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How close a relationship does a capital market have with other markets? A reexamination based on the equal variance test ☆

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ABSTRACT

The cointegration test cannot discriminate closer relationships from cointegrating relationships. In most applications, we must assess the degrees of cointegrating relationships, for example, to examine the comovement between international stock markets using the cointegration methodology. Lee et al. (2012) introduced a variance test of cointegration equilibrium errors to measure the similarity of these relationships. However, the key assumption of cross-sectional independence between a panel of two country-pair squared cointegrating equilibrium errors in their model is not desirable. The appearance of cross-sectional dependence of individual (stock) markets in a panel is a common existence. The current paper shows that the consideration of cross-sectional dependence and the method of estimating long-run variance are important. Our results, which extend the cross-sectional dependence of some Asian stock markets during the Asian financial crisis (1997–1998) documented by Lee et al. (2012), indicate that the similarity of background and business cooperation (or trading activities) are all crucial factors for determining the price patterns by the "equal variance test" proposed in this paper. The analysis of the 2007–2009 global financial crisis is included to confirm the robustness of the results.

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1. Introduction

Over the past two decades, researchers have paid increased attention to the comovement patterns among international stock markets. Early empirical studies investigated the comovement patterns among international stock markets based on a simple correlation analysis of returns or dynamic conditional correlation in the multivariate generalized autoregressive conditional heteroscedasticity (GARCH) framework.[1] However, focusing on stock returns and returns volatility rather than equity prices may yield unstable and often conflicting short-term empirical results (Kasa, 1992; Manning, 2002; Yang et al., 2006). Thus, the numerous studies that examine the comovement patterns among asset prices using either a bivariate or multivariate cointegration methodology can be used to complement the investigation of international stock markets (e.g., Kasa, 1992; Richards, 1995; Rangvid, 2001; Ghosh et al., 2005; Yang et al., 2006; Valadkhani and Chancharat, 2008; Lee et al., 2012).

Cointegration among stock markets can naturally result from the existence of a common feature among stock markets (Engle and Susmel, 1993). Based on this realization, a large number of authors have attempted to explain the factors underlying the comovement among stock markets. Most of the recent studies on this topic are devoted to determining the relative importance of both economic and geographical ties, but the cause of comovement remains enigmatic. Some studies tend to support the dominant importance of economic ties. For example, Johnson and Soenen (2002) showed that an increased export share by Asian economies to Japan and greater foreign direct investment from Japan to other Asian economies contributed to greater comovement. More recently, Didier et al. (2012) indicated that comovement in the stock market is driven largely by financial linkages. Fernández-Avilés et al. (2012) showed that stock market linkages are unrelated to geographical proximity. However, other authors have showed that both economic and geographical ties are important, or that geographical ties influence the pattern of stock prices. For example, Madaleno and Pinho (2012)reported results suggesting that geographically and economically closer markets exhibit a higher correlation and more short-run comovements. Lee et al. (2012) confirmed that geographic ties, not trading activities/business cooperation, would be reflected by most of the comovement patterns among stock markets.

An assessment of the relationships among international stock markets is crucial to exploring the comovement patterns or the factors underlying the comovement of capital markets in the cointegration framework. Lee et al. (2012) proposed a residual-based variance test to discriminate closer relationships from cointegrating relationships by comparing the variances of the cointegrating equilibrium errors from the statistics calculated from the ordinary least squares (OLS)-estimated squared cointegrating residuals. This test can be treated as an extension of the concept of cointegration.

Unlike analysis based on returns, such as correlation analysis, the degree of cointegration provides information on the long-term common trend. Alexander (1999) indicated that cointegration and correlation are related but are different concepts. A high correlation of returns does not necessarily imply a high cointegration in prices. Fig. 1 in Appendix B shows that the degree of cointegration can be measured by comparing the variances of cointegrating errors. Because the scale of the variables is a determining factor of the magnitude of the variances of cointegrating equilibrium errors, the variance tests should be used to compare the degrees of cointegrating relationships that contain a common dependent or independent variable. Nevertheless, in most applications, we must assess the degrees of cointegrating relationships between a panel of two pairs of countries which contains a common country, satisfying the requirement of including a common dependent or independent variable in the relationship. Therefore measuring the degree of cointegration by conducting the variance test can be a useful complement to the analysis of comovement patterns among international stock markets based on correlations of returns.

However, the variance test proposed by Lee et al. (2012) has some limitations. The appearance of cross-sectional dependence of individual time series in a panel is a common existence. According to Lee et al. (2012), the key assumption of cross-sectional independence between a panel of two country-pair squared cointegrating equilibrium errors is not desirable. Some common unobservable factors or omitted variables can lead to cross-sectionally dependent cointegrating equilibrium errors, especially for country-pair regressions, hence lead to cross-section dependence between squared cointegrating equilibrium

[1] Arshanapalli et al. (1995) summarized the limitations of the methodologies based on the stock returns. Furthermore, Li et al. (2012a) reviewed a wide range of related terms and methodologies used in the literature of interdependence in financial markets.

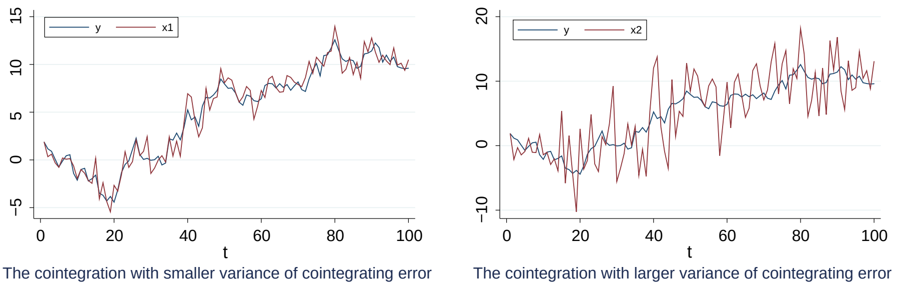


Fig. 1. The variances of cointegrating errors and the degree of cointegration.

errors. In one extreme case, even cross-sectional independent cointegrating errors can induce cross-correlated squared cointegrating equilibrium errors among the panel data. Consequently, it may be necessary to relax the assumption by allowing for cross-sectional dependence between squared cointegrating equilibrium errors.

Another drawback of the test proposed by Lee et al. (2012) is that it uses White (1980) heteroscedasticity-autocorrelation consistent (HAC) estimator to estimate variance–covariance of squared cointegrating equilibrium errors. The use of a HAC estimator involves the specification of a kernel and a truncation lag or bandwidth. The bandwidth choice determines the fraction of the available covariance information that goes into the calculation of the long run variances. Kiefer et al. (2000) showed that even if a data-dependent method is used to choose the truncation lag (bandwidth), arbitrary choices of the truncation lag are inevitable. Furthermore, HAC has a poor finite sample performance, (see, for example, Kiefer et al., 2000; Kiefer and Vogelsang, 2005; Phillips et al., 2006, 2007). Kiefer et al. (2000) proposed an alternative method of constructing robust test statistics; in this method, estimates of the variance–covariance matrix are not explicitly required to construct the test. This approach requires a nonsingular data-dependent stochastic transformation to the OLS estimates. Therefore, arbitrary choices of the truncation lags in HAC can be avoided, and the test based on KVB approach is asymptotically invariant to serial correlation/heteroskedasticity nuisance parameters. Furthermore, Ray and Savin (2008) and Ray et al. (2009) showed that the HAC-based method has an unsatisfactory size control, whereas the KVB-based approach provides a substantially more accurate approximation to the finite sample distribution.

The method proposed in this paper improves the variance test proposed by Lee et al. (2012) in two ways. First, we allow the cross-sectional dependence between squared cointegrating equilibrium errors. To achieve this, we employ the concept of the near-epoch dependence (NED) on a mixing process because, under suitable size of the underlying mixing and moment restrictions, NED is general enough to enable the application of the central limit theorem to the squared cointegrating equilibrium errors. The NED approach can also accommodate a variety of possible process of the squared cointegrating equilibrium errors, such as GARCH(p,q). This accommodation is important given the considerable evidence that the conditional variance–covariance matrix in financial time series can be described as a GARCH-type model. Second, the method proposed in this paper reconstructs Lee et al.'s (2012) variance test by using the KVB approach. The proposed test has the advantage of being simple and intuitive. The limit distribution of the proposed test is free of nuisance parameters, and the critical values of the proposed test are also tabulated. The simulations presented in this study indicate that the finite performance of Lee et al. (2012) test is sensitive to the serial correlation of cointegrating errors and cross-sectional dependence between cointegrating errors, whereas the proposed test in this paper has favorable finite sample performance.[2] Hence, the proposed test can be expected to yield more robust empirical results.

The test proposed in this paper was used to reexamine the comovement patterns among stock markets including Taiwan, the United States, and other Asian markets during the 1997 Asian financial crisis, which were examined by Lee et al. (2012). Lee et al.(2012) indicated that the main reason for focusing on Taiwan is that much trade has taken place between Taiwan and countries other than those in the Asian region

[2] This finite-sample performance was measured through a restricted data generating process by Monte Carlo simulations. These results are reported in Appendix C.

(e.g., the United States) over the past several decades. Therefore, focusing on the relationships between Taiwan and other Asian markets enables the exploration of the relative importance of both economic and geographical ties.

The results of this study are in sharp contrast with those reported by Lee et al. (2012) even though both studies draw from the same data. First, we find closer relationships between Taiwan and the Philippines, Australia, the United States, and Thailand than between Taiwan and other countries considered during the 1997 Asian financial crisis. By contrast, Lee et al. (2012) did not find that Taiwan had any close relationships with other country for this stage. Second, this paper does not show a closer relationship between Taiwan and Malaysia than between Taiwan and other countries, including the United States for the post-crisis period, whereas Lee et al. (2012) found a closer relationship here. Thus, our results support the existence of the same extent of comovements among these selected markets after the 1997 Asian financial crisis. Therefore, after considering the lack of a cointegration relationship before the 1997 crisis, and the cointegrating relationships are confirmed and the extent of the comovement among the selected stock markets shows no significant difference in the post-crisis period, we conclude that the linkage among stock markets was strengthened after the Asian financial crisis. The relatively close relationship between Taiwan and the United States during this crisis confirms the leadership and influence of the US economy, whereas geographic ties cannot be rejected in considering the close relationships between Taiwan and its adjacent markets, such as Hong Kong, Singapore, South Korea, and Malaysia, during and after the 1997 crisis. It is also shown that our empirical results are robust for the 2007–2009 financial crisis when the analysis is extended to the sample period July 1, 2007 to December 31, 2012.

The remainder of the paper is organized as follows. Section 2 presents the test methodology, and presents the null distribution of the proposed test. This section also presents tabulated critical values of the proposed test. Section 3 presents empirical results. Finally, Section 4 provides the concluding remarks. Throughout this paper, $\xrightarrow{d}$ denotes convergence in distribution; $\xrightarrow{a.s.}$ denotes almost sure convergence; $[Tr]$ denotes the largest integer not exceeding Tr, and $\|.\|_p$ denotes $(\mathrm{E}|.|^p)^{1/p}$.

2. The methodology: Test for equality of variances

2.1. Model and assumptions

In this section, we extend the variance test proposed by Lee et al. (2012) to accommodate the cross-sectional dependence between the squares of cointegrating equilibrium errors, and use the KVB approach to reconstruct the test statistics.

Let $\mathbf{x}_t = (S_{TW,t}, S_{1t})'$ and $\mathbf{x}_y = (S_{TW,t}, S_{2t})'$ be two vectors of a 2×1 $I(1)$ time series process, where each vector contains one deterministic trend and one stochastic trend:

$$\begin{aligned} \mathbf{x}_t &= \tau^x t + \Gamma_t^x, \\ \mathbf{y}_t &= \tau^y t + \Gamma_t^y, t = 1, 2, \ldots, T, \end{aligned} \tag{1}$$

where τ^x, τ^y are constant vectors of the parameters of deterministic trends, 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 $\eta_t^x = \mathbf{C}^x(L)\varepsilon_t^x, \eta_t^y = \mathbf{C}^y(L)\varepsilon_t^y$, where $\mathrm{C}^x(.)$, $\mathrm{C}^y(.)$ represents a matrix-valued polynomial, and ε_t^x and ε_t^y are vector white noises.

Assume that there exists only one cointegration relationship in the vectors $\mathbf{x}_t$ and $\boldsymbol{y}_t$ such that $\alpha_x \mathbf{x}_t \sim I(0)$ and $\alpha_y \mathbf{y}_t \sim I(0)$, respectively. Let the cointegration vectors be normalized such that $\alpha_x = (1, -\beta_x)'$, and $\alpha_y = (1, -\beta_y)'$, Lee et al. (2012) used multivariate Beveriage–Nelson decomposition to show that these two cointegrating relations $\alpha_x' \mathbf{x}_t \sim I(0)$ and $\alpha_y' \mathbf{y}_t \sim I(0)$ can be represented by

$$\begin{aligned} S_{TW,t} - \beta_x S_{1t} - \mu_x - \delta_x t &= \nu_t \sim I(0), \\ S_{TW,t} - \beta_y S_{2t} - \mu_y - \delta_y t &= \varsigma_t \sim I(0), \end{aligned} \tag{2}$$

where μ_x, δ_x, μ_y, and δ_y are constants and ν_t and ς_t are de-meaned and de-trended stationary cointegrating equilibrium errors.

Given that $\mathbf{x}_t$ and $\mathbf{y}_t$ are cointegrated, we are interested in discriminating the closer relationship from the cointegrating relationships. To this end, we compare the magnitude of the variances of cointegrating equilibrium errors by developing the equal variance test of the null hypothesis,

$$H_0 : \sigma_\nu^2 = \sigma_\varsigma^2, \quad \text{v.s.} \quad H_1 : \sigma_\nu^2 \gtrless \sigma_\varsigma^2,$$

in which a smaller variance of cointegrating equilibrium error is associated with a closer linkage between the variables in the cointegrating regression model. Hence, the rejection of the null hypothesis can discriminate the closer relationship from the cointegrating relationships.

