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An Empirical Analysis of the Thai and Major 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 (Australia, Hong Kong, Indonesia, Japan, Korea, Malaysia, the Philippines, Singapore, Taiwan, the UK and the US), using monthly data spanning December 1987 to December 2005. Both the Engle-Granger two-step procedure (assuming no structural breaks) and the 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. Based on the empirical results obtained from these two residual-based cointegration tests, potential long-run benefits exist from diversifying the investment portfolios internationally to reduce the associated systematic risks across countries. However, in the short run, three unidirectional Granger causalities run from the stock returns of Hong Kong, the Philippines and the UK to those of Thailand, pair-wise. Furthermore, there are two unidirectional causalities running from the stock returns of Thailand to those of Indonesia and the US. We also found empirical evidence of bidirectional Granger causality, suggesting that the stock returns of Thailand and three of its neighbouring countries (Malaysia, Singapore and Taiwan) are interrelated. No previous study examines the possibility that the pair-wise long-run relationship between the stock prices of Thailand and those of both emerging and developed markets may have been subject to a structural break.

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1. Introduction

There are many reasons why stock markets of different countries may have significant co-movements. For example global capital movements and the presence of economic ties and regional policy coordination among countries can directly or indirectly interconnect their stock prices through time. According to Phylaktis and Ravazzolo (2005), unlike other crises, the Asian crisis engulfed a group of countries that were both financially and economically integrated prior to the crisis. However, Chan et al. (1997) argue that although common economic and geographic factors were considered as crucial factors, they were not necessarily major causes of national stock markets to follow the same stochastic trend. It is also argued that there is less evidence of stock market integration after major stock market crises and hence international diversification among stock markets can be undertaken more effectively due to the lack of long-run co-movements of international stock prices (Patev et al., 2006). In the context of the Malaysian stock market, for example, Ibrahim and Aziz (2003) provide some evidence that the Asian crisis appears to give rise to irregularity in the interactions between stock prices and macroeconomic variables.

A growing interest in the integration of international stock markets is evident in the number of empirical studies that examine the various aspects of stock market linkages. These studies were mainly 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; while Phylaktis and Ravazzolo (2002) report their test results using international capital asset pricing models.

In addition to these studies, cointegration techniques in the literature are widely used to investigate the long-run relationships between stock markets. These studies can be classified into three groups. First, some focus mainly on developed markets in the US, Canada, Europe and Japan (for example, Kasa, 1992; Richards, 1995; Choudhry, 1996; Kanas, 1998a; Hamori and Imamura, 2000; Ahlgren and Antell, 2002) and find some evidence that there are interdependent linkages among the stock markets of developed countries. Second, other studies in the literature examine the stock price linkages among only emerging stock markets, without capturing the important influence of stock markets in developed countries. They find only weak evidence of a relationship among the Asian stock markets (for more details see Chaudhuri, 1997; Sharma and Wongbangpo, 2002; Worthington et al., 2003; Yang et al., 2003).

The last group of studies examines the interdependencies between developed and emerging markets but they do 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 pair-wise long-run relationships between stock markets in developed countries and the stock markets of emerging countries, other studies (such as Chang, 2001; Ng, 2002; Climent and Meneu, 2003) do not find any empirical evidence suggesting that stock market dependence exists among such countries. These studies have deepened our understanding of the interplay among international stock market linkages; however, allowing for a possible break in cointegration vectors, this paper specifically examines the interplay between the stock markets in Thailand and 11 other countries, including both developed and emerging markets.

Compared to previous studies, this paper differs in two aspects. First, no previous study examines the possibility that the pair-wise 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, this paper employs the Gregory and Hansen (GH, 1996) cointegration test, which allows for a structural break in the cointegrating vector. Gregory and Hansen argue 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 focus on developed markets, and few examine both emerging and developed markets. In contrast, this study examines whether the Thai stock market is linked with the stock markets of its major trading partners. No existing study focuses specifically on the Thai stock market, although some include Thailand in their sample countries (for example, Masih and Masih, 1999; Chang, 2001; Ng, 2002; Sharma and Wongbangpo, 2002; Climent and Meneu, 2003; Worthington et al., 2003; Phylaktis and Ravazzolo, 2005).

