compare three different sub-periods, namely, the pre-crisis, crisis, and post-crisis periods, in order to explore the changing impacts of the 1997 Asian financial crisis.

To sum up, we adopt two viewpoints to observe the impacts of the co-movements of international stock markets in this paper. First, the empirical results are compared for three periods, i.e., before, during, and after the Asian crisis based on the unequal variance test that this paper develops to find out how close the relationship is. Second, this paper contributes to the literature by explaining the factors underlying the spillover effect and the co-movement of capital markets.

We find that Taiwan's relationships with Hong Kong and the US are closer than those between Taiwan and other countries for the whole sample period. If different sub-periods are considered, during the 1997 financial crisis, there is no special relationship between Taiwan and other countries although a cointegration relationship does exist between Taiwan and the other countries. Besides, there is much closer relationship between Taiwan and Malaysia after the 1997 financial crisis. It appears that the similarity of background of the capital market will reflect the more similar price pattern after the financial crisis. The findings of this study have implications for a closer relationship between similar and adjacent countries than major trading partners under the cointegration relationship.

2. The methodology

To conduct this study, in what follows Section 2.1 presents the cointegration test, and Section 2.2 develops the "unequal variance test".

2.1. Cointegration test

Granger (1986) and Engle and Granger (1987) indicate that the long-run relationship between two closely-related time series can be characterized by the concept of cointegration. The series are said to be cointegrated if there exists a linear combination of I(0), i.e., there is integral cointegration. Engle and Granger (1987) provide support for the formal concept of integral cointegration. If two non-stationary time series are able to be co-dominated by a linear combination of the two, they will have the same dominant property or they will be generally cointegrated. The linear combination called the cointegrating equation may be interpreted as a long-run equilibrium relationship.

Besides the Engle–Granger methodology, the maximum likelihood estimators discussed by Johansen (1988) and Stock and Watson (1988), as well as the Johansen and Juselius cointegration tests (1990) are also another main way of testing for cointegration. Although there are two important ways of testing for cointegration, the unequal variance test is derived from the Engle–Granger methodology. Therefore, this paper determines whether there exists an equilibrium relationship based on the Engle–Granger methodology.

By supposing that two variables are believed to be integrated of order 1, Engle and Granger (1987) propose a two-step test to determine whether two I(1) variables are cointegrated of order (1, 1). Using a unit root test to pretest each variable to determine its order of integration is the first step in the analysis. If the variables are both I(1), the next step is to estimate the long-run equilibrium relationship that takes the form:

$$\begin{aligned} S_{TW,t} &= \mu + \gamma S_{1t} + \theta t + \nu_t \\ S_{TW,t} &= u + \kappa S_{2t} + \delta t + \varsigma_t, \end{aligned} \tag{1}$$

where $S_{TW,t}$ is the price of the stock market for Taiwan in time t; μ, γ, θ, u, κ, and δ are constant coefficients; $S_{1t}$and S_{2t} are the two related prices of the stock markets of countries 1 and 2 being compared in time t, respectively; and ν_t and ς_t are residual terms of the regression model.

If we do not reject the null hypothesis that the residual terms contain a unit root, we cannot reject the hypothesis that the variables are not cointegrated. In this paper we use the GLS Dickey–Fuller unit root test

(Elliott et al., 1996; hereafter DF–GLS) to examine the stationarity of the residual sequences. The DF–GLS test[1] follows the estimation of the standard ADF test equation using the GLS detrended variables. The critical values are τ_μ including a constant, and τ_τ including both a constant and trend.

2.2. The unequal variance test

The cointegration model seeks to answer the question as to whether a cointegration relationship exists, but does not describe "how" close the relationship is. This paper extends the concept that a cointegration relationship exists between the two series if the residual series of the linear combination of the two series is stationary. By using the residual series of the two series for the whole period, the unequal variance test seeks to examine this stationarity and may lead to a better understanding of the relationship. The proposed testing procedures are as follows:

The model that seeks to compare the variances of the cointegration equilibrium errors is considered in what follows. Let the vectors $x_t=[S_{TW,t},\ S_{1t}]'$and $y_t=[S_{TW,t},\ S_{2t}]'$ be both 2×1 vectors of a I(1) times series process which contains one or more trends:

$$\begin{aligned} x_t &= \Gamma_t^x+\tau^x t,. \\ y_t &= \Gamma_t^y+\tau^y t,\ t=1,\ 2,\cdots,\ T. \end{aligned} \tag{2}$$

where $\tau^x,\ \tau^y$ is the vector of parameters, and $\Gamma_t^x=\sum_{j=1}^{t}\eta_j^x$, $\Gamma_t^y=\sum_{j=1}^{t}\eta_j^y$ are the stochastic trends with $\Gamma_0^x=\Gamma_0^y=0$. Here, we assume that $\eta_t^x=C^x(L)\varepsilon_t^x$, $\eta_t^y=C^y(L)\varepsilon_t^y$ with $C^x(\cdot)$, $C^y(\cdot)$ representing a matrix-valued polynomial, possibly of infinite order, and $\varepsilon_t^x,\ \varepsilon_t^y$ being an i.i.d. zero mean and finite variance matrix random sequence.

By multivariate Beveridge–Nelson decomposition (see for example, Hamilton, 1994, p. 545) we have

$$\begin{aligned} x_t &= \Gamma_t^x+\tau^x t \\ &= \sum_{j=1}^{t} C^x(L)\varepsilon_t^x+\tau^x t \\ &= C^x(1)\sum_{j=1}^{t}\varepsilon_j^x+\tau^x t+\omega_t-\omega_0 \end{aligned}$$

where $C^x(1)=(C_0^x+C_1^x+C_2^x+\ldots)$ and $\omega_t=\sum_{s=0}^{\infty}\xi_s\varepsilon_{t-s}^x$ for $\xi_s=-(C_{s+1}+C_{s+2}+C_{s+3}+\ldots)$, and $\{\xi_s\}_{s=0}^{\infty}$ are absolutely summable. Therefore, ω_t is a stationary process. If the cointregrating vector is α, then

$$\begin{aligned} \alpha' x_t &= \alpha'\Gamma_t^x+\alpha'\tau^x t \\ &= \alpha' C^x(1)\sum_{j=1}^{t}\varepsilon_j^x+\alpha'\tau^x t+\alpha'\omega_t-\alpha'\omega_0 \\ &= \mu+\theta t+\nu_t. \end{aligned}$$

Since $\alpha' x_t\sim\mathrm{I}(0)$, it must be the case that $\alpha' C^x(1)=0$. The relation between Γ_t^x, ν_t and ε_t^x is clear from the equation above. y_t can be proved similarly. If the variables in the $\mathrm{I}(1)$vectors are cointegrated with exactly one cointegration relationship, then there exist $\alpha=[1,-\gamma]'$ and $\beta=[1,-\kappa]'$ such that

$$\begin{aligned} \alpha' x_t-\mu-\theta t &= \nu_t\sim\mathrm{I}(0),. \\ \beta' y_t-u-\delta t &= \varsigma_t\sim\mathrm{I}(0). \end{aligned} \tag{3}$$

[1] GLS detrended variables is defined as $y_t^d\equiv y_t-\hat\beta' z_t$, where $\hat\beta$ is obtained from the regressing y_t on z_t in Elliott et al. (1996) Eq. (1) when $d_t=\beta' z_t$ and $z_t=1$ with $\bar\alpha=1-7/T$ or $z_t=[1,\ t]$ with $\bar\alpha=1-13.5/T$.