Before constructing the test and performing statistical inference, we must be precise about the dependence structure of the squared cointegrating equilibrium errors $\{(\nu_t^2,\varsigma_t^2)'\}_{t=0}^{\infty}$ in Eq. (2) to ensure the validity of a functional central limit theorem. In modeling the dependence or memory of a time series in econometrics, nonparametric methods, of which martingale difference and mixing are perhaps most commonly used, are very popular because of their flexibility and model-free characteristics. In constructing the variance test, Lee et al. (2012) assumed that the squared cointegrating equilibrium errors are cross-sectionally uncorrelated adapted mixingale where the mixingale process has attributed both mixing processes and martigale difference processes.[3] However, martingale differences are sequences of a rather special kind and many important dependent processes are not mixing, e.g., infinite moving average (MA(∞)) under general conditions and stable first-order linear autoregressive models with Bernoulli-distributed *i.i.d.* shocks (Andrews, 1984). The mixing condition is also difficult or impossible to check because establishing conditions where uniform mixing holds is harder, while the best known sufficient condition (e.g., Davidson, 1994, Theorem 14.4) requires the shock process to be bounded with probability 1, ruling out normality for example. In addition, there is considerable evidence to model the conditional variance and covariance in financial time series by using a GARCH-type model since Engle (1982) and Bollerslev (1986). Thus, the squared cointegrating equilibrium errors in Eq. (2) may be governed by a GARCH model. Unfortunately, it is not known under what conditions GARCH processes are mixing (Hansen, 1991). Therefore, the mixing assumption might not be the best candidate in modeling the squared cointegrating equilibrium errors.

Instead of assuming that the squared cointegrating equilibrium errors are adapted mixingale and cross-sectionally uncorrelated, the current paper employs the near-epoch dependence (NED) which can go back at least to Ibragimov (1962). We also allow for cross-sectional dependence between the squared cointegrating equilibrium errors in Eq. (2). The NED assumption has the benefits of holding in cases where mixing fails and of being potentially verifiable. Furthermore, the NED assumption allows for a variety of possible generating mechanisms for the squared cointegrating equilibrium error process. Davidson (2002) showed that many popular nonlinear models, including GARCH, are NED.[4] Davidson (2002) also proved that, subject to a suitable size of the underlying mixing and moment restrictions, NED is sufficient for the central limit theorem to hold. This theorem can be easily extended to the vector-valued NED case using the standard Cramér–Wold device, as shown in Appendix A. The central limit theorem is a key to derive the equal variance test. In order to give a more convenient statement in what follows, we first introduce the specific definition of NED.

Definition 1. (NED) (Davidson, 1994)

Let $z_t(\ldots,e_{t-1},e_t,e_{t+1},\ldots)$ denote a scalar random sequence, which is measurable function of the underlying process $\{e_s, -\infty < s < \infty\}$, where e_s is α-mixing of size $-r/(r-2)$ for $r > 2$ or ϕ-mixing of size $-r/(2r-2)$ for $r \geq 2$. Let $\mathcal{F}_s^t = \sigma(e_s,\ldots,e_t)$ be a sequence of σ-fields in which $s < t$. Denote E_{t-m}^{t+m} as the expectation conditional on $\mathcal{F}_{t-m}^{t+m}$. Then, z_t is said to be L_p-NED on a mixing process e_t for $p > 0$ if

$$\left\| z_t - E_{t-m}^{t+m}(z_t) \right\|_p \leq d_t v(m), \tag{3}$$

[3] For the definitions of mixing and martingale, see Chapter 3 of White (2001).

[4] The concept of NED on a mixing process was discussed and improved by Gallant and White (1988), Davidson (1992, 1993, 2001, 2004), and Davidson and De Jong (2000), Lu and Linton (2007), Li et al. (2012a,b), Qiu and Lin (2011), Jenish (2012), and Jenish and Prucha (2012), among others.

where d_t is a sequence of positive constants, and $v(m) \to 0$ as $m \to 0.z_t$ is said to be L_p-NED of size $-u$ if $v(m) = O(m^{-u-\varepsilon})$ for $\varepsilon > 0$.

Here, the process $\{e_t\}$ can be treated as the underlying deriving innovation process, which is assumed to satisfy the mixing condition. NED is a condition on the mapping from $\{e_t\}$ to z_t, and says nothing about the amount of dependence in the z_t series itself. The NED process can be approximated by a mixing input process in the NED sense, and does not require the assumption that the process itself is mixing. Of course, every mixing process is also NED on itself, and thus the class of process that are NED on a mixing includes the class of mixing process. Following Li et al. (2012b), a vector-valued NED can be introduced as follows.

Definition 2 (vector-form NED). Let $\mathbf{z}_t(\ldots,e_{t-1},e_t,e_{t+1},\ldots) = (z_{1t}(\ldots,e_{t-1},e_t,e_{t+1},\ldots),\ z_{2t}(\ldots,e_{t-1},e_t,e_{t+1},\ldots))'$ be a 2×1 random vector sequence, which is a measurable function of the underlying mixing process $\{e_s, -\infty < s < \infty\}$, where e_s is an α-mixing of size $-r/(r-2)$ for $r > 2$ or ϕ-mixing of size $-r/(2r-2)$ for $r \geq 2$. The process $\mathbf{z}_t = (z_{1t},z_{2t})'$ is said to be L_p-NED on a mixing process e_t for $p > 0$ if

$$\left\| z_{1t} - E_{t-m}^{t+m}(z_{1t}) \right\|_p + \left\| z_{2t} - E_{t-m}^{t+m}(z_{2t}) \right\|_p \leq d_t v(m), \tag{4}$$

where d_t is a sequence of positive constants, and $v(m) \to 0$ as $m \to 0$. $\mathbf{z}_t$ is said to be L_p-NED of size $-u$ if $v(m) = O(m^{-u-\varepsilon})$ for $\varepsilon > 0$.

To derive the asymptotic distributions of the proposed equal variance test, we need the following assumptions on the cointegrating equilibrium errors and their squared processes.

Assumption 1. The cointegrating equilibrium error process $\mathbf{u}_t = (\nu_t,\varsigma_t)'$ is stationary and ergodic with $E(\nu_t,\varsigma_t)' = (0,0)'$, $E(\nu_t^2,\varsigma_t^2)' = (\sigma_u^2,\sigma_\varsigma^2)'$.

Assumption 2. The squared cointegrating equilibrium error process $\mathbf{z}_t = (\nu_t^2 - \sigma_\nu^2, \varsigma_t^2 - \sigma_\varsigma^2)'$ is L_2-NED of size $-\frac{1}{2}$ on a process $\{e_t\}_{t=-\infty}^{\infty}$, where e_s is an α-mixing of size $-r/(r-2)$ for $r > 2$ or ϕ-mixing of size $-r/(2r-2)$ for $r \geq 2$. In addition, $\sup_t \|\mathbf{z}_t\|_r < \infty$.

These assumptions are general enough to enable us to apply the central limit theorem to the vector of squared cointegrating errors (i.e., $\mathbf{z}_t$) and accommodate a variety of possible processes of the squared cointegrating equilibrium errors, including GARCH. Assumption 1 enables us to employ the law of large numbers to estimate the variances of ν_t and ς_t and it also restricts ν_t^2 and ς_t^2 to be L_2-bounded. This in turn leads to the existence of the global variance of the squared cointegrating error in Eq. (2). Assumption 2 allows for the cross-sectional dependence and the auto-correlation of the squared cointegrating equilibrium errors.

2.2. The asymptotic distribution of the OLS estimation of squared cointegrating equilibrium errors

The matrix format of Eq. (2) in a sample of size T can be rewritten as

$$\begin{aligned} \mathbf{S} &= \mathbf{G}\theta_x + \nu, \\ \mathbf{S} &= \mathbf{H}\theta_y + \varsigma, \end{aligned} \tag{5}$$

where $\mathbf{S} = (S_{TW,1},S_{TW,2},\ldots,S_{TW,T})'$, $\mathbf{G} = (1,\mathbf{S}_{1t},\mathbf{t})_{t=1}^{T}$, $\mathbf{H} = (1,\mathbf{S}_{2t},\mathbf{t})_{t=1}^{T}$, $1 = (1,1,\ldots,1)$, $\mathbf{S}_{1t} = (S_{11},S_{12},\ldots,S_{1T})$, $\mathbf{S}_{2t} = (S_{21},S_{22},\ldots,S_{2T})$, $\mathbf{t} = (1,2,\ldots,T)$, $\theta_x = (\mu_x,\beta_x,\delta_x)'$, $\theta_y = (\mu_y,\beta_y,\delta_y)'$, $\nu = (\nu_1,\nu_2,\ldots,\nu_T)$, and $\varsigma = (\varsigma_1,\varsigma_2,\ldots,\varsigma_T)'$.

To identify the closer long-run relationship, following Lee et al. (2012), we compare the variances of ν_t and ς_t by constructing a variance test of the null hypothesis:

$$H_0 : \sigma_\nu^2 = \sigma_\varsigma^2,$$

against the alternative hypothesis,

$$H_1 : \sigma_\nu^2 \neq \sigma_\varsigma^2,$$

where σ_ν^2 and σ_ς^2 are the variances of ν_t and ς_t, respectively. The OLS estimates of parameters in Eq. (5) are super-consistent if cointegration relationships exist (see Stock, 1987). Therefore, we first examine the properties of OLS estimates of the regressions in the form of Eq. (5) under Assumptions 1 and 2 before constructing the test. After denoting the corresponding OLS estimates of the regression models in Eq. (5) as $\theta_x = \left[\hat{\mu}_x, \hat{\beta}_x, \hat{\delta}_x\right]'$ and $\theta_y = \left[\hat{\mu}_y, \hat{\beta}_y, \hat{\delta}_y\right]'$, we have

$$\begin{aligned} \mathbf{S} &= \mathbf{G}\hat{\theta}_x + \hat{\nu}, \\ \mathbf{S} &= \mathbf{H}\hat{\theta}_y + \hat{\varsigma}. \end{aligned} \tag{6}$$

The null distribution of the equal variance test can be derived from the joint distribution of the sum of squared estimated cointegrating equilibrium residuals. The following theorem shows that the joint distribution of sums of the OLS-estimated squared cointegrating equilibrium residuals is asymptotically standard normal under suitable normalization.

Theorem 1. *Define* $\hat{\sigma}_\nu^2 = \frac{\hat{\nu}'\hat{\nu}}{T-3} = \frac{\sum_{t=1}^T \hat{\nu}_t^2}{T-3}$ *and* $\hat{\sigma}_\varsigma^2 = \frac{\hat{\varsigma}'\hat{\varsigma}}{T-3} = \frac{\sum_{t=1}^T \hat{\varsigma}_t^2}{T-3}$. *Then, we have*

$$\mathbf{V}_T^{-1/2} T^{1/2} \left(\hat{\sigma}_\nu^2 - \sigma_\nu^2 \hat{\sigma}_\varsigma^2 - \sigma_\varsigma^2 \right) \xrightarrow{d} N(0, \mathbf{I}), \tag{7}$$

where $\mathbf{V}_T = \mathrm{Var}\left(T^{-1/2} \sum_{t=1}^T \mathbf{z}_t \right)$ *in which* $\mathbf{z}_t = (\nu_t^2 - \sigma_\nu^2, \varsigma_t^2 - \sigma_\varsigma^2)'$.

Proof of Theorem 1. See Appendix A.Theorem 1 shows that the joint distribution of the sums of OLS-estimated squared cointegrating equilibrium residuals is a multivariate normal distribution. The following result shows that the multivariate asymptotic normality can be transformed into univariate normality by a further transformation.

Theorem 2. *Define* $\sigma_m^2 = \mathrm{Var}(T^{-1/2}(\nu'\nu - \varsigma'\varsigma))$. *Then, we have*

$$\sigma_m^{-1} T^{1/2} \left[\left(\hat{\sigma}_\nu^2 - \hat{\sigma}_\varsigma^2 \right) - \left(\sigma_\nu^2 - \sigma_\varsigma^2 \right) \right] \xrightarrow{d} N(0, 1) \equiv W(1). \tag{8}$$

Proof of Theorem 2. See Appendix A.Theorem 2 clearly shows that the variance, σ_m^2, consists of three types of dependence: the cross-sectional dependence between ν_t^2 and ς_t^2 at the same point in time, the auto-dependence of ν_t^2 and ς_t^2, and the cross-dependence between ν_t^2 and ς_s^2 ($t \neq s$). The complicated dependence structure between ν_t^2 and ς_t^2 implies that it is difficult to estimate the variance, σ_m^2, directly and accurately. The following section presents this topic in greater detail.

2.3. The test statistics

The results of Theorem 2 can be used to test the null hypothesis, $H_0: \sigma_\nu^2 = \sigma_\varsigma^2$, if σ_m^2 can be consistently estimated. Lee et al. (2012) constructed a variance test by employing the HAC approach to estimate the variance σ_m^2 under the additional assumption that $\mathrm{Cov}(\nu_t^2, \varsigma_s^2) = 0$, $\forall$ t, s. However, this method requires the choice of a truncation lag (bandwidth). In practice, even if a data-dependent method is used to choose the bandwidth, arbitrary choices are inevitable, and these choices can produce inconclusive statistical inference.

Kiefer et al. (2000) proposed an alternative method of constructing test statistics strategy that does not require an estimate of the variance, σ_m^2. The key idea of this approach is to transform OLS estimates to construct t-type statistics in which the joint distribution of the transformed OLS estimates becomes asymptotically invariant to the serial correlation/heteroscedasticity nuisance parameters in σ_m^2. Thus, a robust test can be obtained. The method proposed in this paper can be used to construct the equal variance

Table 1
The critical values of the functionals of Browian motion: $\frac{\mathrm{W(1)}}{\left\{\int_0^1 [\mathrm{W(r)}\text{-}\mathrm{rW(1)}]^2 \mathrm{dr}^{1/2}\right\}}$.

T	1.00%	2.50%	5.00%	10.00%	15.00%	50.00%	90.00%	95.00%	97.50%	99.00%
25	−8.612	−6.796	−5.366	−3.877	−3.026	−0.017	3.881	5.338	6.792	8.596
50	−8.602	−6.753	−5.359	−3.869	−3.009	−0.010	3.893	5.367	6.822	8.586
100	−8.596	−6.758	−5.352	−3.867	−3.001	0.000	3.855	5.373	6.811	8.596
250	−8.595	−6.720	−5.365	−3.853	−3.000	0.010	3.884	5.368	6.767	8.551
500	−8.600	−6.761	−5.338	−3.853	−2.997	0.006	3.872	5.321	6.751	8.592
1000	−8.598	−6.748	−5.351	−3.864	−3.006	0.000	3.867	5.352	6.761	8.597
5000	−8.542	−6.715	−5.364	−3.865	−2.967	0.000	3.889	5.363	6.720	8.551

Notes: The Monte Carlo simulation is used to calculate the critical value of the equal variance test, based on 100,000 replications. The Wiener process, $W(r)$ is approximated by normalized sums of i.i.d. $N(0,1)$ pseudo random deviates. The simulations were written in the GAUSS programming language. The standard errors of critical values are reported in Appendix D which confirm that simulating 100,000 Monte-Carlo replications is statistically sufficient.

test using the same idea.[5] Instead of constructing a consistent estimate of the variance, σ_m^2, we eliminate σ_m^2 in the statistic, $\sigma_m^{-1}T^{1/2}\left[\left(\hat{\sigma}_\nu^2-\hat{\sigma}_\varsigma^2\right)-\left(\sigma_\nu^2-\sigma_\varsigma^2\right)\right]$, in Theorem 2 by a partial sum of the de-meaned-OLS squared residuals. This partial sum has a random limit that depends on σ_m^2. This result is stated in the following Lemma.

