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Dynamic linkages between the Thai and international stock markets

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This paper investigates the existence of cointegration and causality between the stock market price indices of Thailand and its major trading partners using monthly data (1987-2005). The Engle-Granger two-step procedure and Gregory and Hansen (1996) test (allowing for one structural break) provide no evidence of a long-run relationship between the stock prices of Thailand and these countries. We argue that there exist potential long-run benefits from diversifying the investment portfolios internationally to reduce the associated systematic risks across countries. However, in the short run there are three unidirectional Granger causalities running from the stock returns of Hong Kong, the Philippines and the UK to that of Thailand. Furthermore, there are two unidirectional causalities running from the stock returns of Thailand to those of Indonesia and the US. Bidirectional Granger causality test results suggest that the stock returns of Thailand and three of its neighbouring countries are interrelated.

Keywords

Dynamic, linkages, between, Thai, international, stock, markets

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DYNAMIC LINKAGES BETWEEN THE THAI AND INTERNATIONAL STOCK MARKETS

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Abstract

This paper investigates the existence of cointegration and causality between the stock market price indices of Thailand and its major trading partners using monthly data (1987-2005). The Engle-Granger two-step procedure and Gregory and Hansen (1996) test (allowing for one structural break) provide no evidence of a long-run relationship between the stock prices of Thailand and these countries. We argue that there exist potential long-run benefits from diversifying the investment portfolios internationally to reduce the associated systematic risks across countries. However, in the short run there are three unidirectional Granger causalities running from the stock returns of Hong Kong, the Philippines and the UK to that of Thailand. Furthermore, there are two unidirectional causalities running from the stock returns of Thailand to those of Indonesia and the US. Bidirectional Granger causality test results suggest that the stock returns of Thailand and three of its neighbouring countries are interrelated.

Keywords: Stock markets, Thailand, Structural break, Causality and Cointegration.

1. Introduction

There is a growing interest in the integration of international stock markets as can be seen by a number of empirical studies examining the various aspects of stock market linkages. These studies have mainly been motivated by the stock market crash in October 1987 and subsequent Asian financial crisis in 1997. For instance, Susmel and Engle (1994), Fraser and Power (1997), Kanas (1998b) and Fratzscher (2002) examine volatility spillovers across stock markets, Phylaktis and Ravazzolo (2002) report their test results using international capital asset pricing models. In addition to these studies, in the literature cointegration techniques have been widely used to investigate the long-run relationship among stock markets. These studies can be classified into three groups.

First, some of previous studies have mainly focused on developed markets in the US, Canada, Europe and Japan. For example, Kasa (1992), Richards (1995), Choudhry (1996), Kanas (1998a), Hamori and Imamura (2000) and Ahlgren and Antell (2002) found some evidences that there are interdependent linkages among the stock markets of developed countries. Second, other studies in the literature have examined the stock price linkages among only emerging stock markets without capturing the important influence of stock markets in developed countries. They found only weak evidence that there is a relationship among the Asian stock markets. For more details see Chaudhuri (1997), Sharma and Wongbangpo (2002), Worthington *et al.* (2003) and Yang *et al.* (2003).

Finally, the last group of studies did examine the interdependencies between developed and emerging markets but they did not incorporate the effect of possible structural changes in the long-run relationships such as the 1987 great crash and the Asian financial crisis in 1997. Due to earlier inconclusive results, there is no consensus among previous studies as to whether international stock markets are interdependent. For instance, while Masih and Masih (1999) and Syriopoulos (2004) found some pairwise long-run relationships between stock markets in developed countries and the stock markets of emerging countries, the other studies such as Chang (2001), Ng (2002) and Climent and Meneu (2003) did not find any empirical evidence suggesting that stock market dependence among such countries. All these studies have indeed deepened our understanding of interplay among international stock market linkages. However, allowing for a possible break in cointegration vectors in this paper we examine specifically the interplay between the stock markets in Thailand and 11 countries, which include both developed and emerging markets.