Although x_t and y_t are cointegrated themselves, we might be interested in the question as to which equilibrium error, ν_t or ς_t has a smaller variance, which represents a closer long-run relationship among the variables in that vector. That is, we are interested in testing the following hypothesis:

$$H_0 : \sigma_\nu^2 = \sigma_\varsigma^2 \, v.s.$$
$$H_1 : \sigma_\nu^2 \neq \sigma_\varsigma^2.$$

As shown in Hansen (1992) and Davidson (2000), of the OLS estimates of the parameters in (1), $(\hat{\gamma},\ \hat{\kappa})$ converge at the rate T, $\left(\hat{\theta}, \hat{\delta}\right)$ converge at the rate $T^{3/2}$ and $(\hat{\mu}, \hat{u})$ converge at the rate $T^{1/2}$. Constructing a test of σ_ν^2 and σ_ς^2 can therefore be based on the OLS residual sum of squares. To this end, we employ a very general dependent class of stochastic processes referred to as the mixingale condition due to McLeish (1974)'s specification of the square of the cointegration equilibrium errors ν_t and ς_t. The result is as follows. First of all, we assume the following conditions:

Assumption 1. Assume that both ν_t and ς_t are stationary ergodic processes and that $E\nu_t^2 = \sigma_\nu^2$, $E\varsigma_t^2 = \sigma_\varsigma^2$, $E\nu_t^4 < \infty$, $E\varsigma_t^4 < \infty$. We further assume that $\{\nu_t^2 - \sigma_\nu^2,\ I_t\}$ and $\{\varsigma_t^2 - \sigma_\varsigma^2,\ I_t\}$ are adapted mixingale with an α_m of size -1 and where $Cov(\nu_s^2,\ \varsigma_\ell^2) = 0, \forall s,\ \ell$.

Let $\{Z_t,\ I_t\}$ be an adapted stochastic sequence with $E(Z_t^2) < \infty$. Then $\{Z_t,\ I_t\}$ is an adapted mixingale if there exist finite nonnegative sequence $\{c_t\}$ and $\{\gamma_m\}$ such that $\gamma_m \to 0$ and $m \to \infty$ and $(E(E(Z_t|I_{t-m})^2))^{1/2} \leq c_t\gamma_m$. We say γ_m is of size $-a$ if $\gamma_m = O(m^{-a-\varphi})$ for some $\varphi > 0$. As the name is intended to suggest, mixingale process has attributed both mixing processes and martingale difference processes. They can be thought of as processes that behave "asymptotically" like martingale difference process, analogous to mixing processes, which behave "asymptotically" like independent processes. Here $\{I_t\}$ is a sequence of σ-fields, $I_{t-1} \subset I_t$. Please see Chapter 5 of White (1999).

The assumptions are general enough to allow the equilibrium error to be dependent identically distributed and enable us to apply the central limit theorem to its square.

Lemma 1. *Given Assumption 1. Let* $\hat{\sigma}_{TW\_S1}^2 = \frac{1}{T-3}\sum_{t=1}^{T}\left(S_{TW,t} - \hat{\mu} - \hat{\gamma}S_{1t} - \hat{\theta}t\right)^2$ *and* $\hat{\sigma}_{TW\_S2}^2 = \frac{1}{T-3}\sum_{t=1}^{T}\left(S_{TW,t} - \hat{u} - \hat{\kappa}S_{2t} - \hat{\delta}t\right)^2$ *then*[2]

$$\frac{T^{1/2}\left(\hat{\sigma}_{TW\_S1}^2 - \sigma_\nu^2\right)}{V_\nu^{1/2}} \overset{A}{\sim} N(0,1) \tag{4}$$

$$\frac{T^{1/2}\left(\hat{\sigma}_{TW\_S2}^2 - \sigma_\varsigma^2\right)}{V_\varsigma^{1/2}} \overset{A}{\sim} N(0,1) \tag{5}$$

and

$$\frac{\left(\hat{\sigma}_{TW\_S1}^2 - \hat{\sigma}_{TW\_S2}^2\right) - \left(\sigma_\nu^2 - \sigma_\varsigma^2\right)}{\left(\frac{V_\nu}{T} + \frac{V_\varsigma}{T}\right)^{1/2}} \stackrel{.}{\sim} N(0,1) \tag{6}$$

where $V_\nu = Var(T^{-1/2}\sum_{t=1}^{T}\nu_t^2)$ and $V_\varsigma = Var(T^{-1/2}\sum_{t=1}^{T}\varsigma_t^2)$.

Proof. In proving the first part of this lemma, we show the case for the x_t vector. The proof of the y_t vector is the same. Define Z as a sample data matrix of the regressor, $z_t = [1, S_{1t}, t]'$, $\xi = \left[\mu,\ \gamma,\ \theta\right]'$, $\hat{\xi} = \left[\hat{\mu},\ \hat{\gamma},\ \hat{\theta}\right]'$, likewise r (column vector) for the regressand, $S_{TW,t}$ and ν as the disturbance ν_t. The deviation of the OLS estimate from its population value is given by

$$\hat{\xi} - \xi = (Z'Z)^{-1}(Z'V). \tag{7}$$

[2] If the trend is not included in the unit root test, $\hat{\sigma}_{TW\_S1}^2 = \sum_{t=1}^{T}\left(S_{TW,t} - \hat{\mu} - \hat{\gamma}S_{1t}\right)^2/T-2$, and $\hat{\sigma}_{TW\_S2}^2 = \sum_{t=1}^{T}\left(S_{TW,t} - \hat{\mu} - \hat{\gamma}S_{2t}\right)^2/T-2$.

The order of the probability of the individual terms in Eq. (7) from the standard asymptotic results of the sample moments of I(1) (such as Hamilton, p. 486) is as follows:

$$\hat{\xi}-\xi=\begin{bmatrix} O_p(T) & O_p\left(T^{3/2}\right) & O_p\left(T^2\right) \\ O_p\left(T^{3/2}\right) & O_p\left(T^2\right) & O_p\left(T^{5/2}\right) \\ O_p\left(T^2\right) & O_p\left(T^{5/2}\right) & O_p\left(T^3\right) \end{bmatrix}^{-1}\begin{bmatrix} O_p\left(T^{1/2}\right) \\ O_p(T) \\ O_p\left(T^{3/2}\right) \end{bmatrix}. \tag{8}$$

Define a rescaling matrix as,

$$\Im_T=\begin{bmatrix} T^{1/2} & 0 & 0 \\ 0 & T & 0 \\ 0 & 0 & T^{3/2} \end{bmatrix}$$