The 1997 Asian financial crisis first began with the floating of the Thai baht in July 1997, and soon after, the crisis 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, on the Stock Exchange of Thailand (SET), market turnover was 93.8 per cent, there were 465 listed domestic companies, and the value traded was $US109,949 million. The SET was classified as the ninth highest among emerging markets in terms of these three measures, and the nineteenth, twentieth and twenty-fourth on a global scale. In terms of market capitalisation, the SET reached a record high $US115,400 million, which ranked twelfth-highest among all emerging markets and thirty-first in the world (Standard and Poor's, 2005).

This paper investigates the long-run and short-run relationships between the Thai stock market and those of its major trading partners: Australia, Hong Kong, Indonesia, Japan, Korea, Malaysia, the Philippines, Singapore, Taiwan, the UK and the US. We chose these eleven countries because of their relatively high share of Thai exports and imports. It should be noted that Japan and the US are Thailand's two biggest trading partners. Malaysia, Singapore, Indonesia and the Philippines are all members of the Association of Southeast Asian Nations (ASEAN), which aims to remove trade barriers among its member countries. Hong Kong, Taiwan and Australia are also among Thailand's top-ten trading partners, followed by Korea and the UK, which are just outside the top ten.

The remainder of the paper is structured as follows. Section 1 discusses briefly the empirical methodology adopted in the paper. Section 2 describes the summary statistics of the data. Section 3 presents the empirical results of cointegration and causality tests. Finally, Section 4 provides some concluding remarks.

2. Empirical methodology

We initially performed 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 this equation:

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

3

Where y_t denotes the time series being tested; $y_t = \ln(P_t^i)$, $\ln(P_t^i)$ 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) was 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 was used as an alternative nonparametric model to control for serial correlation. Using the PP test ensures that the higher-order serial correlations in the ADF equation were handled properly. That is, 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) argues 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) developed methods to search endogenously 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:

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

Where $DU_t = 1$ if $t > TB$, otherwise zero; TB denotes the time of break; and $DT_t = t - TB$ if $t > TB$, otherwise zero.

The 'trimming region', in which we 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 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 needed to test for the existence of any long-run relationship between the stock prices of Thailand and its major trading partners. We employed the Engle-Granger two-step procedure first by obtaining the resulting residuals of the following equation, and then conducting a unit root test on them:

$$y_t = \mu_0 + \beta t + \varphi x_t + \varepsilon_t \tag{3}$$

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 a long-run relationship between these two variables exists. The resulting error-correction model (*ECM*) from such a model can then be written as:

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

Where $\lambda_i s$ are the estimated short-term coefficients; $\delta_i s$ denotes 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 OLS method.

The general-to-specific methodology can then be used to omit insignificant variables in equation (4) based on 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 is statistically significant.

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 also used the GH (1996) 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:

$$y_t = \mu_0 + \theta DU_t + \mu_1 x_t + \varepsilon_t \tag{5}$$

The second model, a level shift and change in trend (*C/T*), takes this 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 0.15T-0.85T, this technique searches for a particular *TB*, which minimises the value of the ADF* statistic for $\hat{\varepsilon}_t$. The GH (1996) method tests 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 conducted 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 VAR equations will then be used for the 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, these 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) $H_o : \lambda_1 = \lambda_2 = ... = \lambda_{k1} = 0$ 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 $\Delta x_t \rightarrow \Delta y_t$. Similarly, if, in equation (9), $H_o' : \lambda_1' = \lambda_2' = ... = \lambda_{k1}' = 0$ is rejected, then we can conclude that Δy_t causes Δx_t or $\Delta y_t \rightarrow \Delta x_t$. If both null hypotheses are rejected simultaneously, there would be a bidirectional causality between the two variables, that is, $\Delta y_t \leftrightarrow \Delta x_t$. 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 run properties of the data are not lost when dealing with the first difference form. If in equation (10), $H_o : \lambda_1 = \lambda_2 = ... = \lambda_{k1} = 0$ is rejected, then $\Delta x_t \rightarrow \Delta y_t$ (Δx_t Granger causes Δy_t), or $\Delta x_t \rightarrow \Delta y_t$. In the same way, if in equation (11), $H_o' : \lambda_1' = \lambda_2' = ... = \lambda_{k1}' = 0$ is rejected, then one can conclude that $\Delta y_t \rightarrow \Delta x_t$. If both H_o and H_o' are rejected, the causality between the two variables is bidirectional, or $\Delta y_t \leftrightarrow \Delta x_t$.