Lemma 1. *Define the partial sum* $\hat{S}_T(r)=\sum_{t=1}^{[Tr]}\left(\hat{\nu}_t^2-\hat{\varsigma}_t^2-\frac{1}{T}\sum_{s=1}^{T}\left(\hat{\nu}_s^2-\hat{\varsigma}_s^2\right)\right)$, *where* $0<r\le 1$. *Then*

$$\widehat{\mathcal{M}}^2=T^{-2}\sum_{[Tr]=1}^{T}\hat{S}_T^2(r)\xrightarrow{d}\sigma_m^2\int_0^1[W(r)-rW(1)]^2dr, \tag{9}$$

where W(r) is a standard Brownian motion.

Proof of Lemma 1. See Appendix A.The null hypothesis, $H_0:\sigma_\nu^2=\sigma_\varsigma^2$, can be tested using t-type statistics constructed as the usual t-statistics from Eq. (8), in which the usual standard error (σ_m) is replaced with the square root of $\hat{\mathcal{M}}^2$. In other words, consider the transformation $\hat{\mathcal{M}}^{-1}T^{1/2}\left[\left(\hat{\sigma}_\nu^2-\hat{\sigma}_\varsigma^2\right)-\left(\sigma_\nu^2-\sigma_\varsigma^2\right)\right]$, then it follows directly from Eqs. (8) and (9) that

$$\widehat{\mathcal{M}}^{-1}T^{1/2}\left[\left(\hat{\sigma}_\nu^2-\hat{\sigma}_\varsigma^2\right)-\left(\sigma_\nu^2-\sigma_\varsigma^2\right)\right]\xrightarrow{d}\frac{W(1)}{\left\{\int_0^1[W(r)-rW(1)]^2dr\right\}^{1/2}}. \tag{10}$$

Compared with Eq. (8), this transformation results in a limiting distribution in Eq. (10), where the nuisance parameter σ_m^2 is eliminated. The test based on Eq. (10) is likely to be more robust. The main result is summarized in the following theorem.

Theorem 3. *Under the null hypothesis of equal variance,* $H_0:\sigma_\nu^2=\sigma_\varsigma^2$, *the test statistic* $\mathcal{Z}=\widehat{\mathcal{M}}^{-1}T^{1/2}\left(\hat{\sigma}_\nu^2-\hat{\sigma}_\varsigma^2\right)$ *is asymptotically distributed as functionals of Brownian motion:*

$$\mathcal{Z}=\widehat{\mathcal{M}}^{-1}T^{1/2}\left(\hat{\sigma}_\nu^2-\hat{\sigma}_\varsigma^2\right)\xrightarrow{d}\frac{W(1)}{\left\{\int_0^1[W(r)-rW(1)]^2dr\right\}^{1/2}}. \tag{11}$$

[5] Assuming that the squared cointegrating equilibrium errors are adapted mixingale, Gao (2009) proposed a similar test statistic by defining $M^2=T^{-2}\sum_{[Tr]=1}^{T}S_T^2(r)$ in which $S_T(r)=\sum_{t=1}^{[Tr]}\left(\hat{\nu}_t^2-\hat{\varsigma}_t^2\right)$. However, this test is inconsistent. Specifically, we have shown that Gao's test converges to $\pm\sqrt{3}$ when the null of equal variance fails to hold. Hence, the power of the test is zero, even in large sample. The detailed proof is available from the authors upon request.

Proof of Theorem 3. By substituting the condition $\sigma_\nu^2 = \sigma_\varsigma^2$ into the left-hand side of Eq. (10), this result is obvious.

This new test statistic has a nonstandard asymptotic distribution and is free of nuisance parameters. The limiting distribution in Eq. (*11*) is identical to the distribution given by Eq. (*7*) in Kiefer et al. (2000). The critical values of the test represented by Eq. (*11*) can be computed using simulations, and are tabulated in Table 1. Finite-sample properties of the test and that of Lee et al. (2012) is provided in Appendix C.

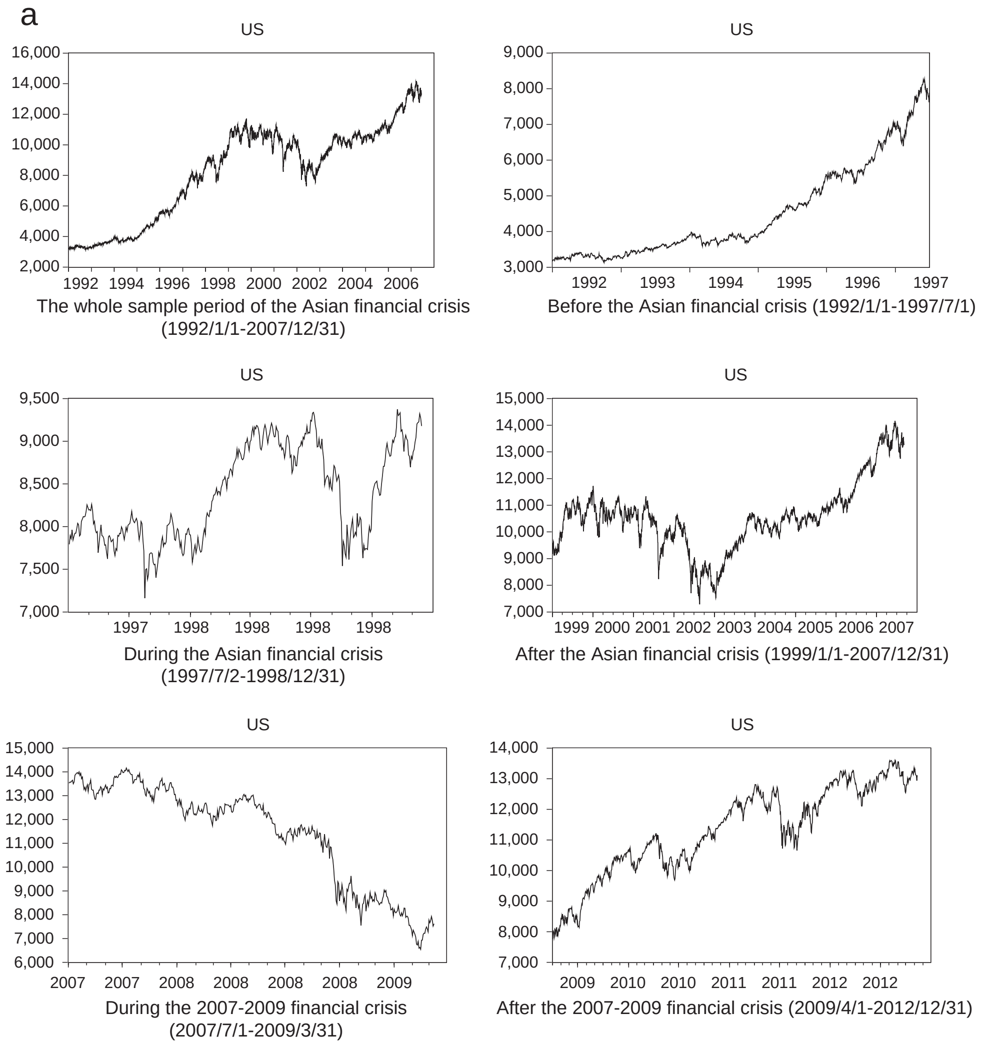


Fig. 2. (a) The US index (1992.1.1–2012.12.31). (b) Eight indexes over the whole sample period of the Asian financial crisis (1992.1.1–2007.12.31). (c) Eight indexes before the Asian financial crisis (1992.1.1–1997.7.1). (d) Eight indexes during the Asian financial crisis (1997.7.1–1998.12.31). (e) Eight indexes after the Asian financial crisis (1999.1.1–2007.12.31). (f) Eight indexes during the 2007–2009 financial crisis (2007.7.1–2009.3.31). (g) Eight indexes after the 2007–2009 financial crisis (2009.4.1–2012.12.31).

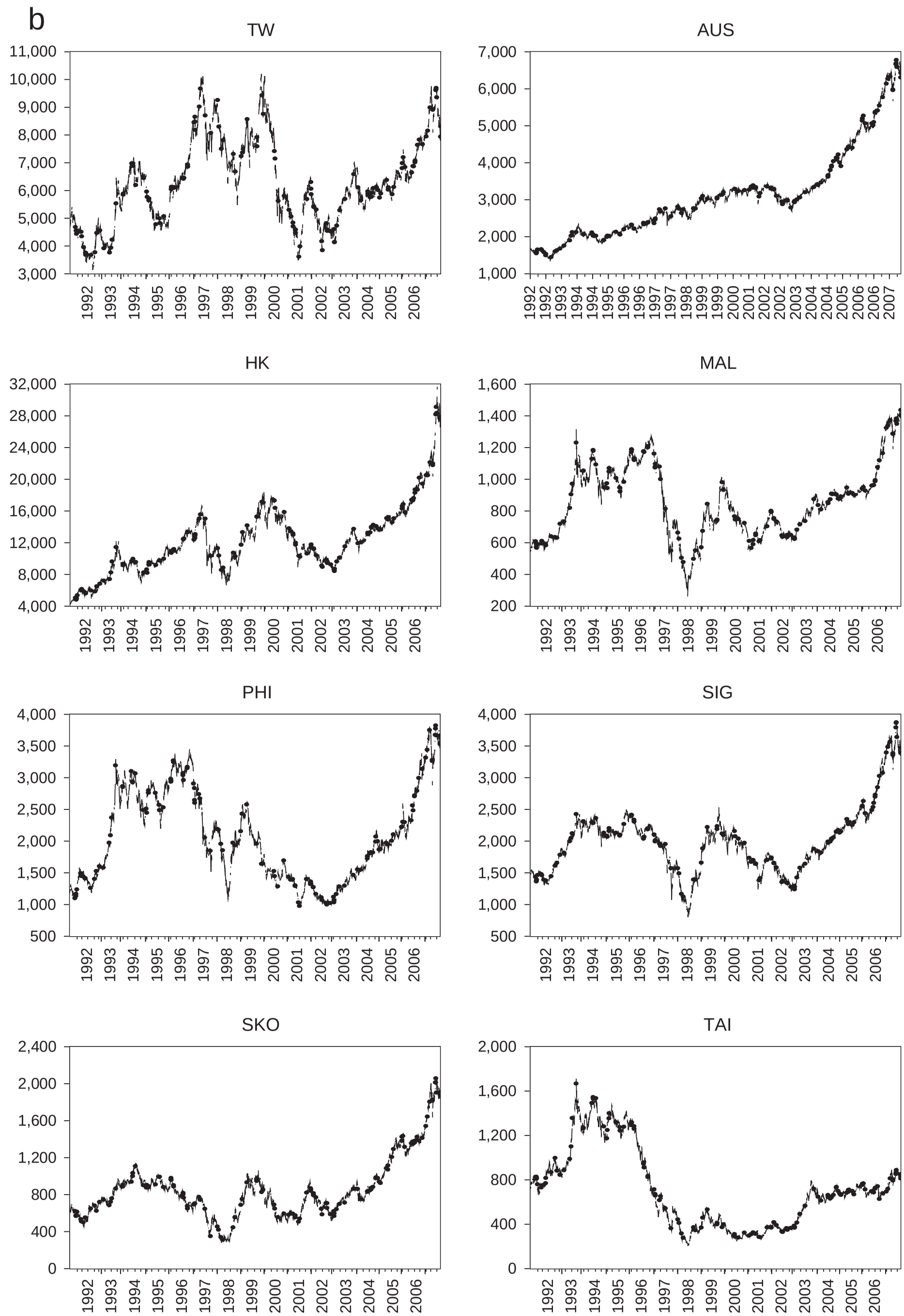
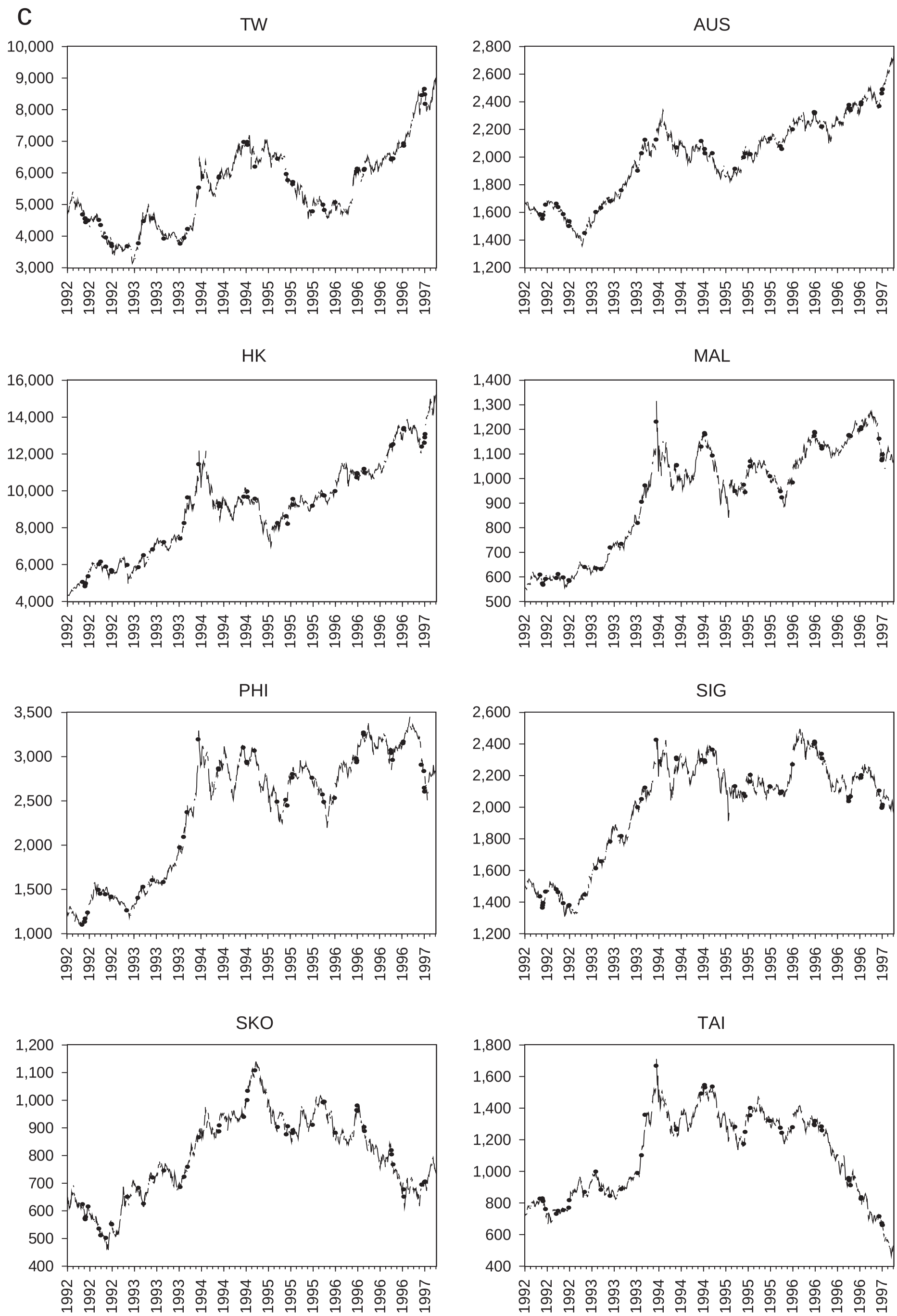


Fig. 2 (*continued*).