By multiplying the rescaling matrices $\Im_T$ in Eq. (8), we obtain

$$\begin{aligned}\Im_T\left(\hat{\xi}-\xi\right) &= \Im_T(Z'Z)^{-1}\Im_T\Im_T^{-1}(Z'\nu) \\ &= \left[\Im_T^{-1}(Z'Z)\Im_T^{-1}\right]^{-1}\Im_T^{-1}(Z'\nu) \\ &= O(1).\end{aligned} \tag{9}$$

This results show the convergence rate of the OLS estimators. Notice that the population residual sum of squares can be written as

$$\begin{aligned}(r-Z\xi)'(r-Z\xi) &= \left(r-Z\hat{\xi}+Z\hat{\xi}-Z\xi\right)'\left(r-Z\hat{\xi}+Z\hat{\xi}-Z\xi\right) \\ &= \hat{\nu}+Z\left(\hat{\xi}-\xi\right)\Big)'\left(\hat{\nu}+Z\left(\hat{\xi}-\xi\right)\right) \\ &= \left(\hat{\nu}'\hat{\nu}\right)-\left[\left(\hat{\xi}-\xi\right)'\left(Z'Z\right)\left(\hat{\xi}-\xi\right)\right].\end{aligned}$$

The estimator for the disturbance's variance can therefore be expressed as

$$\hat{\sigma}_\nu^2=\frac{1}{T}\left(r-Z\hat{\xi}\right)'\left(r-Z\hat{\xi}\right)=\frac{1}{T}\left(\nu'\nu\right)+\frac{1}{T}\left(\left(\hat{\xi}-\xi\right)'\left(Z'Z\right)\left(\hat{\xi}-\xi\right)\right),$$

or represented by

$$\begin{aligned}\hat{\sigma}_\nu^2-\frac{1}{T}(\nu'\nu) &= \frac{1}{T}\left\{\left[\left(\hat{\xi}-\xi\right)'\Im_T\right]\left[\Im_T^{-1}(Z'Z)\Im_T^{-1}\right]\left[\Im_T\left(\hat{\xi}-\xi\right)\right]\right\} \\ &= \frac{1}{T}O(1)O(1)O(1)=o(1)\end{aligned} \tag{10}$$

Therefore $\hat{\sigma}_\nu^2$ and $\frac{1}{T}(\nu'\nu)$ are asymptotically equivalent.
However,

$$\frac{\sqrt{T}}{T}(\nu'\nu)-\sqrt{T}\sigma_\nu^2=T^{-1/2}\sum_{t=1}^{T}\left(\nu_t^2-\sigma_\nu^2\right),$$

where $\{(\nu_t^2-\sigma_\nu^2),\ I_t\}$ is the adapted mixingale with α_m of size -1 by assumption. Since we have assumed that ν_t is a stationary ergodic process, then $(\nu_t^2-\sigma_\nu^2)$ will also be stationary and ergodic. The proofs can be found in Stout (1974 p. 110, 182) and White (1999, p. 44).

Hence, based on the central limit theorem of a stationary ergodic adapted mixingale of Scott (1973) we have

$$\frac{T^{-1/2}\sum_{t=1}^{T}\left(\nu_t^2-\sigma_\nu^2\right)}{V_\nu^{1/2}} \overset{A}{\sim} N(0,1) \tag{11}$$

Substituting Eq. (10) into Eq. (11), we conclude that

$$\frac{\sqrt{T}\left(\hat{\sigma}_\nu^2-\sigma_\nu^2\right)}{V_\nu^{1/2}} \overset{A}{\sim} N(0,1).$$

Finally, it is obvious that the usual OLS estimate of the variance of the disturbance, $\hat{\sigma}^2_{TW\_S1}$, has the same limiting distribution as $\hat{\sigma}_\nu^2$,$\sqrt{T}(\hat{\sigma}^2_{TW\_S1}-\sigma_\nu^2)-\sqrt{T}(\hat{\sigma}_\nu^2-\sigma_\nu^2)\to^{P}0$,and hence

$$\frac{T^{1/2}\left(\hat{\sigma}^2_{TW\_S1}-\sigma_\nu^2\right)}{V_\nu^{1/2}} \overset{A}{\sim} N(0,1). \tag{12}$$

The same results apply to the vector y_t. This completes the proof of the first part of this Lemma.

For the second part, for a large T,

$$\hat{\sigma}^2_{TW\_S1} \sim N\left(\sigma_\nu^2, T^{-1}V_\nu\right)$$

$$\hat{\sigma}^2_{TW\_S2} \sim N\left(\sigma_\varsigma^2, T^{-1}V_\varsigma\right).$$

Since by assumption $Cov(\nu_s^2,\ \varsigma_\ell^2)=0,\forall s,\ \ell$, so

$$\hat{\sigma}^2_{TW\_S1}-\hat{\sigma}^2_{TW\_S2} \sim N\left(\sigma_\nu^2-\sigma_\varsigma^2, T^{-1}V_\nu+T^{-1}V_\varsigma\right),$$

or equivalently

$$\frac{\left(\hat{\sigma}^2_{TW\_S1}-\hat{\sigma}^2_{TW\_S2}\right)-\left(\sigma_\nu^2-\sigma_\varsigma^2\right)}{\left(\frac{V_\nu}{T}+\frac{V_\varsigma}{T}\right)^{1/2}} \sim N(0,1).$$

□

From this result, it is then straightforward to construct a test regarding the equality of the variance of the two cointegration errors from the above result.

Theorem 1. *Let $\hat{V}_\nu$ and $\hat{V}_\varsigma$ be consistent estimators of V_ν and V_ς in Eqs. (4) and (5), then under the null hypothesis that $H_0:\sigma_\nu^2=\sigma_\varsigma^2$, the statistic*

$$Z_0=\frac{\left(\hat{\sigma}^2_{TW\_S1}-\hat{\sigma}^2_{TW\_S2}\right)}{\left(\frac{\hat{V}_\nu}{T}+\frac{\hat{V}_\varsigma}{T}\right)^{1/2}} \sim N(0,1). \tag{13}$$

Here, $\hat{\sigma}^2_{TW\_S1}=\frac{1}{T-3}\sum_{t=1}^{T}\left(S_{TW,t}-\hat{\mu}-\hat{\gamma}S_{1t}-\hat{\theta}t\right)^2$, $\hat{\sigma}^2_{TW\_S2}=\frac{1}{T-3}\sum_{t=1}^{T}\left(S_{TW,t}-\hat{u}-\hat{\kappa}S_{2t}-\hat{\delta}t\right)^2$. To consistently estimate $V_\nu=Var(T^{-1/2}\sum_{t=1}^{T}\nu_t^2)$ and $V_\varsigma=Var(T^{-1/2}\sum_{t=1}^{T}\varsigma_t^2)$, the Newey–West estimator could be used. In our case it takes this form:

$$\hat{V}_\nu=\hat{\gamma}_0+2\sum_{j=1}^{q}[1-j/(q+1)]\hat{\gamma}_j \tag{14}$$

where $\hat{\gamma}_0 = T^{-1}\sum_{t=1}^{T}\hat{\nu}_t^4, \hat{\gamma}_j = T^{-1}\sum_{t=j+1}^{T}\hat{\nu}_t^2\hat{\nu}_{t-j}^2$, and $1-j/(q+1)$ are the Bartlett kernel. $\hat{V}_\varsigma$ can be proved similarly.