3. Data and summary statistics

The data included in this paper include the stock prices of these 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 December 1987 to December 2005, with a base value of 100 in December 1987. All stock price indices were obtained from Morgan Stanley Capital International (MSCI) which is one of the widely used sources of financial data in the literature (Hamori and Imamura, 2000; Ahlgren and Antell, 2002; Ng, 2002; Climent and Meneu, 2003; Worthington et al., 2003) in terms of the degree of comparability and avoidance of dual listings. Since this paper is concerned with the comparative performance of the international stock markets, all price indices (*P*) are denominated in US dollars. The MSCI indices for different markets are computed using the same consistent formula which is value weighted. The rate of returns $ln(P_t/P_{t-1})$ calculated from the MSCI price indices which consists of both capital gain and income gain .

Table I presents the descriptive statistics of the data, including sample means, medians, maximums, minimums, standard deviations, skewness, kurtosis as well as the Jarque-Bera statistics and *p*-values. 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 are Indonesia (the most volatile). The standard deviations of stock returns are lowest in developed economies (that is, the US, the UK, Australia and Japan), and the most volatile in Indonesia, Thailand, Taiwan and Korea. 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 three exceptions being the monthly stock returns in Australia, Japan and the UK.

4. Empirical results

As mentioned earlier, we first used the ADF and PP tests to determine the order of integration of the 12 stock prices studied. The lowest value of the AIC was used to determine the optimal lag length in the estimation procedure. Based on the results of the unit root tests presented in Table II, the ADF and PP tests reject the random walk hypothesis for only the stock price index in Taiwan at the five and one 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 the 12 countries are I(1).

Table I. Descriptions of the data (stock return) employed, December 1987-December 2005

Variable	Mean	Median	Maximum	Minimum	Standard deviation	Skewness	Kurtosis	Jarque-Bera	*p*-value
$\Delta \ln P_t^{TH}$	0.003	0.007	0.359	-0.416	0.119	-0.394	4.802	34.804	0.000
$\Delta \ln P_t^{AU}$	0.006	0.005	0.157	-0.166	0.053	-0.244	3.464	4.091	0.129
$\Delta \ln P_t^{HK}$	0.008	0.007	0.284	-0.344	0.077	-0.203	5.290	48.907	0.000
$\Delta \ln P_t^{IN}$	0.005	0.009	0.662	-0.525	0.145	0.415	7.320	174.181	0.000
$\Delta \ln P_t^{JA}$	0.000	-0.002	0.217	-0.216	0.066	0.077	3.437	1.944	0.378
$\Delta \ln P_t^{KO}$	0.005	-0.001	0.534	-0.375	0.111	0.306	5.914	79.815	0.000
$\Delta \ln P_t^{MA}$	0.004	0.005	0.405	-0.361	0.091	-0.200	6.731	126.730	0.000
$\Delta \ln P_t^{PH}$	0.002	0.005	0.360	-0.347	0.095	-0.021	4.744	27.405	0.000
$\Delta \ln P_t^{SG}$	0.006	0.009	0.228	-0.231	0.071	-0.502	5.365	59.702	0.000
$\Delta \ln P_t^{TA}$	0.004	0.002	0.381	-0.410	0.113	-0.034	4.179	12.556	0.002
$\Delta \ln P_t^{UK}$	0.006	0.004	0.138	-0.111	0.045	0.083	3.137	0.420	0.810
$\Delta \ln P_t^{US}$	0.008	0.011	0.106	-0.151	0.041	-0.556	3.871	18.022	0.000

Source: Morgan Stanley Capital International, http://www.msci.com/equity/index2.html.