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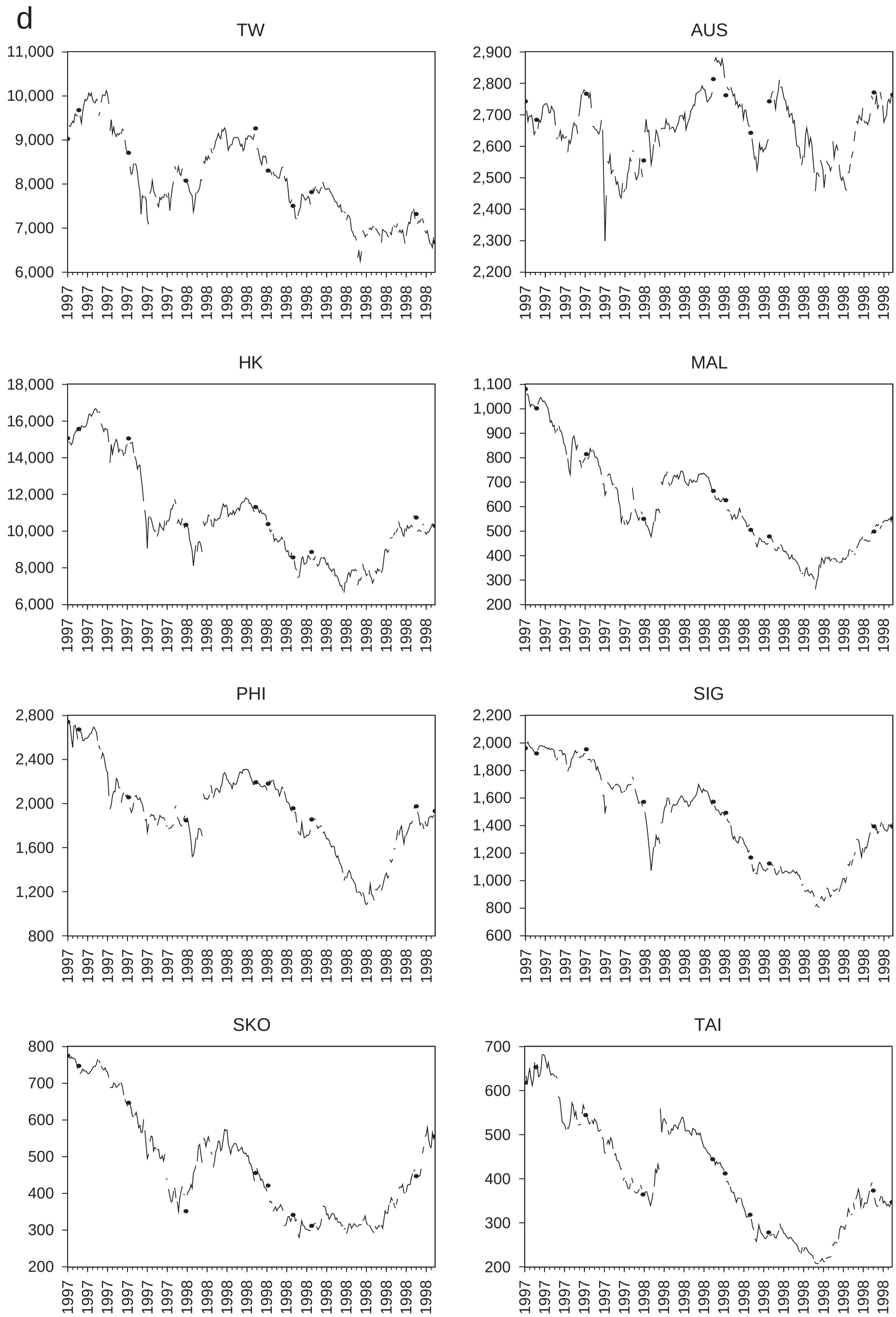

Fig. 2 (*continued*).

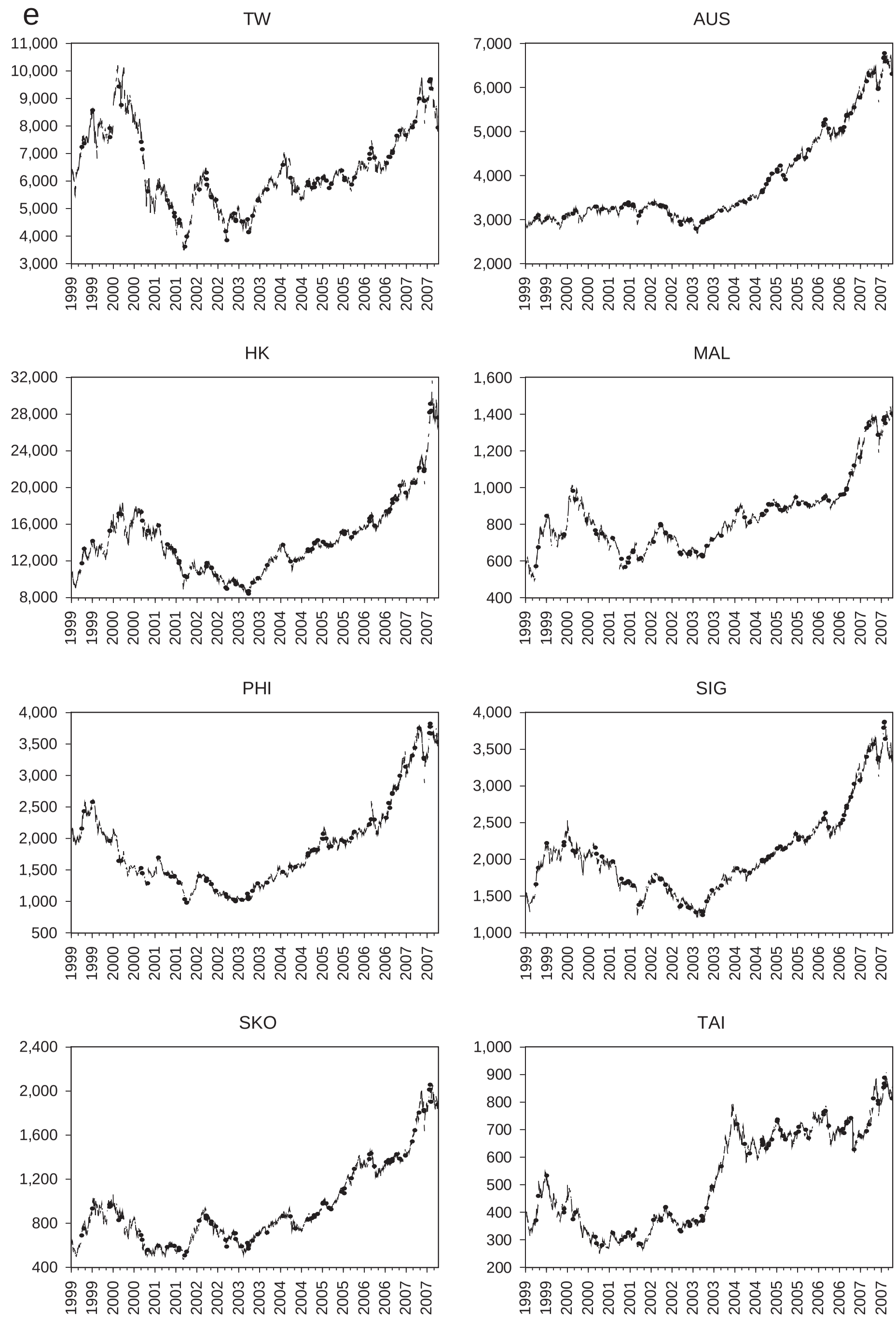

Fig. 2 (*continued*).

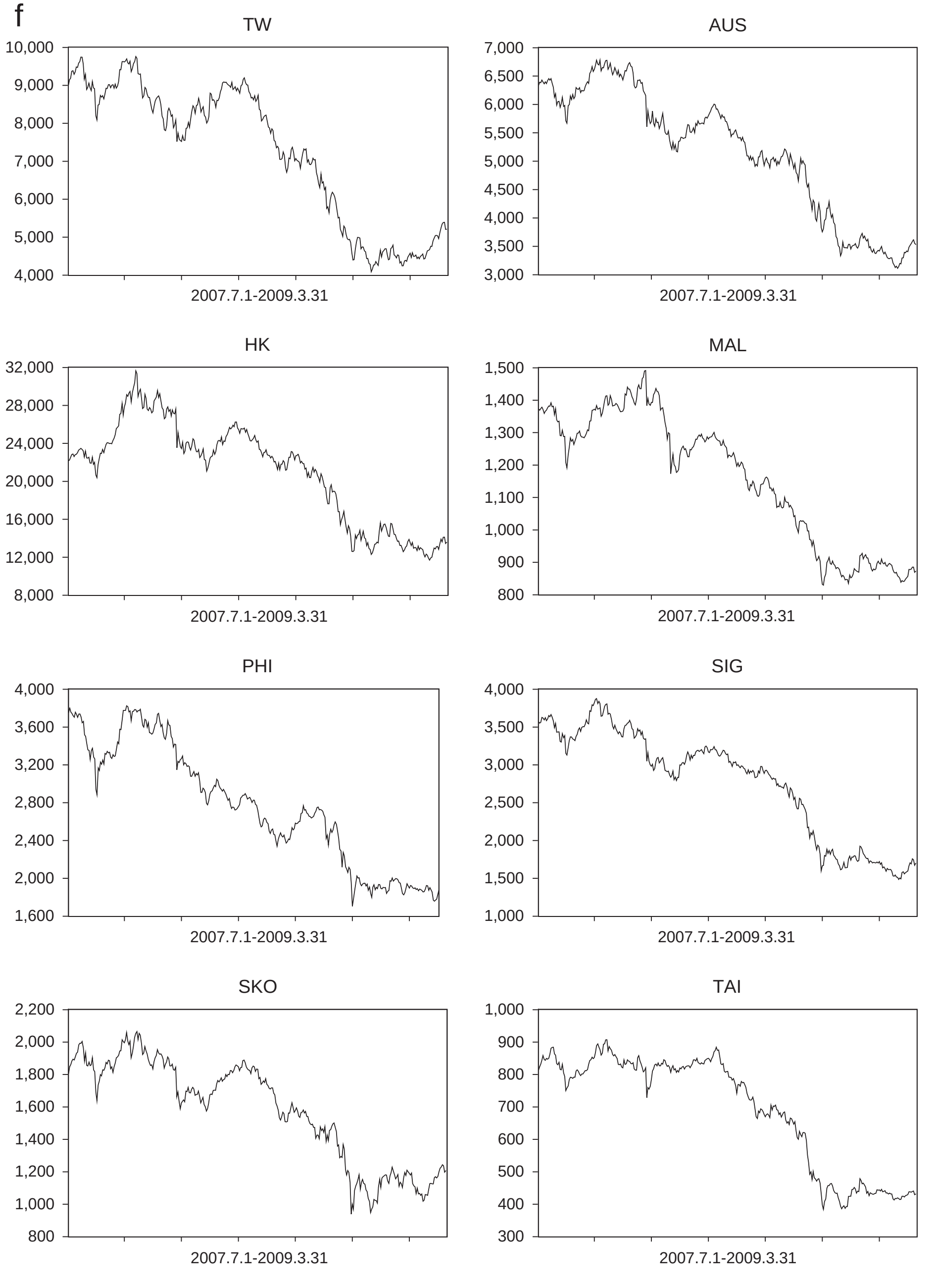

Fig. 2 (*continued*).