Therefore, compared to previous studies, this paper is different from two aspects. First, none of previous studies examined a possibility that the pairwise long-run relationship between the stock prices of two countries may have been subject to a structural break. In addition to the Engle-Granger two-step procedure, in this paper we employ the Gregory and Hansen (GH, 1996) cointegration test which allows for a structural break in the cointegrating vector. GH (1996) argued that structural breaks have important implications for cointegration analysis because these breaks can decrease the power of cointegration tests and lead to the under rejection of the null hypothesis of no cointegration.

Second, as discussed earlier most previous studies have focused on developed markets and few studies have examined both emerging markets and developed markets, whereas this study examines whether the Thai stock market is linked with the stock markets of its major trading partners. None of the existing studies focuses specifically on the Thai stock market, although some studies include Thailand in their sample countries (for example see Masih and Masih (1999), Chang (2001), Ng (2002), Sharma and Wongbangpo (2002), Climent and Meneu (2003), Worthington *et al.* (2003) and Phylaktis and Ravazzolo (2005)).

It is important to recognize that, the 1997 Asian financial crisis first began with the floating of the Thai baht in July 1997 and soon after spread rapidly to the Philippines, Malaysia, Indonesia and Korea. Following this crisis relatively small depreciations also engulfed Singapore and Japan (Barro, 2001). Therefore, Thailand can be considered as an important case among other emerging markets. In 2004, market turnover, number of listed domestic companies and value traded on the Stock Exchange of Thailand (SET) was 93.8 per cent, 465 companies and US$109949 million, respectively. The SET was classified as the 9th highest among emerging markets in terms of all above three measures, and the 19th, 20th and 24th on a global scale, respectively. In terms of market capitalization, the SET reached a record high US$115400 million which ranked 12th highest among all emerging markets and 31st in the world (Standard and Poor's, 2005).

The main objective of this paper is to investigate the long-run and short-run relationships between the Thai stock market and these of its major trading partners namely Australia, Hong Kong, Indonesia, Japan, Korea, Malaysia, the Philippines, Singapore, Taiwan, the UK and the US. We chose these countries because the share of the Thai exports and imports to/from these eleven countries were relatively high. In 2005 Japan and the US were Thailand's two biggest trading partners. Malaysia, Singapore, Indonesia and the Philippines are all members of the Association of Southeast Asian Nations (ASEAN) which aimed at removing trade barriers among its member countries. Hong Kong, Taiwan and Australia are also all among Thailand's top ten trading partners followed with Korea and the UK lying just outside the top ten.

The remainder of the paper is structured as follows. Section 2 discusses briefly the empirical methodology adopted in the paper. Section 3 describes the summary statistics of the data. Section 4 presents the empirical results of cointegration and causality tests. Finally, Section 5 provides some concluding remarks.

2. Empirical Methodology

We initially perform the Augmented Dickey-Fuller (ADF) unit root test to examine the time series properties of the data without allowing for any structural breaks. The ADF test is conducted using the following equation:

$$\Delta y_t = \mu + \beta t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + \varepsilon_t \tag{1}$$

where y_t denotes the time series being tested, $y_t = \ln P_t^i$ which $\ln P_t^i$ is the natural logarithm of the stock market price index in country i, Δ is the first different operator, t is a time trend term, k denotes the optimal lag length and ε_t is a white noise disturbance term. In this paper, the lowest value of the Akaike Information Criterion (AIC) has been used as a guide to determine the optimal lag length in the ADF regression. These lags augment the ADF regression to ensure that the error term is white noise and free of serial correlation. In addition, the Phillips-Perron (PP) test has been used as an alternative nonparametric model to control for serial correlation. By using the PP test, one can ensure that the higher order serial correlations in the ADF equation have been handled properly.

In other words, the ADF test corrects for higher order autocorrelation by including lagged differenced terms on the right-hand side of the ADF equation, whereas the PP test corrects the ADF t-statistic by removing the serial correlation in it. This nonparametric t-test uses the Newey-West heteroscedasticity autocorrelation consistent estimate and is robust to heteroscedasticity and autocorrelation of unknown form.