Table II. Unit root test results

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

Variable	ADF test		PP test	
	Constant and trend	Optimal lag	Constant and trend	Bandwidth
$\ln P_t^{TH}$	-2.372	12	-2.046	5
$\Delta \ln P_t^{TH}$	-4.656***	6	-14.169***	7
$\ln P_t^{AU}$	-2.573	0	-2.478	7
$\Delta \ln P_t^{AU}$	-9.002***	4	-16.265***	12
$\ln P_t^{HK}$	-2.086	0	-2.050	8
$\Delta \ln P_t^{HK}$	-14.003***	0	-14.001***	11
$\ln P_t^{IN}$	-3.350	8	-2.595	5
$\Delta \ln P_t^{IN}$	-10.271***	1	-12.274***	3
$\ln P_t^{JA}$	-2.188	0	-2.387	3
$\Delta \ln P_t^{JA}$	-14.151***	0	-14.151***	1
$\ln P_t^{KO}$	-1.668	0	-1.744	1
$\Delta \ln P_t^{KO}$	-14.103***	0	-14.103***	4
$\ln P_t^{MA}$	-3.053	9	-2.332	4
$\Delta \ln P_t^{MA}$	-3.862**	10	-12.440***	0
$\ln P_t^{PH}$	-2.099	1	-2.006	2
$\Delta \ln P_t^{PH}$	-11.696***	0	-11.700***	3
$\ln P_t^{SG}$	-2.537	0	-2.552	1
$\Delta \ln P_t^{SG}$	-14.393***	0	-14.393***	1
$\ln P_t^{TA}$	-3.759**	1	-4.068***	5
$\Delta \ln P_t^{TA}$	-13.130***	0	-13.145***	2
$\ln P_t^{UK}$	-1.551	2	-1.805	6
$\Delta \ln P_t^{UK}$	-13.546***	1	-15.718***	9
$\ln P_t^{US}$	-1.178	0	-1.146	3
$\Delta \ln P_t^{US}$	-15.805***	0	-15.794***	3

Notes: a) ** and *** indicate that the corresponding null hypothesis is rejected at the 5 and 1 per cent significance levels, respectively. b) Critical values at the 5 and 1 per cent are -3.43 and -4.00, respectively (MacKinnon, 1991).

Table III. 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).

Figure 1. Plot of the international stock price indices and market returns
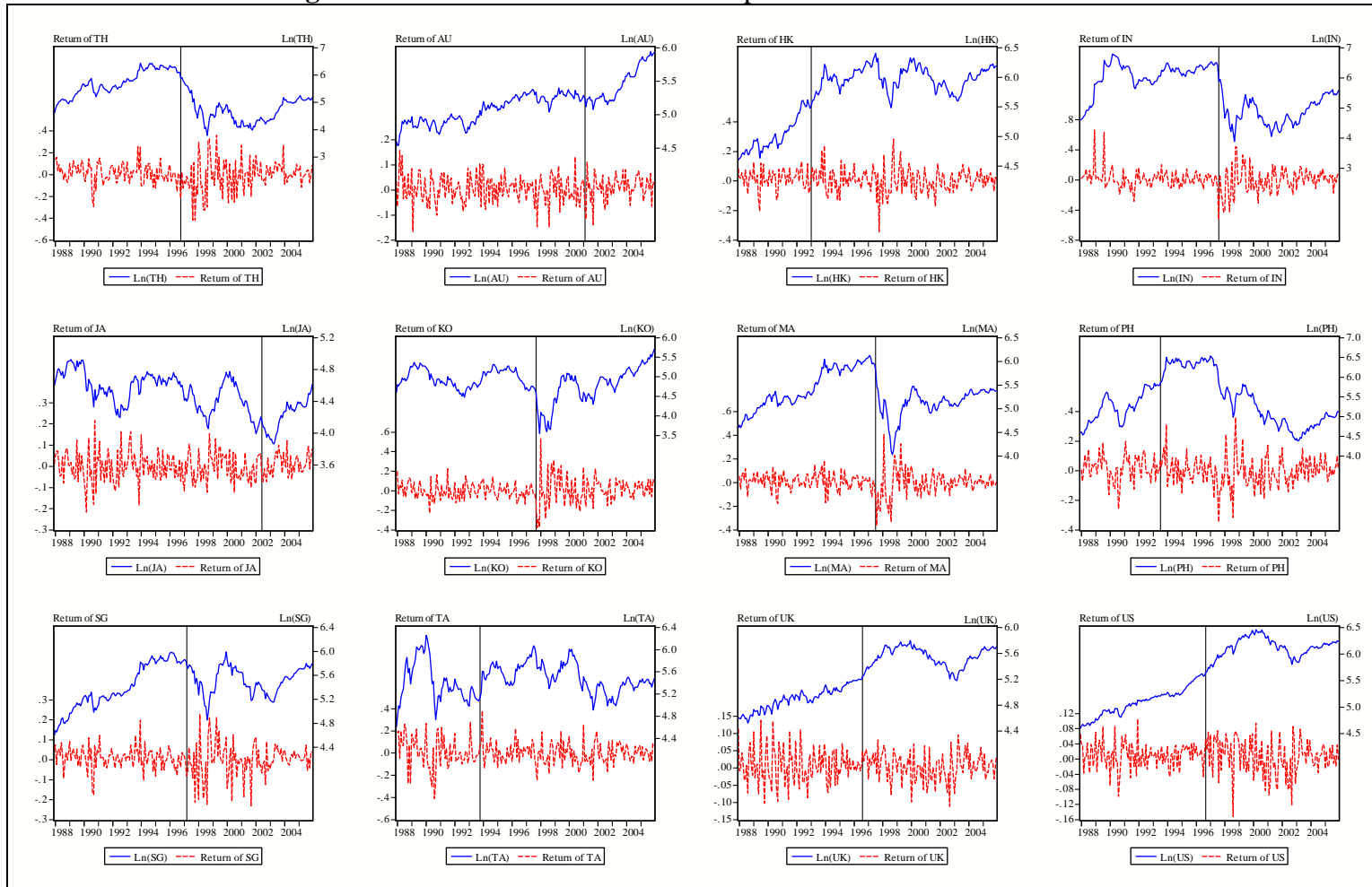
Source: Morgan Stanley Capital International, http://www.msci.com/equity/index2.html.
Note: The vertical line shows the time of the structural break obtained by the ZA (1992) method.