3. Empirical results

This section applies the variance test presented in Section 2 to investigate the comovement patterns and assess the degree of comovement between Taiwan and other countries. As argued by Lee et al. (2012), the main reason why Taiwan was chosen as the focus of this study is that much trade has taken place

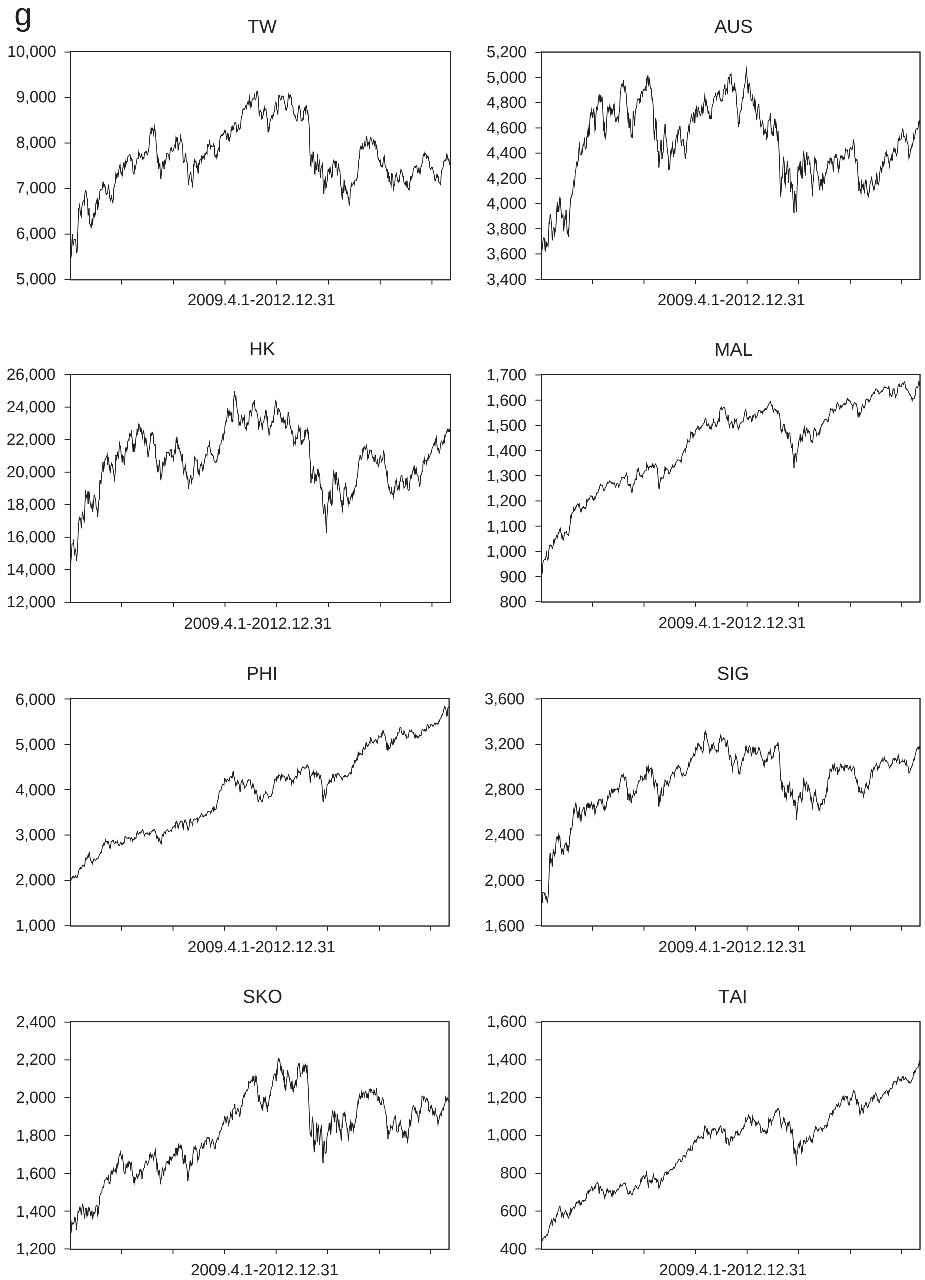

Fig. 2 (*continued*).

between Taiwan and countries (e.g., the United States) other than those in the Asian region over the past several decades. By examining the long-run relationships among stock markets, including the United States market and other Asian markets, this study investigates the factors behind the international stock market comovements. Specifically, we attempt to answer the following question: Is this comovement because of the adjacent region and the similarities of background of the capital markets, or the international trade and business cooperation, or both?

3.1. Data and unit root tests

The data analyzed in this study comprised of daily closing share price indexes of the eight East Asian stock indices: 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 index for Thailand (TAI), and the Taiwan Stock Exchange weighted price index for Taiwan (TW) over the period January 1, 1992 to December 31, 2007. The Dow Jones Industrial Average Index was also included in this analysis, because of the widely-held view of the leadership of the United States and the trading activities and business cooperation between Taiwan and the United States. The data was sourced from the Datastream database, and the sample period was divided into the pre-crisis stage (1992/1/1–1997/7/1), crisis stage (1997/7/2–1998/12/31), and post-crisis stage (1999/1/1–2007/12/31). The sample period and countries in the sample were the same as those considered by Lee et al. (2012) to enable a comparison with their empirical results. To check the robustness of our results, we then update the data to end at December 31, 2012 in Section 3.4. This could extend our analysis to include the recent global financial crisis as well. Fig. 2 in Appendix E shows a graphical depiction of these chosen stock price indexes.

The following paragraphs present a series of unit root tests prior to testing for cointegration. We employ the DF-GLS test suggested by Elliott et al. (1996) to test for stationarity of stock indices. The results of the whole sample and the sub-samples show that the null hypothesis of a unit root cannot be rejected in all stock prices at the 5% level of significance. After first differencing the variables, we applied the DF-GLS test to test for the order of integration for each stock prices over the full sample and sub-samples. Like similar findings in the literature (Kasa, 1992; Masih and Masih, 2001), each of these stock prices is stationary in their first-differences. Thus, it might make sense to conclude that all stock prices are integrated of order one, $I(1)$, for the full sample period and three sub-sample periods.

3.2. Cointegration tests

The Engle–Granger two-step method and Johansen–Juselius procedure are now well-known in cointegration analysis, of which the Engle–Granger procedure has advantage of being intuitive and simple. After confirming the common integration properties of the stock prices, we therefore apply the Engle and Granger (1987) cointegration test to test the presence of long-run equilibrium of non-stationary series between the indexes of the Taiwan stock market and seven other markets. Cointegration means a possible allowance for departure from the equilibrium in the short run, but not in the long run. In terms of investment, the existence of the cointegration relationships among international stock markets may limit the benefit of international diversification. Thus, the results displayed in Table 2 imply that diversifying

Table 2
Engle–Granger co-integration test with Taiwan.

period	test	AUS	HK	MAL	PHI	SIG	SKO	TAI	US
All	PP	−2.427*	−2.549*	−2.284	−2.458*	−2.269	−2.261	−2.254	−2.815**
	ADF	−2.350	−2.561*	−2.255	−2.481*	−2.237	−2.193	−2.149	−2.812**
Pre-crisis	PP	−1.578	−1.814	−1.022	−0.977	−1.178	−1.244	−1.479	−1.450
	ADF	−1.533	−1.749	−0.988	−0.987	−1.114	−1.181	−1.386	−1.387
During the crisis	PP	−3.638***	−2.379	−2.825**	−3.344**	−2.119	−2.339	−2.607*	−3.217**
	ADF	−3.416***	−2.379	−2.851**	−3.361**	−2.155	−2.367	−2.654*	−3.371***
Post-crisis	PP	−2.563*	−2.540*	−3.669***	−3.053**	−3.283**	−3.253**	−1.984	−3.392***
	ADF	−2.499*	−2.600*	−3.771***	−3.093**	−3.378**	−3.215**	−2.137	−3.478***

Notes: ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively. The standard PP and ADF unit root tests with no constant or time trend are employed to complete the Engle-Granger procedure. The critical values of residual based unit root tests are from Phillips and Ouliaris (1990).

across national markets might be very beneficial before a financial crisis, but not after or during. The international diversification between Taiwan and other markets in the long-run is also relatively beneficial. Following the residual-based Engle–Granger two-step approach, we first use the OLS to estimate the regressions in the form of Eq. (2). Except for the intercept for Australia during the crisis, all the coefficients of the regression model in Eq. (2) are significant.[6] These results are consistent with these reported by Lee et al. (2012). Following the results of Phillips and Ouliaris (1990) for residual-based cointegration tests, we employ the PP test and ADF test to detect the presence of cointegration. When the unit root tests are applied to the residuals from a spurious regression, the critical values differ from those employed for testing a unit root in raw time series (Phillips and Ouliaris, 1990; Hamilton, 1994).

Table 2 presents the results of the Engle–Granger cointegration test. Given the integrational properties of the residuals, the PP and ADF statistics in this study can tell us if a cointegration relationship exists for the null hypothesis of unit root in the residual being rejected. These results show that the contegration relationship may change in different periods of a country sample. For the full sample period, the cointegrating relationships only occur between Taiwan and half of the eight capital markets. Both of the unit root tests appeared to be robust for sub-periods samples. No cointegration relationship appears across the pair-wise countries before the crisis. Furthermore, there are cointegration relationships between Taiwan and almost all of the other markets after the crisis, whereas five countries exhibit comovement with Taiwan. These findings regarding changing cointegration relationships have several implications. First, the number of cointegration relationships may increase abruptly during the crisis, and become prolonged in the post-crisis period. Thus, it appears that the linkages among stock markets were strengthened during and after the crisis, which is highly consistent with the results of Yang et al. (2006). It is related to the famous contagion effect because investors associate it with risk in thinking, especially for emerging markets. Second, some type of relationship may exist only during the crisis, but not before or after the crisis, because of uncertainty economic circumstances. For example, the comovement between Taiwan and Thailand markets is only founded during the crisis. Finally, the cointegration relationships between Taiwan and Singapore and South Korea exist only after the crisis, and not before or during the crisis. Thus, we infer that stock markets may be partially tied together by economic globalization or international trade.

The cointegration results in Table 2 confirms that the linkages among stock markets are strengthened after the crisis, which is consistent with the results of Arshanapalli and Doukas (1993), Yang et al. (2003), and Yang et al. (2006). However, other studies have presented conflicting conclusions. For example, King and Wadhwani (1990) and King et al. (1994) showed that the correlation between national stock market returns increases only temporarily during a financial crisis. Tuluca and Zwick (2001) showed that the Asian financial crisis had temporary strengthening effects on global equity market relationships. Chen et al. (2002) concluded that the Asian financial crisis and the Russian crisis did not have a dramatic effect on the interdependence across Latin American stock markets, and the long-run cointegration relationship disappeared in the period following the Russian crisis. Manning (2002) showed that the convergence process in Asian emerging markets was abruptly halted and somewhat reversed by the Asian financial crisis in 1997. Faccio and Parsley (2009) suggested that the apparent persistency of the increase in stock market comovement after the beginning of the Asian crisis is actually the result of a series of periodic increases in comovement after short-lived shocks, rather than a long-term post-crisis effect.

3.3. The equal variance test

The cointegration analysis seeks to detect the existence of cointegration, but cannot discriminate the closer relationship from cointegrating relationships. In Section 2 we propose a robust equal variance test to measure the degree of cointegration.

As Lee et al. (2012) emphasized, the equal variance test clarifies the cointegration relationship by assessing the closeness of the linkage between two variables. Recall that rejecting the null hypothesis of equality of variances enables us to conclude that smaller variances among the equilibrium errors exhibit closer linkages between the variables in the cointegrating regression models compared to larger variances

[6] The detailed contents can be found in Tables 2 and 3 in Lee et al. (2012).

among the equilibrium errors. Therefore, the equal variance test is employed to distinguish the different levels of the comovement among stock markets.

Unlike the variance test proposed by Lee et al. (2012), which assumes the uncorrelated squared cointegrating equilibrium errors between a pair-wise of variables, the new variance test proposed in this paper accommodates the cross-sectional dependence between the squared cointegrating errors. The simulations of variance tests in Table 6 in Appendix C are adequate for the cases of cross-sectionally dependent/independent cointegrating equilibrium errors. The test proposed by Lee et al. (2012) might spuriously reject the null and decrease its power to detect the closer relationship under cross-sectional dependent case, whereas the equal variance test has the correct size and a larger power. Because the cross-dependence between the squared cointegrating errors is suitable for the real world, re-evaluating the linkages between stock markets by applying the new variance test proposed in this paper enables us to avoid misleading conclusions when assessing cointegrating relationships.

Table 3 shows the results of the variance test described in Theorem 3. First, the tests were applied with the full period in Panel A, followed by the same tests in other sub-periods in Panels B, C, and D. By rejecting the null hypothesis that the variances formed by each cointegration regression are equal, we conclude that the closer linkages between the variables in cointegration regression models are the result of the smaller variance among the disturbances. The positive (minus) number in Table 3 indicates that the variances of cointegrating errors between the countries in the left-most column are larger (smaller) than those in the corresponding upper row. For example, the −3.962 value in the third column of Panel A in Table 3 means that we can reject the null $\hat{\sigma}^2_{TW\_HK} = \hat{\sigma}^2_{TW\_PHI}$ at the 10% significance level. Therefore there is a closer cointegration relationship between Taiwan and Hong Kong than that between Taiwan and the Philippines.

Over the full sample period, the relationships between Taiwan and Thailand have the weakest connection, as shown by the all positive signs in the bottom row in Panel A, Table 3. The second-farthest relationship is between Taiwan and South Korea.[7]

Panels B, C, and D in Table 3 present the results for the pre-crisis, crisis, and post-crisis periods, separately. In Panel B, all the numbers are insignificant and we cannot reject the null hypothesis that the two sets of squared cointegrating errors are equal. These results are consistent with those of Lee et al. (2012).

The results in Panel C indicate that the relationship between Taiwan and the Philippines is closest during the crisis period. Next, the relationships between the stock prices of Taiwan and Australia and that between Taiwan and the United States are likely to be closer than those between Taiwan and other countries during the crisis. The existence of a relatively close relationship between Taiwan and the United States compared to that between Taiwan and other countries excluding Australia and Philippines during the crisis tends to confirm the widely-held view of the leadership of the United States, which is consistent with many previous studies. For example, Arshanapalli and Doukas (1993) found the United States stock market had a considerable effect on the French, German, and UK markets in the post-crash period. Masih and Masih (1999) confirmed the dominant role of the United States over both the short- and long-term, and the existence of a significant short- and long-term relationship between the established OECD and the emerging Asian markets. Bessler and Yang (2003) further showed that the United States market is the only market that has a consistently strong impact on price movements in other major stock markets in the long-run. In comparison, the test presented by Lee et al. (2012) cannot detect any close relationships for this stage, and therefore fails to assess how far and how close the relationships are among stock markets during the crisis period.

The results in Panel D show that there is no closer relationship between Taiwan and other countries, except for Thailand, in the post-crisis period. Thus, the degree of the comovement between Taiwan and each stock market considered seems to be similar.

Compared to the empirical results in Lee et al. (2012), this study arrives at the same conclusion that the relationship between Taiwan and Thailand is special because the variances of the residuals of the regression model as expressed in Eq. (2) are surprisingly large over the entire sample and post-crisis

[7] Similarly, Valadkhani and Chancharat (2008) found no evidence of long-run relationships between Thailand and Taiwan after considering structural breaks. The closest relationship appears between Taiwan and Hong Kong. Lee et al. (2012) showed a closer relationship between Taiwan and the United States than between Taiwan and Philippines for the full sample, but this finding not supported in this paper.

Table 3
The equal variance tests between different countries based on Taiwan.