An important shortcoming associated with the ADF and PP tests is that they do not allow for the effect of structural breaks. Perron (1989) argued that if a structural break in a series is ignored, unit root tests can be erroneous in rejecting null hypothesis. Zivot and Andrews (ZA, 1992) have developed methods to endogenously search for a structural break in the data. We employ their model C which allows for one structural break in both the intercept and slope coefficients in the following equation:

$$\Delta y_t = \mu + \beta t + \theta DU_t + \gamma DT_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + \varepsilon_t \qquad (2)$$

where $DU_t = 1$ if $t > TB$, otherwise zero; TB denotes the time of break, $DT_t = t - TB$ if $t > TB$, otherwise zero. The "trimming region", in which we have searched for TB covers the $0.15T$-$0.85T$ period, where T is the sample size. Following Chaudhuri and Wu (2003) and Narayan and Smyth (2005), we have selected the break point (TB) based on the minimum value of the t statistic for α. In this study, k_{max} is set equal to 12.

After determining the order of integration of each variable, we need to test for the existence of any long-run relationship between the stock prices of Thailand and its major trading partners. We employ the Engle-Granger two-step procedure by first obtaining the resulting residuals of the following equation and then conducting a unit root test on them:

$$y_t = \mu_0 + \beta t + \varphi x_t + \varepsilon_t \qquad (3)$$

where y_t and x_t are the natural log of the stock price indices of Thailand and one of its major trading partners, respectively.

According to Engle and Granger (1987), if both y_t and x_t are I(1) and $\hat{\varepsilon}_t$ is I(0), then there exists a long-run relationship between these two variables. The resulting error-correction model (*ECM*) from such a model then can be written as:

$$\Delta y_t = \phi + \sum_{i=0}^{k1} \lambda_i \Delta x_{t-i} + \sum_{i=1}^{k2} \delta_i \Delta y_{t-i} + \eta ECM_{t-1} + v_t \qquad (4)$$

where $\lambda_i s$ are the estimated short-term coefficients, $\delta_i s$ denote the estimated coefficients of the lagged dependent variables added to ensure v_t or the disturbance term is white noise, η is the feedback effect capturing the speed of adjustment whereby short-term dynamics converge to the long-term equilibrium path indicated in equation (3) and ECM_t or $\hat{\varepsilon}_t$ is obtained from equation (3) by the ordinary least squares (OLS) method.

The general-to-specific methodology can then be used to omit insignificant variables in equation (4) on the basis of a battery of maximum likelihood tests. In this method, joint zero restrictions are imposed on explanatory variables in the unrestricted (general) model to obtain a parsimonious model. The null hypothesis of no cointegration is rejected if $\eta < 0$ and statistically significant.

One may argue that the lack of evidence of cointegration in previous studies in the literature could be attributed to the ignorance of the structural break in cointegrating vector. To address this issue, we have also used the GH test. GH (1996) postulate three alternative models similar to those proposed by ZA (1992) to capture the changes in parameters of the cointegrating vector. First, the level shift model (*C*) which assumes a change only in the intercept as shown below:

$$y_t = \mu_0 + \theta DU_t + \mu_1 x_t + \varepsilon_t \qquad (5)$$

The second model, a level shift and change in trend (*C/T*) which takes the following form:

$$y_t = \mu_0 + \theta DU_t + \beta t + \mu_1 x_t + \varepsilon_t \tag{6}$$

The third model, which allows for changes in both the intercept and slope of the cointegration vector (*C/S*), is presented as:

$$y_t = \mu_0 + \theta DU_t + \beta t + \mu_1 x_t + \mu_2 x_t DU_t + \varepsilon_t \tag{7}$$

where DU_t is defined as previously in equation (2). Intuitively within the range of 0.15*T*-0.85*T*, this techniques searches for a particular *TB*, which minimizes the value of the ADF* statistic for $\hat{\varepsilon}_t$. GH (1996) test the null hypothesis of no cointegration against the alternative hypothesis of cointegration with a single structural break at time *TB*, which is determined endogenously.