Proof. Since under the null hypothesis that $H_0:\sigma_\nu^2=\sigma_\varsigma^2$, we have

$$\frac{\left(\hat{\sigma}^2_{TW\_S1}-\hat{\sigma}^2_{TW\_S2}\right)}{\left(\frac{\hat{V}_\nu}{T}+\frac{\hat{V}_\varsigma}{T}\right)^{1/2}}-\frac{\left(\hat{\sigma}^2_{TW\_S1}-\hat{\sigma}^2_{TW\_S2}\right)}{\left(\frac{V_\nu}{T}+\frac{V_\varsigma}{T}\right)^{1/2}}$$

$$=\left[\left(\frac{\hat{V}_\nu}{T}+\frac{\hat{V}_\varsigma}{T}\right)^{-1/2}\left(\frac{V_\nu}{T}+\frac{V_\varsigma}{T}\right)^{1/2}-1\right]\left(\frac{V_\nu}{T}+\frac{V_\varsigma}{T}\right)^{-1/2}\left(\hat{\sigma}^2_{TW\_S1}-\hat{\sigma}^2_{TW\_S2}\right)\xrightarrow{p}0.$$

Table 1
Descriptive statistics in different periods.

Period	AUS	HK	MAL	PHI	SIG	SKO	TAI	TW	US
Panel A All 1992/1/1–2007/12/31									
Mean	3088.160	12,125.820	853.175	2059.506	1992.756	847.689	714.046	6278.449	8206.862
Median	2960.000	11,372.780	841.720	1956.495	2011.970	791.550	678.795	6118.835	9088.400
Maximum	6781.000	31,638.220	1447.040	3824.200	3875.770	2064.850	1709.640	10,202.200	14,164.530
Minimum	1357.200	4306.970	262.700	979.340	805.040	280.000	207.310	3135.560	3136.600
Std. dev.	1183.595	4230.408	223.316	717.486	518.258	320.475	353.400	1519.267	3065.236
Skewness	1.158	1.091	0.284	0.432	0.920	1.328	0.691	0.333	−0.347
Kurtosis	4.002	5.317	2.627	2.078	4.673	5.233	2.458	2.463	1.831
Jarque–Bera	806.257	1282.467	58.403	201.978	782.719	1524.141	278.913	92.487	233.881
Panel B Pre-crisis 1992/1/1–1997/7/1									
Mean	2003.497	9060.767	937.792	2409.368	2009.141	801.945	1097.598	5506.083	4436.114
Median	2050.650	9271.020	997.835	2670.230	2107.450	818.560	1162.905	5356.480	3862.850
Maximum	2725.900	15,196.790	1314.460	3447.600	2493.700	1138.750	1709.640	9030.280	7796.500
Minimum	1357.200	4306.970	546.630	1086.110	1311.000	459.090	464.770	3135.560	3136.600
Std. dev.	297.682	2530.669	212.711	699.372	317.835	149.684	271.887	1279.333	1200.295
Skewness	−0.156	0.144	−0.519	−0.555	−0.835	−0.138	−0.190	0.404	0.979
Kurtosis	2.267	2.317	1.879	1.786	2.425	2.190	1.780	2.564	2.826
Jarque–Bera	27.964	24.191	102.859	119.296	137.562	32.309	72.023	37.166	170.432
Panel C During the crisis 1997/7/2–1998/12/31									
Mean	2655.311	10,694.960	607.895	1915.673	1437.953	473.133	408.898	8092.874	8395.045
Median	2666.000	10,295.150	568.010	1892.565	1459.950	447.555	384.455	7902.925	8387.050
Maximum	2881.400	16,673.270	1080.610	2753.150	2007.200	775.260	682.160	10,116.840	9333.100
Minimum	2299.200	6660.420	262.700	1082.180	805.040	280.000	207.310	6251.380	7161.200
Std. dev.	99.374	2565.251	193.234	379.293	341.304	143.989	126.109	949.627	525.070
Skewness	−0.307	0.780	0.540	−0.060	−0.022	0.606	0.277	0.306	0.064
Kurtosis	2.789	2.680	2.514	2.784	1.828	2.241	2.064	2.116	1.629
Jarque–Bera	5.189	31.295	17.299	0.757	16.961	25.206	14.582	14.242	23.384
Panel D Post-crisis 1999/1/1–2007/12/31									
Mean	3845.700	14,302.990	843.127	1864.982	2079.980	942.264	526.710	6444.776	10,542.820
Median	3315.900	13,637.240	810.885	1684.000	1983.365	837.190	496.645	6142.220	10,527.450
Maximum	6781.000	31,638.220	1447.040	3824.200	3875.770	2064.850	907.280	10,202.200	14,164.530
Minimum	2687.100	8435.040	494.570	979.340	1213.820	472.130	250.600	3493.660	7286.270
Std. dev.	1060.097	4013.712	199.616	691.437	582.149	364.261	179.392	1413.309	1263.721
Skewness	1.241	1.397	1.154	1.077	1.113	1.152	0.112	0.456	0.436
Kurtosis	3.326	5.624	4.147	3.547	3.854	3.648	1.520	2.571	3.661
Jarque–Bera	439.674	1030.949	465.769	346.266	399.045	401.813	157.142	71.251	83.986

Notes: The indexes are those for 9 securities markets, namely, the Sydney Australia All-Ordinaries index for Australia (AUS), the Hang Seng Index for Hong Kong (HK), the Kuala Lumpur index for Malaysia (MAL), the Manila Composite index for the Philippines (PHI), the Strait Times Index for Singapore (SIG), the Korea Composite Stock Price Index for South Korea (SKO), the Bangkok SET (Stock Exchange of Thailand) index for Thailand (TAI), the Taiwan Stock Exchange weighted price index for Taiwan (TW), and the New York Dow Jones industrial average for the US (US).

Table 2
Unit root tests in different periods.