Panel A: All 1992/1/1–2007/12/31

	$\hat{\sigma}^2_{TW\_HK}$	$\hat{\sigma}^2_{TW\_US}$	$\hat{\sigma}^2_{TW\_PHI}$	$\hat{\sigma}^2_{TW\_AUS}$	$\hat{\sigma}^{2\ a}_{TW\_SIG}$	$\hat{\sigma}^{2\ a}_{TW\_MAL}$	$\hat{\sigma}^{2\ a}_{TW\_SKO}$	$\hat{\sigma}^{2\ a}_{TW\_TAI}$
$\hat{\sigma}^2_{TW\_HK}$	0.000	−1.504	−3.962*	−5.862**	−5.806**	−6.739**	−8.811***	−7.116**
$\hat{\sigma}^2_{TW\_US}$	1.504	0.000	−1.328	−2.447	−2.404	−3.095	−3.922*	−7.853**
$\hat{\sigma}^2_{TW\_PHI}$	3.962*	1.328	0.000	−2.227	−3.403	−7.692**	−7.752**	−4.421*
$\hat{\sigma}^2_{TW\_AUS}$	5.862**	2.447	2.227	0.000	−0.537	−0.761	−3.024	−1.703
$\hat{\sigma}^{2\ a}_{TW\_SIG}$	5.806**	2.404	3.403	0.537	0.000	−0.689	−4.198**	−1.535
$\hat{\sigma}^{2\ a}_{TW\_MAL}$	6.739**	3.095	7.692**	0.761	0.689	0.000	−2.837	−1.781
$\hat{\sigma}^{2\ a}_{TW\_SKO}$	8.811***	3.922*	7.752**	3.024	4.198*	2.837	0.000	−0.738
$\hat{\sigma}^{2\ a}_{TW\_TAI}$	7.116**	7.853**	4.211*	1.703	1.535	1.781	0.738	0.000

Panel B: Pre-crisis 1992/1/4–1997/7/1

	$\hat{\sigma}^{2\ a}_{TW\_PHI}$	$\hat{\sigma}^{2\ a}_{TW\_HK}$	$\hat{\sigma}^{2\ a}_{TW\_MAL}$	$\hat{\sigma}^{2\ a}_{TW\_AUS}$	$\hat{\sigma}^{2\ a}_{TW\_TAI}$	$\hat{\sigma}^{2\ a}_{TW\_SIG}$	$\hat{\sigma}^{2\ a}_{TW\_SKO}$	$\hat{\sigma}^{2\ a}_{TW\_US}$
$\hat{\sigma}^{2\ a}_{TW\_PHI}$	0.000	−0.076	−0.343	−0.385	−0.883	−2.031	−2.101	−0.958
$\hat{\sigma}^{2\ a}_{TW\_HK}$	0.076	0.000	−0.008	−0.591	−3.719	−1.376	−1.315	−3.003
$\hat{\sigma}^{2\ a}_{TW\_MAL}$	0.343	0.008	0.000	−0.366	−1.014	−2.948	−2.941	−1.106
$\hat{\sigma}^{2\ a}_{TW\_AUS}$	0.385	0.591	0.366	0.000	−1.570	−1.426	−1.337	−2.380
$\hat{\sigma}^{2\ a}_{TW\_TAI}$	0.883	3.719	1.015	1.570	0.000	−0.004	−0.005	−0.056
$\hat{\sigma}^{2\ a}_{TW\_SIG}$	2.031	1.376	2.948	1.426	0.004	0.000	−0.018	−0.011
$\hat{\sigma}^{2\ a}_{TW\_SKO}$	2.101	1.315	2.941	1.337	0.005	0.018	0.000	−0.008
$\hat{\sigma}^{2\ a}_{TW\_US}$	0.958	3.003	1.106	2.380	0.056	0.011	0.008	0.000

Panel C: During the crisis 1997/7/7-1998/12/31

	$\hat{\sigma}^2_{TW\_PHI}$	$\hat{\sigma}^2_{TW\_AUS}$	$\hat{\sigma}^2_{TW\_US}$	$\hat{\sigma}^2_{TW\_MAL}$	$\hat{\sigma}^2_{TW\_TAI}$	$\hat{\sigma}^{2\ a}_{TW\_HK}$	$\hat{\sigma}^{2\ a}_{TW\_SIG}$	$\hat{\sigma}^{2\ a}_{TW\_SKO}$
$\hat{\sigma}^2_{TW\_PHI}$	0.000	−1.080	−1.910	−7.254**	−5.854**	−5.370**	−3.714	−5.401**
$\hat{\sigma}^2_{TW\_AUS}$	1.080	0.000	−1.756	−1.944	−2.997	−3.135	−6.430**	−5.703**
$\hat{\sigma}^2_{TW\_US}$	1.910	1.756	0.000	−1.561	−3.088	−3.140	−5.089*	−6.082**
$\hat{\sigma}^2_{TW\_MAL}$	7.254**	1.944	1.561	0.000	−2.334	−2.586	−2.450	−3.967
$\hat{\sigma}^2_{TW\_TAI}$	5.854**	2.997	3.088	2.334	0.000	−1.620	−2.256	−3.687
$\hat{\sigma}^{2\ a}_{TW\_HK}$	5.370**	3.135	3.140	2.586	1.620	0.000	−1.030	−3.026
$\hat{\sigma}^{2\ a}_{TW\_SIG}$	3.714	6.430**	5.089*	2.450	2.256	1.030	0.000	−0.296
$\hat{\sigma}^{2\ a}_{TW\_SKO}$	5.401**	5.703**	6.082**	3.967*	3.687	3.026	0.296	0.000

Panel D: Post-crisis 1999/1/1–2007/12/31

	$\hat{\sigma}^2_{TW\_MAL}$	$\hat{\sigma}^2_{TW\_SIG}$	$\hat{\sigma}^2_{TW\_SKO}$	$\hat{\sigma}^2_{TW\_HK}$	$\hat{\sigma}^2_{TW\_PHI}$	$\hat{\sigma}^2_{TW\_US}$	$\hat{\sigma}^2_{TW\_AUS}$	$\hat{\sigma}^{2\ a}_{TW\_TAI}$
$\hat{\sigma}^2_{TW\_MAL}$	0.000	−0.841	−1.986	−2.194	−1.684	−1.735	−2.527	−9.774***
$\hat{\sigma}^2_{TW\_SIG}$	0.841	0.000	−1.622	−4.322*	−2.220	−2.407	−3.648	−6.294**
$\hat{\sigma}^2_{TW\_SKO}$	1.986	1.622	0.000	−1.790	−1.435	−1.484	−2.640	−6.673**
$\hat{\sigma}^2_{TW\_HK}$	2.194	4.322*	1.790	0.000	−0.708	−0.947	−2.754	−4.337*
$\hat{\sigma}^2_{TW\_PHI}$	1.684	2.220	1.435	0.708	0.000	−0.167	−3.842	−2.907
$\hat{\sigma}^2_{TW\_US}$	1.735	2.407	1.484	0.947	0.167	0.000	−6.198**	−2.709
$\hat{\sigma}^2_{TW\_AUS}$	2.527	3.648	2.640	2.754	3.842	6.198**	0.000	−1.657
$\hat{\sigma}^{2\ a}_{TW\_TAI}$	9.774***	6.294**	6.673**	4.337*	2.907	2.709	1.657	0.000

Notes: This table displays the test statistics, Z, constructed in Theorem 3, which compares the magnitude of the variances of cointegrating equilibrium errors, hence assess how close and how far the relationships are. The variances of disturbances can be estimated as follows: $\hat{\sigma}^2_{TW\_Si} = \frac{1}{T-3}\sum_{t=1}^{T}\left(S_{TW,t}-\hat{\mu}-\hat{\beta}S_{it}-\hat{\delta}t\right)^2$, where S_i is one of the 8 indices. Smaller variances of disturbances exhibit closer linkages between the independent and dependent variables in the cointegration regression model as opposed to larger variances of disturbances. The superscript, [a], indicates that the rows or columns have non-stationary disturbances. The indexes are the Taiwan Stock Exchange weighted price index for Taiwan (TW) and those of 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), and the New York Dow Jones industrial average for the United States (US). In addition, ***, **,and * denote statistical significance at the 1%, 5%, and 10% levels, respectively.

periods. However, a cointegrating relationship appears during the crisis. Unlike Lee et al. (2012), we cannot observe a closer relationship between Taiwan and Malaysia than between Taiwan and other countries for the post-crisis period. Therefore, unlike Lee et al. (2012), we cannot reach the conclusion that adjacent regions with similar capital markets will exhibit price patterns that are more similar to those of Taiwan than those of countries with which Taiwan frequently trades or cooperates. In particular, we find a closer relationship between Taiwan and the United States than between Taiwan and other countries. This finding confirms to commonly held impression that the stock markets of the United States and Taiwan are strongly linked. These findings appear to support that similarities in background and trading activities are crucial factors in the price patterns, which is consistent with Madaleno and Pinho (2012). Nevertheless, this result should be interpreted carefully because there are other potential explanations. For example, a large export share by Asian economies to the United States may contribute to comovement among Asian stock markets, which might make geographical ties unfavorable and therefore cannot be excluded. Overall, these results show that, after considering the cross-sectional dependence between squared cointegrating errors, the evidence supporting a closer cointegration relationship between capital markets based on adjacent regions on statistical grounds is much weaker than suggested by Lee et al. (2012). According to the finite-sample properties of the variance tests in Appendix C, failure to control size and low power may explain the empirical results of Lee et al. (2012).

3.4. Robustness analysis

The sample period of previous analysis was restricted to January 1, 1992 to December 31, 2007, which enabled us to compare our results with those of Lee et al. (2012). To verify the robustness of the empirical results and understand the recent financial crisis, we extended the sample period of the analysis to July 1, 2007 to December 31, 2012. The Financial Crisis of 2007–2009 might have been the greatest shock to the United States and worldwide financial systems since the 1930s. We therefore focus on the 2007–2009 crisis in this section. Based on the analysis of Milesi-Ferretti and Tille (2011) and Ben-David et al. (2012), we divided the sample period into: a crisis period (2007/7/1–2009/3/31), which began with the Quant Meltdown in the summer of 2007 and ended with the trough of the stock market in March 2009, and a post-crisis period (2009/4/1–2012/12/31).

Similarly to the previous analysis, all stock prices were confirmed to be integrated of order one, $I(1)$, for the sub-sample periods. By using the residual-based Engle–Granger two-step approach, we first estimated the regressions in the form of Eq. (2). The results of the Engle–Granger cointegration test are reported in Table 4.

Different co-movement relationships were observed both during and after the 2007–2009 financial crisis. Regarding the post-crisis period (2009/4/1–2012/12/31), the co-integrating relationships between Taiwan and other countries could not be rejected, but most of these relationships could not be observed in the crisis period (2007/7/1–2009/3/31). To explore additional implications of this recent crisis, we measured the similarity of capital markets by conducting the variance tests.

Table 5 presents the empirical results of the variance tests for the crisis period (2007/7/1–2009/3/31) and the post-crisis period (2009/4/1–2012/12/31). Panel B in Table 5 shows that the variance test of Lee et al. (2012) indicates a closer relationship between Taiwan and South Korea than between Taiwan and the United States at the 1% level of significance, which might be associated with geographical ties.

Table 4
Engle–Granger co-integration test with Taiwan in the period 2007/7/1–2012/12/31.

Period	test	AUS	HK	MAL	PHI	SIG	SKO	TAI	US
(2007/7/1–2009/3/31)	PP	−1.67	−1.54	−1.79	−1.82	−1.96	−2.87**	−2.81**	−2.54*
During the crisis	ADF	−1.66	−1.59	−1.85	−1.84	−1.99	−2.86**	−3.00**	−2.32
(2009/4/1–2012/12/31)	PP	−2.91**	−3.09**	−6.67***	−4.33***	−3.26**	−4.53***	−7.96***	−4.14***
Post-crisis	ADF	−2.82**	−3.14**	−5.01***	−2.88**	−3.61***	−4.57***	−4.15***	−3.88***

Notes: ***, **, and * denote statistical significance at the 1%, 5%, and 10% levels, respectively. The standard PP and ADF unit root tests with no constant or time trend are employed to complete the Engle–Granger procedure. The critical values of residual based unit root tests are from Phillips and Ouliars (1990).

Table 5
The variance tests between different countries based on Taiwan in the period 2007/7/1–2012/12/31.

Panel A: The equal variance test proposed in the current paper: during the crisis 2007/7/1–2009/3/31

	$\hat{\sigma}^2_{TW\_TAI}$	$\hat{\sigma}^2_{TW\_SKO}$	$\hat{\sigma}^2_{TW\_SIG}{}^{a}$	$\hat{\sigma}^2_{TW\_US}$	$\hat{\sigma}^2_{TW\_HK}{}^{a}$	$\hat{\sigma}^2_{TW\_AUS}{}^{a}$	$\hat{\sigma}^2_{TW\_MAL}{}^{a}$	$\hat{\sigma}^2_{TW\_PHI}{}^{a}$
$\hat{\sigma}^2_{TW\_TAI}$	0.000	−2.284	−5.079*	−3.934*	−8.224**	−7.361**	−5.604**	−6.746**
$\hat{\sigma}^2_{TW\_SKO}$	2.284	0.000	−5.419**	−2.956	−9.812***	−10.012***	−6.037**	−7.104**
$\hat{\sigma}^2_{TW\_SIG}{}^{a}$	5.079*	5.419**	0.000	−1.623	−4.065*	−7.733**	−4.228*	−7.052**
$\hat{\sigma}^2_{TW\_US}$	3.934*	2.956	1.623	0.000	−0.015	−0.638	−1.092	−3.386
$\hat{\sigma}^2_{TW\_HK}{}^{a}$	8.224**	9.812***	4.065*	0.015	0.000	−2.063	−3.142	−4.789*
$\hat{\sigma}^2_{TW\_AUS}{}^{a}$	7.361**	10.012***	7.733**	0.638	2.063	0.000	−1.794	−5.046*
$\hat{\sigma}^2_{TW\_MAL}{}^{a}$	5.604**	6.037**	4.228*	1.092	3.142	1.794	0.000	−3.151
$\hat{\sigma}^2_{TW\_PHI}{}^{a}$	6.746**	7.104**	7.052**	3.386	4.789*	5.046*	3.151	0.000

Panel B: The variance test proposed by Lee et al. (2012): during the crisis 2007/7/1–2009/3/31

	$\hat{\sigma}^2_{TW\_TAI}$	$\hat{\sigma}^2_{TW\_SKO}{}^{a}$	$\hat{\sigma}^2_{TW\_SIG}{}^{a}$	$\hat{\sigma}^2_{TW\_US}$	$\hat{\sigma}^2_{TW\_HK}{}^{a}$	$\hat{\sigma}^2_{TW\_AUS}{}^{a}$	$\hat{\sigma}^2_{TW\_MAL}{}^{a}$	$\hat{\sigma}^2_{TW\_PHI}{}^{a}$
$\hat{\sigma}^2_{TW\_TAI}$	0.000	−0.698	−1.993**	−2.921***	−2.856***	−3.054***	−2.966***	−3.714***
$\hat{\sigma}^2_{TW\_SKO}{}^{a}$	0.698	0.000	−1.337*	−2.341***	−2.296**	−2.556***	−2.594***	−3.432***
$\hat{\sigma}^2_{TW\_SIG}{}^{a}$	1.993**	1.337*	0.000	−1.097	−1.087	−1.451*	−1.731**	−2.758***
$\hat{\sigma}^2_{TW\_US}$	2.921***	2.341**	1.097	0.000	−0.012	−0.434	−0.888	−2.069**
$\hat{\sigma}^2_{TW\_HK}{}^{a}$	2.856***	2.296**	1.087	0.012	0.000	−0.417	−0.871	−2.048**
$\hat{\sigma}^2_{TW\_AUS}{}^{a}$	3.054***	2.556***	1.451*	0.434	0.417	0.000	−0.500	−1.720**
$\hat{\sigma}^2_{TW\_MAL}{}^{a}$	2.966***	2.594***	1.731**	0.888	0.871	0.500	0.000	−1.208
$\hat{\sigma}^2_{TW\_PHI}{}^{a}$	3.714***	3.432***	2.758***	2.069**	2.048**	1.720**	1.208	0.000