Finally, we conduct the Granger causality test based on the error correction model specified in equation (4). A variable such as Δx_t (the stock returns) Granger causes Δy_t if its past values can explain Δy_t, but past values of Δy_t do not explain Δx_t (Granger, 1969). If the two variables are not cointegrated and η in equation (4) is not negative and significant, the following bivariate vector autoregression (VAR) equations will then be used for causality test:

$$\Delta y_t = \phi + \lambda_0 \Delta x_t + \sum_{i=1}^{k1} \lambda_i \Delta x_{t-i} + \sum_{i=1}^{k2} \delta_i \Delta y_{t-i} + v_t \tag{8}$$

$$\Delta x_t = \phi' + \lambda_0' \Delta y_t + \sum_{i=1}^{k'1} \lambda_i' \Delta y_{t-i} + \sum_{i=1}^{k'2} \delta_i' \Delta x_{t-i} + v_t' \tag{9}$$

On the other hand, if y_t and x_t are cointegrated, the following error correction models are adopted:

$$\Delta y_t = \phi + \lambda_0 \Delta x_t + \sum_{i=1}^{k1} \lambda_i \Delta x_{t-i} + \sum_{i=1}^{k2} \delta_i \Delta y_{t-i} + \eta ECM_{t-1} + v_t \tag{10}$$

$$\Delta x_t = \phi' + \lambda_0' \Delta y_t + \sum_{i=1}^{k'1} \lambda_i' \Delta y_{t-i} + \sum_{i=1}^{k'2} \delta_i' \Delta x_{t-i} + \eta' ECM_{t-1} + v_t' \tag{11}$$

The Granger causality test can be conducted under two assumptions. First, if y_t and x_t are not cointegrated, then we can use equations (8) and (9) to test the following two null hypotheses: If in equation (8) $H_o : \lambda_1 = \lambda_2 = ... = \lambda_{k1}$ is rejected, then $\Delta x_t = \ln P_t^j - \ln P_{t-1}^j$ or the stock price return in country *j* Granger causes $\Delta y_t = \ln P_t^i - \ln P_{t-1}^i$ or the stock price return in country *i*. This can be written as $\Delta x_t \longrightarrow \Delta y_t$. Similarly if in equation (9) $H_o' : \lambda_1' = \lambda_2' = ... = \lambda_{k1}'$ is rejected, then we can conclude that Δy_t causes Δx_t or $\Delta y_t \longrightarrow \Delta x_t$. If both null hypotheses are simultaneously rejected, there would be bidirectional causality between the two variables, *i.e.* $\Delta y_t \longleftrightarrow \Delta x_t$. Second, if y_t and x_t are in fact cointegrated, then we need to use equations (10) and (11) to test the same two hypotheses. The inclusion of *ECM* in these two equations ensures that the long-term rum properties of the data are not lost when dealing with the first difference form. If in equation (10) $H_o : \lambda_1 = \lambda_2 = ... = \lambda_{k1}$ is rejected, then $\Delta x_t \longrightarrow \Delta y_t$ (Δx_t Granger causes Δy_t) or $\Delta x_t \longrightarrow \Delta y_t$. In the same way if in equation (11) $H_o' : \lambda_1' = \lambda_2' = ... = \lambda_{k1}'$ is rejected, then one can conclude that $\Delta y_t \longrightarrow \Delta x_t$. If both H_o and H_o' are jointly rejected, the causality between the two variable is bidirectional or $\Delta y_t \longleftrightarrow \Delta x_t$.

3. Data and Basic Statistics

The data included in this paper include the stock prices of the following 12 countries: Thailand (TH), Australia (AU), Hong Kong (HK), Indonesia (IN), Japan (JA), Korea (KO), Malaysia (MA), the Philippines (PH), Singapore (SG), Taiwan (TA), the UK and the US. Monthly data span from December 1987 to December 2005 with a base value of 100 in December 1987. All stock indices were obtained from Morgan Stanley Capital International.

According to the descriptive statistics of the data (not presented here due to the lack of space), the highest mean return is 0.008 per cent in Hong Kong and the US while the lowest is 0.000 per cent in Japan. The standard deviations range from 0.041 per cent in the US (the least volatile) to 0.145 per cent in Indonesia (the most volatile). The standard deviations of stock returns are lowest in developed economies (*i.e.* the US, the UK, Australia, Japan and Singapore, respectively), and the most volatile in Indonesia, Thailand and Taiwan, respectively. All monthly stock returns, $\ln(P_t / P_{t-1})$, have excess kurtosis which means that they have a thicker tail and a higher peak than a normal distribution. The calculated Jarque-Bera statistics and corresponding p-values are used to test for the normality assumption. Based on the Jarque-Bera statistics and p-values, this assumption is rejected at any conventional level of significance for all stock returns, with the only 3 exceptions being the monthly stock returns in Australia, Japan and the UK.