Period		AUS	HK	JP	MAL	PHI	SIG	SKO	TAI	TW	US
Panel A Price level											
All	Lag	1	1	1	2	2	2	1	2	1	1
	τ_μ	3.064	1.590	−0.488	0.228	0.070	0.425	0.976	−1.150	−0.980	1.067
	Lag	1	1	1	2	2	2	1	2	1	1
	τ_τ	−0.562	−1.391	−1.681	−1.314	−1.118	−1.223	−0.807	−1.121	−2.092	−2.250
Pre-crisis	Lag	1	1	1	2	2	1	1	2	1	1
	τ_μ	1.542	1.721	−0.831	−0.025	0.031	−0.338	−0.851	−0.702	0.920	3.935
	Lag	1	1	1	2	2	1	2	2	1	1
	τ_τ	−1.487	−2.090	−1.290	−2.088	−1.676	−1.067	−1.269	−0.174	−1.033	−0.021
During the crisis	Lag	3	1	3	1	1	2	1	1	1	1
	τ_μ	−1.262	−0.290	0.595	0.212	−0.229	−0.341	−0.382	−0.526	−0.348	−0.592
	Lag	3	1	1	1	1	2	1	1	1	1
	τ_τ	−2.447	−1.233	−2.299	−0.718	−0.952	−1.390	−0.304	−1.584	−2.185	−2.218
Post-crisis	Lag	1	1	1	2	1	1	1	1	1	1
	τ_μ	2.361	1.517	−1.302	1.690	0.424	1.156	1.333	0.542	−1.499	−0.513
	Lag	1	1	1	2	1	1	1	1	2	1
	τ_τ	−0.692	−0.726	−1.304	−1.173	−0.313	−1.099	−1.193	−1.781	−1.650	−2.112
Panel B First difference											
All	Lag	12	1	17	1	3	1	11	14	3	1
	τ_μ	−10.417 ***	−54.077 ***	−3.002 ***	−48.640 ***	−27.412 ***	−51.763 ***	−10.909 ***	−5.474 ***	−28.005 ***	−55.961 ***
	Lag	1	9	17	1	1	1	1	9	1	1
	τ_τ	−55.752 ***	−13.391 ***	−5.520 ***	−49.206 ***	−48.911 ***	−50.991 ***	−53.220 ***	−11.900 ***	−51.738 ***	−55.967 ***
Pre-crisis	Lag	12	1	12	1	1	1	2	9	1	1
	τ_μ	−4.342 ***	−29.909 ***	−2.803 ***	−28.276 ***	−26.166 ***	−30.061 ***	−17.655 ***	−5.822 ***	−30.593 ***	−33.559 ***
	Lag	1	1	4	1	1	1	2	1	1	1
	τ_τ	−29.726 ***	−30.588 ***	−12.013 ***	−28.514 ***	−26.952 ***	−29.937 ***	−19.714 ***	−28.386 ***	−31.381 ***	−33.478 ***
During the crisis	Lag	1	1	1	1	1	1	1	1	1	1
	τ_μ	−11.939 ***	−17.773 ***	−17.892 ***	−14.000 ***	−14.322 ***	−13.128 ***	−13.998 ***	−18.154 ***	−16.256 ***	−16.481 ***
	Lag	1	1	2	1	1	1	1	1	1	1
	τ_τ	−14.006 ***	−17.791 ***	−15.105 ***	−14.879 ***	−15.678 ***	−13.835 ***	−14.999 ***	−18.151 ***	−17.162 ***	−16.248 ***
Post-crisis	Lag	13	13	18	15	13	13	0	10	12	8
	τ_μ	−2.454 **	−9.734 ***	−1.244	−1.534	−1.351	−2.207 **	−37.521 ***	−3.894 ***	−5.653 ***	−4.564 ***
	Lag	13	13	16	7	13	13	0	10	12	1
	τ_τ	−4.723 ***	−7.973 ***	−8.842 ***	−12.874 ***	−2.999 **	−4.253 ***	−39.962 ***	−6.538 ***	−8.068 ***	−30.016 ***

Notes: ***, ** and * denote statistical significance at the 1%, 5%, and 10% levels, respectively. The indexes are those for 9 securities markets, namely, the Sydney Australia All-Ordinaries index for Australia (AUS), the Hang Seng Index for Hong Kong (HK), the Kuala Lumpur index for Malaysia (MAL), the Manila Composite index for the Philippines (PHI), the Strait Times Index for Singapore (SIG), the Korea Composite Stock Price Index for South Korea (SKO), the Bangkok SET (Stock Exchange of Thailand) index for Thailand (TAI), the Taiwan Stock Exchange weighted price index for Taiwan (TW), and the New York Dow Jones industrial average for the US (US). The critical values are τ_μ including a constant, and τ_τ including both a constant and trend, respectively.

Therefore $(\hat{\sigma}^2_{TW\_S1} - \hat{\sigma}^2_{TW\_S2})\left(\hat{V}_\nu/T + \hat{V}_\varsigma/T\right)^{-1/2}$ and $(\hat{\sigma}^2_{TW\_S1} - \hat{\sigma}^2_{TW\_S2})(V_\nu/T + V_\varsigma/T)^{-1/2}$ are asymptotic equivalence. Hence the statistics

$$Z_0 = \frac{\left(\hat{\sigma}^2_{TW\_S1} - \hat{\sigma}^2_{TW\_S2}\right)}{\left(\frac{\hat{V}_\nu}{T} + \frac{\hat{V}_\varsigma}{T}\right)^{1/2}} \sim N(0,1)$$

□

By rejecting the null hypothesis, we could arrive at the conclusion that smaller variances of disturbances exhibit closer linkages between the independent and dependent variables in the cointegrating regression model as opposed to larger variances of disturbances. Therefore, the unequal variance test that this paper develops could answer the question as to how close the relationships are by comparing the other two countries with Taiwan.

Table 3
The OLS residual sum of squares with TW. $S_{TW,t} = \mu + \gamma S_{it} + \theta t + \nu_t$.

Contury	μ	γ	θ	SSE	$\hat{\sigma}^2_{TW\_S_i}$	Rank
Panel A All 1992/1/1–2007/12/31						
AUS	4074.36 ***	1.1438 ***	−0.8740 ***	1,772,485.00	584.01	4
HK	3252.88 ***	0.3295 ***	−0.6386 ***	1,196,618.84	394.27	1
MAL	3715.45 ***	2.2330 ***	0.4329 ***	1,854,175.87	610.93	6
PHI	3292.43 ***	1.0342 ***	0.5634 ***	1,546,318.39	509.50	3
SIG	3625.24 ***	1.1148 ***	0.2841 ***	1,807,487.89	595.55	5
SKO	4787.55 ***	1.2262 ***	0.2971 ***	1,982,261.68	653.13	7
TAI	5737.44 ***	−0.2576 ***	0.4771 ***	2,089,861.13	688.59	8
US	3157.72 ***	0.6913 ***	−1.6801 ***	1,351,411.61	445.28	2
Panel B Pre-crisis 1992/1/4–1997/7/1						
AUS	1156.65 ***	1.7723 ***	1.5081 ***	709,727.86	672.73	4
HK	2549.22 ***	0.2650 ***	1.0504 ***	690,836.24	654.82	2
MAL	2292.06 ***	2.5877 ***	1.4869 ***	691,288.00	655.25	3
PHI	2818.38 ***	0.7800 ***	1.5269 ***	685,075.56	649.36	1
SIG	3233.45 ***	0.4056 ***	2.7529 ***	755,047.93	715.69	6
SKO	3438.47 ***	0.6469 ***	2.9252 ***	755,124.87	715.76	7
TAI	4220.66 ***	−0.3319 ***	3.1157 ***	754,877.98	715.52	5
US	3416.56 ***	0.1863 ***	2.3853 ***	755,487.91	716.10	8
Panel C During the crisis 1997/7/2–1998/12/31						
AUS	−909.11	3.8857 ***	−8.8597 ***	187,433.67	639.71	2
HK	6980.83 ***	0.1713 ***	−4.8477 ***	257,997.72	880.54	6
MAL	6453.07 ***	3.3035 ***	−2.4804 ***	221,281.90	755.23	4
PHI	5831.43 ***	1.5081 ***	−4.2266 ***	161,954.00	552.74	1
SIG	7033.29 ***	1.2219 ***	−4.6965 ***	286,257.09	976.99	7
SKO	7927.72 ***	2.2296 ***	−5.9914 ***	291,688.80	995.52	8
TAI	6765.90 ***	4.4694 ***	−3.3706 ***	234,632.25	800.79	5
US	2623.64 ***	0.8608 ***	−11.8351 ***	200,034.93	682.71	3
Panel D Post-crisis 1999/1/1–2007/12/31						
AUS	2016.27 ***	1.8258 ***	−3.0779 ***	918,501.54	546.40	7
HK	2593.52 ***	0.3392 ***	−1.1873 ***	642,451.42	382.18	4
MAL	729.75 ***	9.2607 ***	−2.4842 ***	383,929.17	228.39	1
PHI	3835.58 ***	1.9341 ***	−1.1845 ***	705,645.81	419.78	5
SIG	2239.58 ***	2.7662 ***	−1.8379 ***	458,093.26	272.51	2
SKO	3826.22 ***	5.1730 ***	−2.6774 ***	513,390.16	305.41	3
TAI	3939.38 ***	8.6277 ***	−2.4201 ***	1,323,606.10	787.39	8
US	−3423.94 ***	1.0138 ***	−0.9723 ***	715,275.24	425.51	6