Panel C: The equal variance test proposed in the current paper: post-crisis 2009/4/1–2012/12/31

	$\hat{\sigma}^2_{TW\_SKO}$	$\hat{\sigma}^2_{TW\_SIG}$	$\hat{\sigma}^2_{TW\_MAL}$	$\hat{\sigma}^2_{TW\_HK}$	$\hat{\sigma}^2_{TW\_AUS}$	$\hat{\sigma}^2_{TW\_US}$	$\hat{\sigma}^2_{TW\_TAI}$	$\hat{\sigma}^2_{TW\_PHI}$
$\hat{\sigma}^2_{TW\_SKO}$	0.000	−1.395	−1.810	−2.201	−3.938*	−6.342**	−11.286***	−10.200***
$\hat{\sigma}^2_{TW\_SIG}$	1.395	0.000	−1.006	−1.775	−8.344**	−13.250***	−10.622***	−10.044***
$\hat{\sigma}^2_{TW\_MAL}$	1.810	1.006	0.000	−1.527	−6.406**	−8.127**	−9.299***	−8.749***
$\hat{\sigma}^2_{TW\_HK}$	2.201	1.775	1.527	0.000	−4.476*	−4.977*	−8.586**	−8.247**
$\hat{\sigma}^2_{TW\_AUS}$	3.938*	8.344**	6.406**	4.676*	0.000	−3.663	−7.635**	−8.059**
$\hat{\sigma}^2_{TW\_US}$	6.342**	13.520***	8.127**	4.977*	3.663	0.000	−4.595*	−7.201**
$\hat{\sigma}^2_{TW\_TAI}$	11.285***	10.622***	9.299***	8.586**	7.635**	4.595*	0.000	−6.615**
$\hat{\sigma}^2_{TW\_PHI}$	10.200***	10.044***	8.749**	8.247**	8.059**	7.201**	6.615**	0.000

Panel D: The variance test proposed by Lee et al. (2012): post-crisis 2009/4/1–2012/12/31

	$\hat{\sigma}^2_{TW\_SKO}$	$\hat{\sigma}^2_{TW\_SIG}$	$\hat{\sigma}^2_{TW\_MAL}$	$\hat{\sigma}^2_{TW\_HK}$	$\hat{\sigma}^2_{TW\_AUS}$	$\hat{\sigma}^2_{TW\_US}$	$\hat{\sigma}^2_{TW\_TAI}$	$\hat{\sigma}^2_{TW\_PHI}$
$\hat{\sigma}^2_{TW\_SKO}$	0.000	−1.352*	−1.537*	−1.945**	−3.145***	−4.221***	−4.064***	−5.311***
$\hat{\sigma}^2_{TW\_SIG}$	1.352*	0.000	−0.289	−0.781	−2.063**	−3.244***	−3.529***	−4.905***
$\hat{\sigma}^2_{TW\_MAL}$	1.537*	0.289	0.000	−0.484	−1.735**	−2.915***	−3.359***	−4.773***
$\hat{\sigma}^2_{TW\_HK}$	1.945**	0.781	0.484	0.000	−1.231	−2.420***	−3.078***	−4.553***
$\hat{\sigma}^2_{TW\_AUS}$	3.145***	2.063**	1.735**	1.231	0.000	−1.231	−2.340***	−3.973***
$\hat{\sigma}^2_{TW\_US}$	4.221***	3.244***	2.915***	2.420***	1.231	0.000	−1.501*	−3.299***
$\hat{\sigma}^2_{TW\_TAI}$	4.064***	3.529***	3.359***	3.078***	2.340***	1.507*	0.000	−1.824**
$\hat{\sigma}^2_{TW\_PHI}$	5.311***	4.905***	4.773***	4.553***	3.973***	3.299***	1.824**	0.000

Notes: Same as those in Table 3.

However, the equal variance test proposed in this paper indicates a similar level of these two cointegrating relationships as shown in Panel A in Table 5, which therefore tends to deny the geographical ties.

According to Panel D in Table 5, the variance test results of Lee et al. (2012) indicates that the levels of cointegration relationships between Taiwan and four Asian economies (i.e., South Korea, Singapore, Malaysia and Hong Kong) are markedly different for the post-crisis period (2009/4/1–2012/12/31). However, our equal variance test indicates a similar level of cointegration relationships as shown in Panel C in Table 5. For the post-crisis period (2009/4/1–2012/12/31), the results of all of the variance tests support that the relationship between Taiwan and the United States is closer than those between Taiwan and certain Asian economies (i.e., Thailand and the Philippines).

Therefore, similarly to the previous analysis of the Asian financial crisis we cannot reject the dominant role of the United States economy in the world. These results show that, after considering the cross-sectional dependence between squared cointegrating errors, the statistical evidence supporting a closer cointegration between the capital markets of adjacent regions is much weaker than suggested by Lee et al. (2012).

4. Conclusion

To improve the knowledge of the level of cointegration relationship, this paper introduces a new generalized test that is based on the test proposed by Lee et al. (2012). The absence of this ability to answer how the level of comovement is for the traditional cointegration analysis including Engle–Granger procedure and Johansen and Juselius approach shows the importance of the variance tests. The ability to assess the closeness of the cointegrating relationships makes the variance tests relevant to applying to various issues, such as constructing portfolios among international stock markets. By allowing cross-sectional dependence between the squared cointegrating equilibrium errors, the proposed variance test provides further insights into the power of a hypothesis test, and is more adequate for the real-world analysis of the cointegration relationship than that in Lee et al. (2012).

We extend the variance test of Lee et al. (2012) and re-examine empirical tests of the price linkages and the degree of comovement between Taiwan and other countries in the event of the 1997 Asian financial crisis. One important feature of the proposed set-up is that it is robust to a variety of possible squared cointegrating equilibrium errors, such as GARCH(p,q). Another advantage of the proposed variance test is that it does not require an estimation of the variance, σ_m^2, by employing the KVB approach.

The results of this study show that the proposed test is more robust because it is asymptotically invariant to serial correlation/heteroscadasticity nuisance parameters in σ_m^2. We also show that the consideration of plausible dependence between capital markets raises questions about the validity of inferences based on the test proposed by Lee et al. (2012), which may lead to different empirical results.

Even though the data source, sample period, and sample countries used in this study are the same as those used by Lee et al. (2012), this study presents different results. We cannot find enough evidence to support the conclusion in Lee et al. (2012)that adjacent regions with similar backgrounds in terms of their capital markets will reflect price patterns. Our results are robust when we focus on the 2007–2009 financial crisis.

In summary, the empirical results of this study find closer relationships between Taiwan and other markets (i.e., the Philippines, the United States, and Australia) during the 1997 Asian financial crisis. Combining the cointegration test with the proposed equal variance test, we conclude that the linkage among stock markets was strengthened after the Asian financial crisis. The leading role of the United States stock market in Taiwan is founded in this paper, and geographical ties cannot be rejected. The findings of this study favor that frequent business cooperation and trading activities may be crucial factors in international stock price patterns. Future studies an employ our methodology to examine the degree of economic integration or convergence between developed and developing economies, or to assess the performance of mutual funds relative to a reference index.

Appendix A. Proofs of the theorems

The following two Lemmas enable us to estimate σ_ν^2 and σ_ς^2 consistently by using the OLS residuals in Eq. (6). The third lemma shows that a vector-form central limit theorem can be used to the vector of squared cointegrating equilibrium errors, i.e., the central limit theorem of Davidson (2002, Theorem 1.2) can be extended to the vector-valued NED case by using the standard Cramér Wold device.

Lemma A1. Under Assumption 1, $\frac{1}{T}\sum_{t=1}^{T}\nu_t^2 \xrightarrow{a.s.} \sigma_\nu^2$ and $\frac{1}{T}\sum_{t=1}^{T}\varsigma_t^2 \xrightarrow{a.s.} \sigma_\varsigma^2$.

Lemma A2. Given Assumption 1, the OLS estimates of parameters in Eq. (6), $\hat{\theta}_x$ and $\hat{\theta}_y$, are consistent. Further, we have

$$\frac{1}{T}\sum_{t=1}^{T}\nu_t^2 = \frac{1}{T}\sum_{t=1}^{T}\hat{\nu}_t^2 + o_p(1),$$

$$\frac{1}{T}\sum_{t=1}^{T}\varsigma_t^2 = \frac{1}{T}\sum_{t=1}^{T}\hat{\varsigma}_t^2 + o_p(1).$$

Lemma A3. Let vector $\mathbf{z}_t = (\nu_t^2 - \sigma_\nu^2, \varsigma_t^2 - \sigma_\varsigma^2)$, and $\mathbf{V}_T = \mathrm{Var}\left(T^{-1/2}\sum_{t=1}^{T}\mathbf{z}_t\right)$. Then, under Assumptions 1 and 2, the distribution of $\mathbf{z}_t$ is asymptotically normal:

$$\mathbf{V}_T^{-1/2}T^{-1/2}\sum_{t=1}^{T}\mathbf{z}_t \xrightarrow{d} N(0,\mathbf{I}), \text{ or equivalently,}$$

$$\mathbf{V}_T^{-1/2}T^{1/2}\begin{pmatrix}\frac{1}{T}\sum_{t=1}^{T}\nu_t^2 - \sigma_\nu^2(12) \\ \frac{1}{T}\sum_{t=1}^{T}\varsigma_t^2 - \sigma_\varsigma^2(13)\end{pmatrix} \xrightarrow{d} N(0,\mathbf{I}). \tag{12}$$

Proof of Lemma A1. Given Assumption 1, both ν_t^2 and ς_t^2 are stationary and ergodic.[8] Hence, applying the law of large numbers,[9] we can obtain the results.

Proof of Lemma A2. The assumptions on the dependence of squared cointegrating equilibrium errors (mixingale or NED) do not affect the asymptotic properties of OLS estimates in Eq. (6) as long as cointegrating equilibrium errors are stationary and ergodic. Therefore, the proof of Lemma A2 here is identical to that of Lemma 1 in Lee et al. (2012, p. 345).

Proof of Lemma A3. Davidson (2002, Theorem 1.2) had derived the central limit theorem (CLT) for a sequence of NED scalars, say x_t, given the conditions: (a) x_t is L_2-NED of size $-\frac{1}{2}$ on a process $\{e_t\}_{t=-\infty}^{\infty}$, where e_s is an α-mixing of size $-r/(r-2)$ for $r>2$ or ϕ-mixing of size $-r/(2r-2)$ for $r \geq 2$; (b) $sup_t\|Ex_t - Ex_t\|_r < \infty$; (c) $T^{-1/2}\sum_{t=1}^{T}x_t$ has finite variance as $T \to \infty$. Define the scalar $U_t = \lambda'\mathbf{V}^{-1/2}\mathbf{z}_t$, where λ is a 2×1 vector satisfying $\lambda'\lambda = 1$ and $\mathbf{V}$ is a 2×2 finite positive-definite matrix, and $\mathbf{z}_t = (z_{1t}, z_{2t})$. To prove Lemma A3, we first verify that the conditions for the CLT on U_t is satisfied and it is natural to extend the results of scalar U_t's CLT to the vector $\mathbf{z}_t$ by the standard Cramér–Wold device theorem.

First, denote $\lambda'\mathbf{V}^{-1/2} = \left(\tilde{\lambda}_1, \tilde{\lambda}_2\right)$. Under Assumption 2, we have

$$\begin{aligned}\left\|U_t - E\left(U_t|\mathcal{F}_{t-m}^{t+m}\right)\right\|_2 &= \left\|\lambda'\mathbf{V}^{-1/2}\mathbf{z}_t - E\left(\lambda'\mathbf{V}^{-1/2}\mathbf{z}_t|\mathcal{F}_{t-m}^{t+m}\right)\right\|_2 \\ &= \left\|\sum_{i=1}^{2}\tilde{\lambda}_i z_{it} - E\left(\sum_{i=1}^{2}\tilde{\lambda}_i z_{it}|F_{t-m}^{t+m}\right)\right\|_2 \\ &\leq \tilde{\lambda}_1\left\|\left(\nu_t^2 - \sigma_\nu^2\right) - E\left(\left(\nu_t^2 - \sigma_\nu^2\right)\middle|\mathcal{F}_{t-m}^{t+m}\right)\right\|_2 + \tilde{\lambda}_2\left\|\left(\varsigma_t^2 - \sigma_\varsigma^2\right) - E\left(\left(\varsigma_t^2 - \sigma_\varsigma^2\right)\middle|\mathcal{F}_{t-m}^{t+m}\right)\right\|_2 \\ &\leq 2v(m) \to 0.\end{aligned} \tag{13}$$

[8] See (White, 2001, Proposition 3.36).
[9] See (White, 2001, Theorem 3.57), Assumptions 1.1 and 1.3 constitute sufficient conditions for the law of large numbers.