4. Empirical Results

As mentioned earlier, we first use the ADF and PP tests to determine the order of integration of the 12 stock prices studied in this paper. The lowest value of the AIC has been used to determine the optimal lag length in the estimation procedure. Based on the results of the conventional unit root tests (available from the authors upon request), the ADF and PP tests reject the random walk hypothesis for only the stock price index in Taiwan at the 5 and 1 per cent significance levels, respectively. However, for all other countries both unit root tests cannot reject the random walk hypothesis. We thus conclude that the stock price indices in 11 out of 12 countries are I(1).

In the second stage, we subject each variable to one structural break. For each series, we then carried out the ZA test (model C) and reported the results in Table 1.

TABLE 1. The Zivot and Andrews Test Results: Break in Both the Intercept and Trend

$$\Delta y_t = \mu + \beta t + \theta DU_t + \gamma DT_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + \varepsilon_t$$

Variable	TB	μ	β	θ	γ	α	k	Inference
$\ln P_t^{TH}$	1996:10	0.420 (3.788)***	0.001 (1.339)	-0.170 (-3.659)***	-0.000 (-0.071)	-0.078 (-3.574)	12	Random walk
$\ln P_t^{AU}$	2001:02	0.792 (3.667)***	0.001 (3.240)***	-0.062 (-2.955)***	0.002 (3.724)***	-0.167 (-3.651)	10	Random walk
$\ln P_t^{HK}$	1993:01	0.652 (4.090)***	0.002 (2.478)**	0.074 (2.320)**	-0.002 (-2.455)**	-0.144 (-4.128)	11	Random walk
$\ln P_t^{IN}$	1997:08	0.831 (5.765)***	0.000 (0.708)	-0.258 (-4.835)***	0.001 (1.206)	-0.137 (-5.695)***	8	Stationary
$\ln P_t^{JA}$	2002:06	0.623 (4.069)***	-0.000 (-2.227)**	-0.068 (-2.565)**	0.003 (2.990)***	-0.132 (-4.089)	9	Random walk
$\ln P_t^{KO}$	1997:09	1.004 (5.425)***	-0.000 (-0.530)	-0.160 (-3.906)***	0.003 (4.267)***	-0.200 (-5.444)**	9	Stationary
$\ln P_t^{MA}$	1997:07	0.883 (6.774)***	0.002 (5.095)***	-0.234 (-6.121)***	-0.001 (-2.492)**	-0.185 (-6.719)***	9	Stationary
$\ln P_t^{PH}$	1993:07	0.440 (3.426)***	0.001 (1.237)	0.073 (1.892)	-0.002 (-2.163)**	-0.090 (-3.468)	12	Random walk
$\ln P_t^{SG}$	1997:03	0.572 (3.781)***	0.001 (2.976)***	-0.075 (-3.081)***	-0.001 (-2.089)**	-0.119 (-3.714)	7	Random walk
$\ln P_t^{TA}$	1993:10	0.885 (4.019)***	-0.002 (-2.045)**	0.109 (2.844)***	0.001 (1.570)	-0.150 (-4.102)	9	Random walk
$\ln P_t^{UK}$	1996:08	0.361 (3.076)***	0.000 (2.016)**	0.032 (2.132)**	-0.001 (-2.148)**	-0.077 (-3.018)	2	Random walk
$\ln P_t^{US}$	1996:09	0.313 (3.407)***	0.001 (2.568)**	0.040 (2.657)***	-0.001 (-2.761)***	-0.066 (-3.338)	7	Random walk

Notes: a) ** and *** indicate that the corresponding null hypothesis is rejected at the 5 and 1 per cent significance levels, respectively. b) Critical values for t_α at the 5 and 1 per cent levels are -5.08 and -5.57, respectively (Zivot and Andrews, 1992).