Note: ***, ** and * denote statistical significance at the 1%, 5%, and 10% levels, respectively. S_i represents the 8 stock indexes, namely, the Sydney Australia All-Ordinaries index for Australia (AUS), the Hang Seng Index for Hong Kong (HK), the Kuala Lumpur index for Malaysia (MAL), the Manila Composite index for the Philippines (PHI), the Strait Times Index for Singapore (SIG), the Korea Composite Stock Price Index for South Korea (SKO), the Bangkok SET (Stock Exchange of Thailand) index for Thailand (TAI), and the New York Dow Jones industrial average for the US (US).

3. Data and empirical results

The data we use comprise the indexes for eight East Asian stock markets, namely, the Sydney Australia All-Ordinaries index for Australia (AUS), the Hang Seng Index for Hong Kong (HK), the Kuala Lumpur index for Malaysia (MAL), the Manila Composite index for the Philippines (PHI), the Strait Times Index for Singapore (SIG), the Korea Composite Stock Price Index for South Korea (SKO), the Bangkok SET (Stock Exchange of Thailand) index for Thailand (TAI), and the Taiwan Stock Exchange weighted price index for Taiwan (TW). Besides these Asian stock markets, the US capital market is decisively important for Taiwan, as it is a prime and leading stock market and has a powerful impact on frequent trades and business cooperation. In this paper we would therefore like to observe how similar the impacts from the US and other Asian countries are using the Dow Jones Industrial Average Index. Our data are daily frequency stock price series ranging from January 1, 1992 to December 31, 2007 that are taken from the Datastream database. To compare the pattern of the Asian financial crisis in 1997, the sample period is divided into three stages: the pre-crisis stage (1992/1/1–1997/7/1), the crisis stage (1997/7/2–1998/12/31), and the post-crisis stage (1999/1/1–2007/12/31). The descriptive statistics for different periods are as shown in Table 1.

To pretest each variable to determine its order of integration, we use the DF–GLS unit root test to examine the stationarity of the variables and specify the number of lags to be added to the test regression which depends on removing the serial correlation in the residuals. According to the statistics in Panel A, Table 2, non-stationarity cannot be rejected for the levels of all stock price series at the 5% significance level based on the DF–GLS test. This means that the stock price indexes are all variables that contain a unit root. This is similar to the findings of numerous studies in which many financial time series contain unit roots that are dominated by stochastic trends. After first differencing the variables in Panel B, Table 2, all variables appear to be stationary processes. It is thus reasonable to conclude that the variables are integrated of order one. The traditional regression model can therefore not be adapted to discuss the relationships between pairs of countries, and a spurious regression model is called for.

The Engle and Granger (1987) cointegration test is conducted to check the long-run equilibrium of non-stationary series between the prices of the Taiwan stock market and other such markets. To estimate the long-run equilibrium relationship in the form of Eq. (1), we follow Engle and Granger's (1987) two-step procedure and estimate the regression model as in Table 3. Except for the intercept term for Australia in Panel C, all the coefficients in Table 3 are significantly different from zero. This is the apodictic sequence and typical symptom because the variables in the model are non-stationary according to the spurious regression model. Granger and Newbold (1974) convincingly show that regression models using unit root time-series data generally lead to spurious results. However, Eq. (1) does not lead to spurious results in the case in which two seemingly non-stationary variables are related. That is, the problems associated

Table 4
Engle–Granger cointegration test with Taiwan.

Period		AUS	HK	MAL	PHI	SIG	SKO	TAI	US
All	τ_μ	−1.596	−2.557 **	−2.201 **	−2.461 **	−1.999 **	−1.751 *	−1.676 *	−2.312 **
	τ_τ	−1.996	−2.610 *	−2.241	−2.480	−2.159	−2.057	−2.089	−2.720 *
Pre-crisis	τ_μ	−1.438	−1.225	−0.995	−1.016	−1.082	−1.123	−1.281	−1.295
	τ_τ	−1.404	−1.265	−0.745	−0.781	−0.905	−0.969	−1.204	−1.222
During the crisis	τ_μ	−1.890 *	−1.784 *	−1.308	−1.427	−1.857 *	−1.687 *	−2.002 **	−2.693 **
	τ_τ	−2.517	−2.017	−1.659	−1.911	−1.968	−1.934	−2.286	−3.073 **
Post-crisis	τ_μ	−1.623 *	−2.443 **	−3.576 ***	−1.288	−3.347 ***	−2.217 **	−1.312	−3.433 ***
	τ_τ	−2.094	−2.653 *	−3.761 ***	−1.977	−3.375 **	−2.705 *	−1.942	−3.470 **

Notes: ***, ** and * denote statistical significance at the 1%, 5%, and 10% levels, respectively. The Engle–Granger cointegration test with Taiwan is conducted for 8 securities markets, namely, the Sydney Australia All-Ordinaries index for Australia (AUS), the Hang Seng Index for Hong Kong (HK), the Kuala Lumpur index for Malaysia (MAL), the Manila Composite index for the Philippines (PHI), the Strait Times Index for Singapore (SIG), the Korea Composite Stock Price Index for South Korea (SKO), the Bangkok SET (Stock Exchange of Thailand) index for Thailand (TAI), the Taiwan Stock Exchange weighted price index for Taiwan (TW), and the New York Dow Jones industrial average for the US (US). This table presents the DF–GLS test of the residuals with Taiwan. The critical values are τ_μ including a constant, and τ_τ including both a constant and trend for the DF–GLS test, respectively.

with the spurious regression model disappear. The residual sequences are inferred to be stationary while the sequences exhibit a cointegration relationship.