Therefore, U_t is L_2-NED of size $-\frac{1}{2}$ on the process $\{e_t\}_{t=-\infty}^{\infty}$. Next, due to assumption 1, we have $E(\nu_t^2 - \sigma_\nu^2)^2 < \infty$, and $E(\varsigma_t^2 - \sigma_\varsigma^2)^2 < \infty$. Hence,

$$\|U_t\|_2 = \left\|\widetilde{\lambda}_1 z_{1t} + \widetilde{\lambda}_2 z_{2t}\right\|_2 \le \widetilde{\lambda}_1 \|z_{1t}\|_2 + \widetilde{\lambda}_2 \|z_{2t}\|_2 < \infty. \tag{14}$$

In other words, the sequence $\{U_t\}$ is uniformly L_2-bounded. Third, we verify the existence of the global variance. From Assumption 2 and the definition of $\mathbf{V}$ and λ, we obtain

$$\mathrm{Var}\left(T^{-1/2}\sum_{t=1}^{T} U_t\right) = \mathrm{Var}\left(T^{-1/2}\sum_{t=1}^{T}\lambda'\mathbf{V}^{-1/2}\mathbf{z}_t\right) = \lambda'\mathbf{V}^{-1/2}\mathbf{V}_T\mathbf{V}^{-1/2}\lambda < \infty, \tag{15}$$

which implies that the variance of U_t exists. Denote this variance as $\mathrm{Var}\left(T^{-1/2}\sum_{t=1}^{T} U_t\right) = \sigma^2$. From Eqs. (13), (14) and (15), the conditions of the CLT (Davidson, 2002, Theorem 1.2) are satisfied and therefore we can apply the (functional) CLT to U_t. Thus, we have

$$\mathcal{U}_T(r) = T^{-1/2}\sum_{t=1}^{[Tr]} U_t = T^{-1/2}\sum_{t=1}^{[Tr]}\lambda'\mathbf{V}^{-1/2}\mathbf{z}_t \xrightarrow{d} \sigma^2 W(r). \tag{16}$$

Set $\mathbf{V} = \lim_{T\to\infty} \mathbf{V}_T$, then $\sigma^2 = \lambda'\mathbf{V}^{-1/2}\mathbf{V}_T\mathbf{V}^{-1/2}\lambda = 1$, and we can rewrite Eq. (16) as

$$T^{-1/2}\sum_{t=1}^{[Tr]}\lambda'\mathbf{V}_T^{-1/2}\mathbf{z}_t \xrightarrow{d} W(r). \tag{17}$$

By using the Cramér–Wold device theorem, we obtain

$$\mathbf{V}_T^{-1/2}T^{-1/2}\sum_{t=1}^{[Tr]}\mathbf{z}_t \xrightarrow{d} \mathbf{W}(r).$$

When $r = 1$, $\mathbf{V}_T^{-1/2}T^{-1/2}\sum_{t=1}^{T}\mathbf{z}_t = \mathbf{V}_T^{-1/2}T^{1/2}\begin{pmatrix}\frac{1}{T}\sum_{t=1}^{T}\nu_t^2 - \sigma_\nu^2 \\ \frac{1}{T}\sum_{t=1}^{T}\varsigma_t^2 - \sigma_\varsigma^2\end{pmatrix} \xrightarrow{d} \mathbf{W}(1) = N(0, \mathbf{I})$. This completes the proof of this Lemma.

Proof of Theorem 1. The results of Theorem 1 are obtained immediately from Lemma A1, Lemma A2, and Lemma A3.

Proof of Theorem 2. For any 2×1 constant vector κ, Theorem 1 shows that

$$\left(\kappa'\mathbf{V}_T\kappa\right)^{-1/2}T^{1/2}\kappa'\begin{pmatrix}\hat{\sigma}_\nu^2 - \sigma_\nu^2 \\ \hat{\sigma}_\varsigma^2 - \sigma_\varsigma^2\end{pmatrix} \xrightarrow{d} N(0,1). \tag{18}$$

Specially, set $\kappa = [1\ -1]'$. Then, we have

$$\begin{aligned}\kappa'\mathbf{V}_T\kappa &= \mathrm{T}^{-1}[1 -1]\cdot\begin{bmatrix}\mathrm{Var}\left(\sum_{t=1}^{T}\nu_t^2\right) & \mathrm{Cov}\left(\sum_{t=1}^{T}\nu_t^2, \sum_{t=1}^{T}\varsigma_t^2\right)(21) \\ \mathrm{Cov}\left(\sum_{t=1}^{T}\nu_t^2, \sum_{t=1}^{T}\varsigma_t^2\right) & \mathrm{Var}\left(\sum_{t=1}^{T}\varsigma_t^2\right)(22)\end{bmatrix}\cdot\begin{bmatrix}1(23) \\ -1(24)\end{bmatrix} \\ &= \mathrm{T}^{-1}\left[\mathrm{Var}\left(\sum_{t=1}^{T}\nu_t^2\right) - 2\mathrm{Cov}\left(\sum_{t=1}^{T}\nu_t^2, \sum_{t=1}^{T}\varsigma_t^2\right) + \mathrm{Var}\left(\sum_{t=1}^{T}\varsigma_t^2\right)\right] \\ &= \mathrm{T}^{-1}\mathrm{Var}\left(\sum_{t=1}^{T}\left(\nu_t^2 - \varsigma_t^2\right)\right) \\ &\equiv \sigma_m^2.\end{aligned} \tag{19}$$

By substituting Eq. (19) into Eq. (18), we obtain the results of Theorem 2. This completes the proof of Theorem 2.

Proof of Lemma 1. Define $d = \sigma_\nu^2 - \sigma_\varsigma^2$, then using the results of Lemma A3 and Theorem 1, the limiting behavior of $T^{-1/2}\mathcal{S}_T(r)$ is

$$T^{-1/2}\mathcal{S}_T(r) = T^{-1/2}\sum_{t=1}^{[Tr]}\left(\hat{\nu}_t^2 - \hat{\varsigma}_t^2 - d + d - \frac{1}{T}\sum_{s=1}^{T}\left(\hat{\nu}_s^2 - \hat{\varsigma}_s^2\right)\right)$$
$$= T^{-1/2}\sum_{t=1}^{[Tr]}\left(\hat{\nu}_t^2 - \hat{\varsigma}_t^2 - d\right) - \frac{1}{T}\sum_{t=1}^{[Tr]}T^{-1/2}\sum_{s=1}^{T}\left(d - \left(\hat{\nu}_s^2 - \hat{\varsigma}_s^2\right)\right) \xrightarrow{d} \sigma_m(W(r) - rW(1)).$$

Let $\widehat{\mathcal{M}}^2 = T^{-2}\sum_{[Tr]=1}^{T}\hat{\mathcal{S}}_T^2(r)$ by continuous mapping theorem we have

$$\widehat{\mathcal{M}}^2 \xrightarrow{d} \sigma_m^2\int_0^1[W(r) - rW(1)]^2 dr.$$

This completes the proof of this Lemma.

Appendix B. The variances of cointegrating errors and the degree of cointegration

The reason for using an equal variance test to assess the degree of cointegration by comparing the variance of equilibrium errors is illustrated in the following simulated data. Consider the data generating process (DGP):[10]

$$x_{1t} = y_t + \nu_t,$$
$$x_{2t} = y_t + \varsigma_t,$$

where $y_t = \sum_{s=1}^{t}\varepsilon_s$, in which $\varepsilon_t \sim i.i.d.N(0,1)$. We assume that $\nu_t \sim i.i.d.N(0,1)$ and $\varsigma_t \sim i.i.d.N(0,4)$. Plots of this DGP for a sample of size 100 are shown in Fig. 1. It is easily observed from the set-up that y_t is cointegrated with x_{1t} and with x_{2t}, but the cointegrating relationship between y_t and x_{1t} is closer than that between y_t and x_{2t} since $\sigma_\nu^2 < \sigma_\varsigma^2$.

Appendix C. Finite sample properties of the test of Lee et al. (2012) and the new variance test

The cointegrating equilibrium errors processes must be stationary and the aim of the proposed variance test is to determine the degree of cointegration by comparing the variances of the cointegrating equilibrium errors. Here for simplicity, we restrict our experiment to examine the finite sample performances of the new variance test and the test of Lee et al. (2012) by assuming that the cointegrating equilibrium errors are known in the simulation. In fact, the cointegrating equilibrium errors are unknown and must be estimated. However, the simulations in this setup are helpful and insightful to understand the performance of the variance tests. We assume that the data generating processes (DGP) of cointegrating equilibrium errors follow a stationary $AR(1)$ process with an exogeneous common factor:

$$\nu_t = af_t + b\nu_{t-1} + \varepsilon_{1t}, \tag{20}$$

$$\varsigma_t = cf_t + d\varsigma_{t-1} + \varepsilon_{2t}, \tag{21}$$

and we employ the variance tests to the hypothesis,

$$H_0 : \sigma_\nu^2 = \sigma_\varsigma^2, \text{ against } H_1 : \sigma_\nu^2 < \sigma_\varsigma^2,$$

[10] Here the DGP contains a common independent variable.

where f_t, ε_{1t} and ε_{2t} are mutually independent and assumed to be *i.i.d.N*(0,1). f_t is the common effect of the process ν_t and ς_t. To compare the finite sample properties of the variance tests, we focus on two scenarios: the cross-sectionally independent and the cross-sectionally dependent case. In this setup, a and c in Eqs. (20) and (21) are used to control the cross-sectionally dependence, and if $a = c = 0$, the DGP degenerates to the cross-sectionally independent, but serially correlated case. b and d in Eqs. (20) and (21) are used to describe the serial correlation, where b, $d \in [0,1)$. It is also worthy to note that, for other sample size, the finite sample properties of the tests are generally similar although they are not reported here to save space. The finite sample properties for $T = 100$ are reported in Table 6.

Notes: The tests were one-sided with the nominal size set at 5%, and were conducted for sample size $T = 100$ using 1000 replications. "Lee" denotes the test of Lee et al. (2012), and "New" denotes the variance test proposed in this paper.

The simulations indicate that: (1) the size control of the variance test of Lee et al. (2012) depends strongly on the serial correlation, thus a strong serial correlation can induce a spurious rejection of the null, while the proposed variance test has relatively good size; (2) a violation of the cross-sectionally uncorrelated squared cointegrating errors can invalidate the variance test of Lee et al. (2012), especially in the cases where the squared cointegrating errors are strongly cross-sectionally dependent. However, the proposed variance test can be used to achieve satisfactory performance.

Appendix D. The standard errors of the critical values in Table 1

In Table 1, we report the simulated critical values of the proposed test. Here, we report the standard errors in each critical value to ensure that simulating 100,000 Monte-Carlo replications is statistically sufficient. Because the empirical results in the current paper depend mainly on the sample size of approximately $T = 250/1000$, we use these cases as examples. 1.0 Notes: The standard error of the critical value is based on 100 times' independent 100,000 Monte-Carlo replications which is used to calculate the critical value of the equal variance test in Table 1.

The standard errors of the critical values.

T	1.00%	2.50%	5.00%	10.00%	15.00%	50.00%	90.00%	95.00%	97.50%	99.00%
250	0.0065	0.0039	0.0027	0.0021	0.0017	0.0010	0.0022	0.0031	0.0042	0.0061
1000	0.0055	0.0036	0.0025	0.0017	0.0015	0.0009	0.0019	0.0028	0.0038	0.0058

Notes: The standard error of the critical value is based on 100 times' independent 100,000 Monte-Carlo replications which is used to calculate the critical value of the equal variance test in Table 1.

In probability theory and statistics, the coefficient of variation ($CV = \frac{\sigma}{\mu}$) is a normalized measure of the dispersion of a probability distribution or frequency distribution. The coefficient of variation is useful because the standard deviation of data must always be understood in the context of the mean of the data. In contrast, the actual value of the CV is independent of the unit in which the measurement has been taken, so it is a dimensionless number. In our case, the CV values were all considerably low. When we consider the 95% critical value where $T = 250$ for example, with the average of 100,000 Monte-Carlo replications executed 100 times being 5.326, the estimated CV of the empirical distribution of this simulated critical values equals $5.82 \times 10^{-4}\left(\widehat{CV} = \frac{\hat{\sigma}}{\hat{\mu}} = \frac{0.0031}{5.326}\right)$. Hence, the dispersion of the empirical distribution is low. We therefore conclude that simulating 100,000 Monte-Carlo replications is statistically sufficient.

Appendix E. Data description: the eight stock price indexes

To obtain more insight on the comovements among international stock markets, we present figures corresponding to the United States and the eight stock price indexes chosen by our empirical analysis: the Dow Jones Industrial Average Index for the United States (US), 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 index for Thailand (TAI), and the Taiwan Stock Exchange weighted price index for Taiwan (TW) over the period January 1, 1992 to

Table 6
Size and power of the variance tests.

	Cross-sectional independent case: size						Cross-sectional independent case: power					
(b,d)	(0.1,0.1)	(0.5,0.5)	(0.75,0.75)	(0.9,0.9)	(0.95,0.95)	(0.98,98)	(0.1,0.2)	(0.1,0.4)	(0.1,0.5)	(0.1,0.6)	(0.1,0.7)	(0.1,0.9)
Lee	0.000	0.004	0.018	0.079	0.137	0.229	0.001	0.007	0.037	0.109	0.329	0.955
New	0.049	0.056	0.051	0.064	0.059	0.096	0.076	0.152	0.246	0.361	0.525	0.708
	Cross-sectional dependent case: size											
(a,b)	(1,0.1)	(1,0.75)	(1,0.9)	(1,0.95)	(1,0.975)	(1,0.99)	(5,0.1)	(5,0.75)	(5,0.9)	(5,0.95)	(5,0.975)	(5,0.99)
(c,d)	(1,0.1)	(1,0.75)	(1,0.9)	(1,0.95)	(1,0.975)	(1,0.99)	(5,0.1)	(5,0.75)	(5,0.9)	(5,0.95)	(5,0.975)	(5,0.99)
Lee	0.000	0.012	0.048	0.101	0.167	0.210	0.000	0.000	0.001	0.003	0.008	0.010
New	0.047	0.048	0.056	0.084	0.098	0.118	0.055	0.051	0.048	0.078	0.085	0.096
	Cross-sectional dependent case: power											
(a,b)	(1,0.1)	(1,0.1)	(1,0.1)	(1,0.1)	(1,0.1)	(1,0.1)	(5,0.1)	(5,0.1)	(5,0.1)	(10,0.1)	(10,0.1)	(10,0.1)
(c,d)	(1,0.2)	(1,0.4)	(1,0.6)	(1,0.7)	(1,0.8)	(1,0.98)	(5,0.3)	(5,0.5)	(5,0.7)	(10,0.3)	(10,0.5)	(10,0.7)
Lee	0.000	0.002	0.088	0.343	0.951	0.995	0.000	0.000	0.261	0.000	0.000	0.259
New	0.061	0.161	0.398	0.581	0.681	0.701	0.237	0.561	0.766	0.306	0.646	0.738

Notes: The tests were one-sided with the nominal size set at 5%, and were conducted for sample size $T = 100$ using 1000 replications. “Lee” denotes the test of Lee et al. (2012), and ”New” denotes the variance test proposed in this paper.

December 31, 2012. The sample period was divided into the pre-crisis stage (1992/1/1–1997/7/1), crisis stage (1997/7/2–1998/12/31), and post-crisis stage (1999/1/1–2007/12/31) when we focused on the 1997 Asian financial crisis. Meanwhile, when focusing on the 2007–2009 crisis, we divided the sample period into: a crisis period (2007/7/1–2009/3/31), which began with the Quant Meltdown in the summer of 2007 and ended with the trough of the stock market in March 2009, and a post-crisis period (2009/4/1–2012/12/31). A graphical depiction of these stock indexes is given in this appendix for the full sample period and the sub-sample periods. See Fig. 2(a)–(g).

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