As mentioned earlier, the ADF and PP test results reveal that most stock prices examined in this paper followed a random walk, whereas the results of the ZA test show that stock prices for three countries (*i.e.* Indonesia, Korea and Malaysia) are now stationary. Despite allowing for one endogenous structural break in the data, the data in the remaining nine countries still contain a unit root. The estimated coefficients μ and θ are statistically significant for all variables except for θ in the case of the Philippines' stock prices. Thus at least there has been one structural break in the intercept during the sample period for all stock prices. The estimated coefficients for β and γ are also statistically significant in 8 and 9 out of 12 countries, respectively, implying the stock price series exhibit an upward or downward trend and there exist at least one structural break in trend in these countries.

The reported *TB*s in the second column of Table 1 have been endogenously determined by the ZA test. It is not surprising to note that the endogenously-determined structural breaks in these stock prices mostly occurred in the Asian crisis period 1996-1997, see *TB*s for Indonesia, Korea, Malaysia, Singapore, Thailand, the UK and the US in Table 1.

Since majority of the stock price indices are non-stationary, we have conducted the Engle-Granger cointegration test. Table 2 shows the results of this test for all 12 countries. As can be seen from the results, the null hypothesis of no cointegration cannot be rejected for all pairwise cases. In order to make robust conclusion we have also conducted the GH test and the results are presented in Table 3. Similar to the Engle-Granger test results, we found that Thai stock price index is not cointegrated with the stock prices any other 11 countries in our sample. This means that there is no pairwise long-run relationship between the stock prices in Thailand and its trading partners. Moreover, according to Table 3, it is interesting to note that the structural break in the cointegrating vector for most countries occurred 1998 (the year after the 1997 Asian financial crisis). However, the cointegration test results remain robust even after capturing the structural breaks in cointegrating vectors.

TABLE 2. The Engle-Granger Two-Step Test Results

	t-statistics	
	ADF test on $\hat{\varepsilon}_t$ (equation 3)[a]	$\hat{\eta}$ coefficient (equation 4)
Thailand-Australia	-2.165(0)	-1.390
Thailand-Hong Kong	-2.412(12)	-0.771
Thailand-Indonesia	-2.965(0)	-1.270
Thailand-Japan	-2.098(0)	-1.520
Thailand-Korea	-2.117(0)	-2.190*
Thailand-Malaysia	-2.884(12)	0.109
Thailand-Philippines	-2.130(12)	-1.610
Thailand-Singapore	-1.297(2)	-0.885
Thailand-Taiwan	-2.406(12)	-1.470
Thailand-UK	-2.309(12)	-2.300*
Thailand-US	-2.468(12)	-3.050*

Notes: a) We do not reject the null (*i.e.* a unit root in $\hat{\varepsilon}_t$) at the 5 per cent level or better as the critical values at the 5 and 1 per cent are -3.43 and -4.00, respectively (MacKinnon, 1991). b) Figures in parentheses are the optimal lag length determined by the AIC.

In sum, the similar results emerged from applying both the Engle-Granger test and the GH test to the data, suggesting that the Thai stock market is not cointegrated pairwise with the stock markets of any of the following countries: Australia, Hong Kong, Indonesia, Japan, Korea, Malaysia, the Philippines, Singapore, Taiwan, the UK and the US. Our results are also consistent with the previous findings of no cointegration between the Thai stock market and some regional stock markets including those of South-East Asia (Ng, 2002) and Pacific Basin (Chang, 2001; Climent and Meneu, 2003).

Finally, in the absence of long-run relationship between the stock prices of Thailand and its major trading partners, we now use the Granger causality test to examine the pairwise short-run interactions between different stock markets. The results of Granger causality test are presented in Table 4. The Wald F-statistics are calculated to test the null hypotheses outlined in the previous section.

TABLE 3. The Gregory and Hansen Test Results

Model *C*: $y_t = \mu_0 + \theta DU_t + \mu_1 x_t + \varepsilon_t$

Model *C/T*: $y_t = \mu_0 + \theta DU_t + \beta t + \mu_1 x_t + \varepsilon_t$

Model *C/S*: $y_t = \mu_0 + \theta DU_t + \beta t + \mu_1 x_t + \mu_2 x_t DU_t + \varepsilon_t$