In the second stage, we subjected each variable to one structural break. For each series, we then carried out the ZA test (model C) and report the results in Table III, below. 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 (that is, 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. There was at least 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 eight and nine out of 12 countries, respectively, implying the stock price series exhibits an upward or downward trend and at least one structural break in trend in these countries exists.

The reported TBs in the second column of Table III were endogenously determined by the ZA test. In addition, Figure 1 shows $\ln(P_t)$ and $\ln(P_t / P_{t-1})$ (monthly return) for each of the 12 countries, as well as their corresponding structural breaks obtained by the ZA test. It is not surprising that the endogenously-determined structural breaks in these stock prices occurred mostly in the Asian crisis period 1996–1997 (see TBs for Thailand, Indonesia, Korea, Malaysia, Singapore, the UK and the US in Table III).

Because the majority of the stock price indices are non-stationary, we conducted the Engle-Granger cointegration test. Table IV shows the results of this test for all 12 countries. The results show that the null hypothesis of no cointegration cannot be rejected for all pair-wise cases. In order to make robust conclusions, we also conducted the GH test, and the results are presented in Table V. Similar to the Engle-Granger test results, we found that the Thai stock price index is not cointegrated with the stock prices of any other of the 11 countries in our sample. This means that there is no pair-wise long-run relationship between the stock prices in Thailand and its trading partners. Importantly, according to Table V, the structural break in the cointegrating vector for most countries occurred in 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.

In sum, similar results emerged from applying both the Engle-Granger test and the GH (1996) test to the data, suggesting that the Thai stock market is not cointegrated pair-wise with the stock markets of any of these 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 the Pacific Basin (Chang, 2001; Climent and Meneu, 2003).

Finally, in the absence of long-run relationships between the stock prices of Thailand and its major trading partners, we then used the Granger causality test to examine the pair-wise short-run interactions between different stock markets. Table VI presents the results of the Granger causality tests. The Wald F-statistics are calculated to test the null hypotheses outlined in the previous section. According to the results presented in Table VI, 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 a unidirectional Granger causality from Thailand's stock return to the stock returns of Indonesia and the US. Summers (2000) argues that a

financial crisis in one country, however big or small, can adversely and psychologically affect investors' perceptions and expectations in other countries. Investors' reactions to acute market shocks when coincided with unwise government policy responses can influence the other markets. For example the Asian crisis influenced the other stock markets in the world (including the US market) as investors started panicking that the financial downturn could also engulf their market due to knock-on effects across international markets. This could partially explain why the stock market return in such a small country such as Thailand Granger causes the return in the US market.