Before showing the unit root test of residual sequences, another issue we may concern is how close the relationship of a country to Taiwan changes in all period and three sub-periods. From the Lemma, The

Table 5
The unequal variance tests between different countries based on Taiwan.

Panel A All 1992/1/1–2007/12/31

	$\hat{\sigma}^2_{TW\_HK}$	$\hat{\sigma}^2_{TW\_US}$	$\hat{\sigma}^2_{TW\_PHI}$	$\hat{\sigma}^2_{TW\_AUS}$ [a]	$\hat{\sigma}^2_{TW\_SIG}$	$\hat{\sigma}^2_{TW\_MAL}$	$\hat{\sigma}^2_{TW\_SKO}$	$\hat{\sigma}^2_{TW\_TAI}$
$\hat{\sigma}^2_{TW\_HK}$	0.000	−1.140	−2.216**	−3.546	−3.650***	−3.996 ***	−4.532 ***	−5.195 ***
$\hat{\sigma}^2_{TW\_US}$	1.140	0.000	−1.194	−2.511	−2.644 ***	−2.961 ***	−3.537 ***	−4.173 ***
$\hat{\sigma}^2_{TW\_PHI}$	2.216 ***	1.194	0.000	−1.216	−1.372	−1.639	−2.228***	−2.796 ***
$\hat{\sigma}^2_{TW\_AUS}$ [a]	3.546	2.511	1.216	0.000	−0.180	−0.426	−1.052	−1.602
$\hat{\sigma}^2_{TW\_SIG}$	3.650 ***	2.644 ***	1.372	0.180	0.000	−0.238	−0.859	−1.397
$\hat{\sigma}^2_{TW\_MAL}$	3.996 ***	2.961 ***	1.639	0.426	0.238	0.000	−0.637	−1.179
$\hat{\sigma}^2_{TW\_SKO}$	4.532 ***	3.537 ***	2.2228 ***	1.052	0.859	0.637	0.000	−0.519
$\hat{\sigma}^2_{TW\_TAI}$	5.195 ***	4.173 ***	2.796 ***	1.602	1.397	1.179	0.519	0.000

Panel B Pre-crisis 1992/1/4–1997/7/1

	$\hat{\sigma}^2_{TW\_PHI}$ [a]	$\hat{\sigma}^2_{TW\_HK}$ [a]	$\hat{\sigma}^2_{TW\_MAL}$ [a]	$\hat{\sigma}^2_{TW\_AUS}$ [a]	$\hat{\sigma}^2_{TW\_TAI}$ [a]	$\hat{\sigma}^2_{TW\_SIG}$ [a]	$\hat{\sigma}^2_{TW\_SKO}$ [a]	$\hat{\sigma}^2_{TW\_US}$ [a]
$\hat{\sigma}^2_{TW\_PHI}$ [a]	0.000	−0.053	−0.055	−0.223	−0.629	−0.619	−0.622	−0.632
$\hat{\sigma}^2_{TW\_HK}$ [a]	0.053	0.000	−0.004	−0.178	−0.600	−0.590	−0.593	−0.604
$\hat{\sigma}^2_{TW\_MAL}$ [a]	0.055	0.004	0.000	−0.168	−0.579	−0.569	−0.572	−0.582
$\hat{\sigma}^2_{TW\_AUS}$ [a]	0.223	0.178	0.168	0.000	−0.416	−0.409	−0.412	−0.420
$\hat{\sigma}^2_{TW\_TAI}$ [a]	0.629	0.600	0.579	0.416	0.000	−0.002	−0.002	−0.006
$\hat{\sigma}^2_{TW\_SIG}$ [a]	0.619	0.590	0.569	0.409	0.002	0.000	−0.001	−0.004
$\hat{\sigma}^2_{TW\_SKO}$ [a]	0.622	0.593	0.572	0.412	0.002	0.001	0.000	−0.003
$\hat{\sigma}^2_{TW\_US}$ [a]	0.632	0.604	0.582	0.420	0.006	0.004	0.003	0.000

Panel C During the crisis 1997/7/–1998/12/31

	$\hat{\sigma}^2_{TW\_PHI}$ [a]	$\hat{\sigma}^2_{TW\_AUS}$	$\hat{\sigma}^2_{TW\_US}$	$\hat{\sigma}^2_{TW\_MAL}$ [a]	$\hat{\sigma}^2_{TW\_TAI}$	$\hat{\sigma}^2_{TW\_HK}$	$\hat{\sigma}^2_{TW\_SIG}$	$\hat{\sigma}^2_{TW\_SKO}$
$\hat{\sigma}^2_{TW\_PHI}$ [a]	0.000	−0.403	−0.602	−0.875	−1.002	−1.310	−1.476	−1.579
$\hat{\sigma}^2_{TW\_AUS}$	0.403	**0.000**	**−0.208**	−0.531	**−0.701**	**−1.043**	**−1.303**	**−1.398**
$\hat{\sigma}^2_{TW\_US}$	0.602	**0.208**	**0.000**	−0.334	**−0.515**	**−0.858**	**−1.139**	**−1.231**
$\hat{\sigma}^2_{TW\_MAL}$ [a]	0.875	0.531	0.334	0.000	−0.191	−0.522	−0.831	−0.915
$\hat{\sigma}^2_{TW\_TAI}$	1.002	**0.701**	**0.515**	0.191	**0.000**	**−0.318**	**−0.636**	**−0.714**
$\hat{\sigma}^2_{TW\_HK}$	1.310	**1.043**	0.858	0.522	**0.318**	**0.000**	**−0.347**	**−0.420**
$\hat{\sigma}^2_{TW\_SIG}$	1.476	**1.303**	**1.139**	**0.831**	**0.636**	**0.347**	**0.000**	**−0.062**
$\hat{\sigma}^2_{TW\_SKO}$	1.579	**1.398**	**1.231**	0.915	**0.714**	**0.420**	**0.062**	**0.000**

Panel D Post−crisis 1999/1/1–2007/12/31

	$\hat{\sigma}^2_{TW\_MAL}$	$\hat{\sigma}^2_{TW\_SIG}$	$\hat{\sigma}^2_{TW\_SKO}$	$\hat{\sigma}^2_{TW\_HK}$	$\hat{\sigma}^2_{TW\_PHI}$ [a]	$\hat{\sigma}^2_{TW\_US}$	$\hat{\sigma}^2_{TW\_AUS}$	$\hat{\sigma}^2_{TW\_TAI}$ [a]
$\hat{\sigma}^2_{TW\_MAL}$	0.000	−1.093	−1.955 *	−3.216 ***	−3.167	−3.183 ***	−4.810 ***	−6.929
$\hat{\sigma}^2_{TW\_SIG}$	1.093	0.000	−0.707	−2.038 **	−2.256	−2.295 **	−3.881 ***	−6.102
$\hat{\sigma}^2_{TW\_SKO}$	1.955	0.707	0.000	−1.446	−1.769	−1.817 **	−3.442 ***	5.744
$\hat{\sigma}^2_{TW\_HK}$	3.216 ***	2.038 **	1.446	0.000	−0.536	−0.607	−2.187 **	−4.595
$\hat{\sigma}^2_{TW\_PHI}$ [a]	3.167	2.256	1.769	0.536	0.000	−0.071	−1.513	−3.845
$\hat{\sigma}^2_{TW\_US}$	3.183 ***	2.295 **	1.817 **	0.607	0.071	0.000	−1.426	−3.748
$\hat{\sigma}^2_{TW\_AUS}$	4.810 ***	3.881 ***	3.442 ***	2.187 **	1.513	1.426	0.000	−2.427
$\hat{\sigma}^2_{TW\_TAI}$ [a]	6.929	6.102	5.744	4.595	3.845	3.748	2.427	0.000