Model	*TB*	ADF*	*k*
Thailand-Australia			
C	1998:06	-3.842	12
C/T	1998:07	-3.609	10
C/S	1998:06	-3.862	12
Thailand-Hong Kong			
C	1998:06	-3.527	12
C/T	2002:10	-3.797	12
C/S	1998:06	-3.444	12
Thailand-Indonesia			
C	1991:12	-3.526	8
C/T	1997:08	-3.301	8
C/S	1991:11	-3.476	8
Thailand-Japan			
C	1998:06	-3.130	12
C/T	1998:06	-3.896	12
C/S	1998:06	-3.129	12
Thailand-Korea			
C	1998:07	-2.719	10
C/T	1998:07	-3.413	10
C/S	1998:07	-2.660	10
Thailand-Malaysia			
C	1998:02	-3.755	12
C/T	2003:06	-3.752	12
C/S	1994:10	-3.461	12
Thailand-Philippines			
C	1995:04	-2.795	12
C/T	2001:09	-3.443	12
C/S	1998:06	-2.834	12
Thailand-Singapore			
C	1996:04	-2.909	12
C/T	2002:10	-3.675	12
C/S	1996:04	-2.908	12
Thailand-Taiwan			
C	1998:06	-3.166	12
C/T	1998:06	-3.706	12
C/S	1998:06	-3.037	12
Thailand-UK			
C	1998:06	-3.247	12
C/T	1998:06	-3.947	12
C/S	1998:06	-3.177	12
Thailand-US			
C	1992:04	-3.298	12
C/T	1998:06	-4.120	12
C/S	1996:07	-3.349	12
Critical values		5 per cent	1 per cent
C		-4.61	-5.13
C/T		-4.99	-5.45
C/S		-4.95	-5.47

Note: Given the reported critical values (GH, 1996), the null is not rejected at the 5 per cent and 1 per cent levels of significance for any pair of countries.

According to the results presented in Table 4, in the short term there is a unidirectional Granger causality running from the stock returns of Hong Kong, the Philippines and the UK to that of Thailand; on the other hand, there is unidirectional Granger causality from Thailand's stock return to the stock returns of Indonesia and the US. In addition, we found bidirectional Granger causality between the market stock returns in Thailand and its three neighbouring countries *i.e.* Malaysia, Singapore and Taiwan. Therefore, the short-run movements of stock returns in these three countries can influence the performance of Thailand's stock market. It can also be concluded that any short-run variation of the stock returns in Thailand can impact on the market returns of its three neighboring countries and vice versa. Hence in order to avoid financial contagion and future crises similar to the one occurred in 1997, it is important for central bankers and individual investors to keep abreast of new developments in international stock markets particularly those for which we found the evidence of bidirectional and unidirectional causality.

5. Conclusions

In this study we have examined the long-run and short-run relationships between the stock prices of Thailand and its major trading partners (Australia, Hong Kong, Indonesia, Japan, Korea, Malaysia, the Philippines, Singapore, Taiwan, the UK and the US) using monthly data for the period December 1987-December 2005. In addition to the Engle-Granger two-step procedure, we have used the Gregory and Hansen (1996) test, which allows for a structural break in the cointegration vector. The major findings of this paper have been summarized below.

First, based on the cointegration results, we found no evidence of long-run relationship between the stock price indices of Thailand and its major trading partners. The policy implication of this finding for international investors is quite straightforward: in the long run there are potential gains (for example reduced systematic risks) which can be benefited by astute investors through portfolio diversification across different international markets. Second, in terms of short-run movements of international stock market returns, we found three-pairwise unidirectional Granger causalities, whereby the returns in Hong Kong, the Philippines and the UK can Granger cause Thailand. Based on these results, one may argue that the performance of stock market in Honk Kong, the Philippines and the UK can have direct bearing on the Thai stock market. On the other hand, there were also two unidirectional Granger causalities running from Thailand to Indonesia and the US. Thus any abnormal movement in Thailand's stock returns may lead to similar changes in Indonesia and the US. Third, we have found evidence of bidirectional Granger causality between the stock returns in Thailand and the three of its neighbouring countries *i.e.* Malaysia, Singapore and Taiwan. The reported causality test results can be useful in any assessment of the Asian stock markets. For examples the interplay between these three pairs of countries Thailand-Malaysia, Thailand-Singapore, Thailand-Taiwan can be useful by central bankers and international investors alike in evaluating the stock market performance.