We also found a bidirectional Granger causality between the market stock returns in Thailand and its three neighbouring countries (that is, 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 affect 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 which occurred in 1997, central bankers and individual investors must keep abreast of new developments in international stock markets — particularly those for which we found the evidence of bidirectional and unidirectional causality.

Table IV. 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.

5. Conclusions

This study examines 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 to December 2005. In addition to the Engle-Granger two-step procedure, we used the Gregory and Hansen (1996) test, which allows for a structural break in the cointegration vector.

Based on the cointegration results, we found no evidence of long-run relationships 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 leveraged 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 pair-wise unidirectional Granger causalities, whereby the returns in Hong Kong, the Philippines and the UK can Granger cause the return in Thailand. Based on these results, the performance of stock markets in Honk Kong, the Philippines and the UK may have a direct bearing on the Thai stock market. However, there were also two unidirectional Granger causalities running from Thailand to Indonesia and the US. Thus any abnormal movement in Thailand's stock returns could lead to similar changes in Indonesia and the US. Third, we found evidence of a bidirectional Granger causality between the stock returns in Thailand and those of three of its neighbouring countries (that is, Malaysia, Singapore and Taiwan). The reported causality test results are useful for any assessment of the Asian stock markets. For example, the interplay between these three pairs of countries (Thailand–Malaysia, Thailand–Singapore and Thailand–Taiwan) can be useful for central bankers and international investors alike in evaluating stock market performance.

The empirical results presented in this paper support the view that international investors have long-run opportunities for portfolio diversification by acquiring stocks from these eleven countries. However, in the short-run the scope of these opportunities is rather limited due to systematic and transitory fluctuations which are inherent to stock markets as evidenced by the causality test results.

Table V. 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 and 1 per cent levels of significance for any pair of countries.

Table VI. 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 $H_o: \lambda_1 = \lambda_2 = ... = \lambda_{k1} = 0$ or $H_o': \lambda_1' = \lambda_2' = ... = \lambda_{k1}' = 0$	
		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	$\Delta \ln P_t^{HK} \longrightarrow \Delta \ln P_t^{TH}$	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	$\Delta \ln P_t^{TH} \longrightarrow \Delta \ln P_t^{IN}$	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 $\Delta \ln P_t^{TH} \longleftrightarrow \Delta \ln P_t^{MA}$	$\Delta \ln P_t^{MA} \longrightarrow \Delta \ln P_t^{TH}$	1.870**	0.046
	$\Delta \ln P_t^{TH} \longrightarrow \Delta \ln P_t^{MA}$	3.771***	0.000
Unidirectional causality	$\Delta \ln P_t^{PH} \longrightarrow \Delta \ln P_t^{TH}$	1.936**	0.049
No causality	$\Delta \ln P_t^{TH} \nrightarrow \Delta \ln P_t^{PH}$	1.628	0.110
Bidirectional causality $\Delta \ln P_t^{TH} \longleftrightarrow \Delta \ln P_t^{SG}$	$\Delta \ln P_t^{SG} \longrightarrow \Delta \ln P_t^{TH}$	2.322*	0.076
	$\Delta \ln P_t^{TH} \longrightarrow \Delta \ln P_t^{SG}$	2.633*	0.051
Bidirectional causality $\Delta \ln P_t^{TH} \longleftrightarrow \Delta \ln P_t^{TA}$	$\Delta \ln P_t^{TA} \longrightarrow \Delta \ln P_t^{TH}$	2.690**	0.011
	$\Delta \ln P_t^{TH} \longrightarrow \Delta \ln P_t^{TA}$	1.798*	0.090
Unidirectional causality	$\Delta \ln P_t^{UK} \longrightarrow \Delta \ln P_t^{TH}$	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	$\Delta \ln P_t^{TH} \longrightarrow \Delta \ln P_t^{US}$	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.

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