Notes: ***, ** and * denote statistical significance at the 1%, 5%, and 10% levels, respectively. The indexes are those of 9 securities markets, namely, the Sydney Australia All-Ordinaries index for Australia (AUS), the Hang Seng Index for Hong Kong (HK), the Kuala Lumpur index for Malaysia (MAL), the Manila Composite index for the Philippines (PHI), the Strait Times Index for Singapore (SIG), the Korea Composite Stock Price Index for South Korea (SKO), the Bangkok SET (Stock Exchange of Thailand) index for Thailand (TAI), the Taiwan Stock Exchange weighted price index for Taiwan (TW), and the New York Dow Jones industrial average for the US (US).

[a] The columns or rows indicate that the disturbance is non-stationary in Table 5, and the boldfaced letters indicate that the disturbance is stationary in Table 5.

$\hat{\sigma}^2_{TW\_Sj}$ sheds light on how close is the relationship of a country to Taiwan, thereby, the SSE also has the same meanings because $\hat{\sigma}^2_{TW\_Sj}$ is derived from SSE. Although the rank in all periods is so different from the three sub-periods (for example, Hong Kong ranks (2, 6, 4) before, during and after the crises, respectively), it is not conflicting about the results. We got one set of coefficients of regression model in the whole period and there is only one linear relationship between the specific country and Taiwan. However we got three sets of coefficients of regression models in three different sub-periods and there are three different linear relationships for each sub-period between the specific country and Taiwan. The distances between the observations and the estimation value for the three slope coefficients of the three sub-periods should be very different from the case with only one slope coefficient of the whole period. This is why we divide sample period to observe before, during, and after crisis.

Nevertheless, the $\hat{\sigma}^2_{TW\_Sj}$ and SSE in Table 3 allow us to see how the residual variances change over the three different periods. Later we report the estimation of the unequal variance test in Table 5 according to the ranks of $\hat{\sigma}^2_{TW\_Sj}$ and SSE.

The empirical results of the unit root test of residual sequences are as shown in Table 4 below:

The results in Table 4 suggest that cointegration is found to exist between Taiwan and almost all the other capital markets for the whole sample period except that for Australia. However, if the whole period is divided into three sub-periods, different co-movement relationships are found both before and after the 1997 financial crisis. No long-run equilibrium between Taiwan's capital market and another capital market is found before the financial crisis since the null hypothesis that there is a non-stationary series cannot be rejected. The null hypothesis can, however, be rejected for six stock markets apart from the Philippines and Malaysia during the period of the financial crisis, and apart from the Philippines and Thailand after the financial crisis. In other words, this implies that there exists an equilibrium relationship between the Taiwan stock market and the other six stock markets both during and after the financial crises.

There are a number of points that give rise to concerns. First, Thailand exhibits a cointegration relationship with Taiwan only during the Asian financial crisis period but not before or after the crisis. This implies that the temporary nature of the co-movement between the Taiwan and Thailand stock markets may be due to an important financial event and be strengthened by the amount of the information made available during the event period and the speed at which it is transmitted. Since the financial crisis spreads rapidly and is overwhelming, the cointegration relationship between the Taiwan and Thailand stock markets can be observed in the crisis period temporarily even though it had not originally been observed. Other evidence of note is that where the co-movement occurs between Malaysia and Taiwan after the financial crisis, but not before or during the financial crisis. The changing relationship is attributable to the fact that the Asian financial crisis has changed and affected the relationships between Taiwan and other countries.

As with the failure of the cointegration test that this paper emphasizes, even though the co-movement exists among international markets, we have no idea how to distinguish different levels of the co-movement among these countries relative to the Taiwan stock market. This paper therefore develops the unequal variance test in order to observe "how close" the linkages are between the Taiwan capital market and other international capital markets.

Next, the unequal variance test is applied to the residuals from this regression model as expressed in Eq. (13). The null hypothesis of the unequal variance test is that the variances formed by the residuals of each cointegrating regression are equal. By rejecting the null hypothesis, we can conclude that smaller variances among the disturbances exhibit closer linkages between the variables in the cointegrating regression models than do bigger variances among the disturbances. One of the strengths of this test is that we can observe how close the linkage between the two variables is, thus making the cointegration relationship clear. The results of the estimation of the unequal variance test are shown in Table 5.

Although the common impressions held by investors are that the stock markets of the US and Taiwan appear to be strongly linked, we find that there exists a closer relationship between Taiwan and Hong Kong than between Taiwan and other countries for the sample period as a whole, as shown in Panel A of Table 5. Panels B, C, and D in Table 5 represent the pre-crisis, during the crisis, and post-crisis periods, respectively. Because there is no evidence of a cointegration relationship between Taiwan and any other countries prior to the financial crisis, Panel B of Table 5 does not merit our attention even though it is posted here. In Panel C of Table 5, while there are no special relationships between Taiwan and other countries, there do exist cointegrating relationships between Taiwan and each of several countries including Australia, Hong Kong, Singapore, South Korea, Thailand and the US. Malaysia is found to be the most closely related to Taiwan in the post-crisis period, and not the US as shown in Panel D of Table 5. The cases of Hong Kong and Taiwan for the whole period and of Malaysia and Taiwan for the post-crisis period indicate that the adjacent region or similarity of background of the capital markets reflects a similar price pattern.

4. Conclusion

This paper focuses on how close the linkages were between Taiwan and other countries in the event of the 1997 Asian financial crisis. The novelty of our paper rests mainly on the unequal variance test that is performed to point out the different relationship conditions. The unequal variance test could help us to find the countries which are closest to the goal country, Taiwan, in international capital markets. Ever since the Asian crisis erupted in 1997, there has been a significant impact on the whole of Asia.

The purpose of this study is to ascertain the co-movement patterns of Taiwan and other countries before the 1997 Asian financial crisis as compared with both during and after the crises. This evidence shows that the co-movement patterns are found to exist during the crisis period but the conclusion that there exists an especially close linkage between Taiwan and any other country cannot be reached. Many investors focus solely on their imagination when diversifying risks, such as countries exhibiting a leader–follower relationship while ignoring the similarities in their stock markets. The co-movements between Taiwan and the US are not as strong as investors imagine and we find the relationship between Taiwan and Hong Kong to be stronger than that between Taiwan and the US for the whole sample period. The financial crisis is found to lead to a greater change in the regional economies, for example, the cointegrating relationship that formerly did not exist has now been switched to one that does exist. It appears that the adjacent regions or similarity of background of the capital markets will be reflected by similar price patterns.

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