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TABLE 4. The Granger Causality Test Results

$$\Delta y_t = \phi + \lambda_0 \Delta x_t + \sum_{i=1}^{k1} \lambda_i \Delta x_{t-i} + \sum_{i=1}^{k2} \delta_i \Delta y_{t-i} + \eta ECM_{t-1} + v_t$$

$$\Delta x_t = \phi' + \lambda'_0 \Delta y_t + \sum_{i=1}^{k'1} \lambda'_i \Delta y_{t-i} + \sum_{i=1}^{k'2} \delta'_i \Delta x_{t-i} + \eta' ECM_{t-1} + v'_t$$

Inference		Null hypothesis $H_o: \lambda_1 = \lambda_2 = \ldots = \lambda_{k1}$ or $H'_o: \lambda'_1 = \lambda'_2 = \ldots = \lambda'_{k1}$ F-statistic	Probability
No causality	$\Delta \ln P_t^{AU} \nrightarrow \Delta \ln P_t^{TH}$	1.034	0.399
No causality	$\Delta \ln P_t^{TH} \nrightarrow \Delta \ln P_t^{AU}$	1.817	0.111
Unidirectional causality	$\Delta \ln P_t^{HK} \rightarrow \Delta \ln P_t^{TH}$	7.013***	0.009
No causality	$\Delta \ln P_t^{TH} \nrightarrow \Delta \ln P_t^{HK}$	0.253	0.616
No causality	$\Delta \ln P_t^{IN} \nrightarrow \Delta \ln P_t^{TH}$	1.322	0.256
Unidirectional causality	$\Delta \ln P_t^{TH} \rightarrow \Delta \ln P_t^{IN}$	4.290***	0.001
No causality	$\Delta \ln P_t^{JA} \nrightarrow \Delta \ln P_t^{TH}$	0.144	0.704
No causality	$\Delta \ln P_t^{TH} \nrightarrow \Delta \ln P_t^{JA}$	1.720	0.191
No causality	$\Delta \ln P_t^{KO} \nrightarrow \Delta \ln P_t^{TH}$	0.358	0.550
No causality	$\Delta \ln P_t^{TH} \nrightarrow \Delta \ln P_t^{KO}$	0.404	0.526
Bidirectional causality $\Delta \ln P_t^{TH} \leftrightarrow \Delta \ln P_t^{MA}$	$\Delta \ln P_t^{MA} \rightarrow \Delta \ln P_t^{TH}$	1.870**	0.046
	$\Delta \ln P_t^{TH} \rightarrow \Delta \ln P_t^{MA}$	3.771***	0.000
Unidirectional causality	$\Delta \ln P_t^{PH} \rightarrow \Delta \ln P_t^{TH}$	1.936**	0.049
No causality	$\Delta \ln P_t^{TH} \nrightarrow \Delta \ln P_t^{PH}$	1.628	0.110
Bidirectional causality $\Delta \ln P_t^{TH} \leftrightarrow \Delta \ln P_t^{SG}$	$\Delta \ln P_t^{SG} \rightarrow \Delta \ln P_t^{TH}$	2.322*	0.076
	$\Delta \ln P_t^{TH} \rightarrow \Delta \ln P_t^{SG}$	2.633*	0.051
Bidirectional causality $\Delta \ln P_t^{TH} \leftrightarrow \Delta \ln P_t^{TA}$	$\Delta \ln P_t^{TA} \rightarrow \Delta \ln P_t^{TH}$	2.690**	0.011
	$\Delta \ln P_t^{TH} \rightarrow \Delta \ln P_t^{TA}$	1.798*	0.090
Unidirectional causality	$\Delta \ln P_t^{UK} \rightarrow \Delta \ln P_t^{TH}$	3.358***	0.006
No causality	$\Delta \ln P_t^{TH} \nrightarrow \Delta \ln P_t^{UK}$	1.577	0.168
No causality	$\Delta \ln P_t^{US} \nrightarrow \Delta \ln P_t^{TH}$	1.422	0.190
Unidirectional causality	$\Delta \ln P_t^{TH} \rightarrow \Delta \ln P_t^{US}$	2.335**	0.020

Note: *, ** and *** indicate that the corresponding null hypothesis is rejected at the 10, 5 and 1 per cent significance levels, respectively.