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Time series analysis for the interest rates relationships among China, Hong Kong, and Taiwan money markets

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Abstract

This paper aims at investigating the regional financial integration of Chinese economic area (CEA), embracing Taiwan, Hong Kong and China after the Asian Financial Crisis. We empirically find that, in the long-run, the money markets of Taiwan, Hong Kong and China have to be mutually linked in order to ensure a long-run equilibrium relationship in the Chinese economic society. In the shot-run, China's interest rate has a significant impact on the interest rates of Taiwan and Hong Kong, but not conversely. The China money market is more rigid due to the existence of market barriers and the China's economic boom in 1990s. Since Taiwan and Hong Kong are typical island-style economies with high degree of market liberalization and their business people have invested huge amount in China market in recent years, they tend to be more sensitive to the innovations and the volatilities from Mainland China market.

JEL classification: C32; G15

Keywords: Interest rate; Money market; Financial integration; CEA; G-VAR

1. Introduction

Nowadays there is a trend that many countries are integrated into one economic organization in a specific geographic region in order to promote sustained economic growth and advance trade and investment. Such examples can be found in the European

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Economic Community (EEC) and the European Free Trade Association (EFTA) in Europe, and the North America Free Trade Association (NAFTA), which involves America, Canada and Mexico in North America.

As the whole world took a close watch on the competence of the dominant power between the two areas mentioned above, there was another market gradually emerging from the Asia-Pacific area. Although there exist many differences in politics and economic development among those countries in the Asia-Pacific area, the economic relationships among one another were getting closer and closer. In 1989, the Asia-Pacific Economic Cooperation (APEC), which now consists of 21 member economies, was formed according to the Australian Government's proposal. Today, APEC has since become the primary regional vehicle for promoting open trade and the practical economic cooperation. Within the last two decades, there indeed has been astonishing performance and rapid growth in economics and open trade in Asia-Pacific area.

Recently, China has officially become a member of the World Trade Organization (WTO) after 15 years of longstanding negotiations. As a result of the negotiations, China has agreed to undertake phased-in series of important commitments to open and liberalize its regime in order to better integrate into the world economy. Therefore, the Asia-Pacific financial markets including China's huge market have become more and more important and should not be ignored. The Asian emerging markets as well as the European and North American markets have now gradually become three major financial markets in the world.

Although the Asian Financial Crisis in 1997 struck the economy of many Asian countries, the financial and economic conditions in these countries have since gradually recovered. Foreign capital investments in China also have become stagnant after the Asian Financial crisis, but these investments are increasing gradually in anticipation of China's entry into the WTO. Hong Kong, part of China's territory, would also benefit from China's accession to the WTO. Because most of China's market access restrictions will be phased out under its WTO commitments, Hong Kong manufacturers operating on the Mainland China will have more opportunities to sell their goods in China's vast domestic markets.

Taiwan, a neighboring country, across the Taiwan Strait, officially became a WTO member on January 1, 2002, immediately following China's accession to the WTO. As both sides gradually ease general trade and investment barriers, further cross-Strait integration and specialization seem inevitable. There is a thought of the Asian economic integration, so-called ''Chinese Economic Area'' (CEA), mainly based on the region of the Chinese society. For example, Cui (1998) found that most of the multinational corporations (MNCs) have practically viewed ''China, Hong Kong, and Taiwan'' as a CEA for the past two decades. Nonetheless, Wang and Schuh (2002) investigated the effects of economic integration of the CEA. The results argued that the three Chinese economies would benefit greatly from further integration by means of liberalizing trade policies.

As mentioned above, China now is playing an increasingly important role in the global financial market, many business people in Hong Kong started to fear that Hong Kong's traditional status as a ''middleman'' might be replaced by more competitive and wide-open China markets. Taiwan, with its huge commercial opportunities, still believes it has become more and more difficult to compete with China. However, no one can neglect the fact that the relationships among three regions are getting closer than ever and all of them would benefit from their regional integration.

De Brouwer (1999) presented positive views on financial integration in East Asia. It argued that economies in the region have financially integrated themselves with the rest of the world over a long period. Moreover, Frankel (1996) pointed out that ‘‘Regionalism’’ may offer a quicker and more efficient way to achieve the goal of worldwide free trade. The regional liberalization does also help build political support for liberalization more generally.

The purpose of this paper is to further investigate the regional financial integration of CEA, embracing Taiwan, Hong Kong and China, after the Asian Financial Crisis (AFC). We employ various time series methodologies to fully investigate the dynamic relationships among the interest rates of these three regions. Some deliberate methods including unit-root test with structural change and the generalized vector autoregressive (G-VAR) model are applied to make the results to be more persuasive. The results of this paper are designated to provide the international investors and policy makers useful information as to make better decisions in the so-called ‘‘Great Chinese Economic Area.’’

The remainder of this paper is organized as follows: Section 2 reviews some relevant literature; Section 3 describes the collected data; Section 4 deals with methods of research and presents the empirical results; and Section 5 concludes this paper.

2. Literature review

For the past few decades, articles have concerned the integration of world money markets and discussed the topics of yields on assets denominated in different geographic markets and their interrelationships. The earlier studies mostly examined the relationships between yields in domestic and offshore markets for US dollars since US dollar played the major role in the international money markets. For examples, Hendershott (1967) and others, all assumed the dominance of the US dollar over other money markets in the world in their study of the international transmission of interest rates. Later works, such as Giddy, Dufey, and Min (1979), Kaen and Hachey (1983), and Swanson (1987) allowed for the two-way feedback effects by employing Granger (1969) causality test for their investigations on the relationships among multinational interest rates.

The works on examining the international transmission of money markets or capital markets can also be found in numerous literatures, e.g., Swanson (1987), Lin and Swanson (1993), Hsieh, Lin, and Swanson (1999), Phylaktis (1999) and Bremnes, Gjerde, and Sattem (2001). Swanson (1987) examined the relationship between Eurodollar deposit rates and yields on negotiable US certificates of deposit (CD) for the period of 1979–1983. The results strongly indicate that the degree of integration increased throughout the period. While the US CD rates is found to be dominant over Eurodollar deposit rates. Lin and Swanson (1993) employed an error correction model to assess the relationships among five major world currencies (US dollar, British pound, German deutsche mark, Swiss franc, and Japanese yen). Their major findings are: (1) the domestic markets of five major world currencies and London Eurodollar market are not integrated completely; (2) the domestic markets of five major world currencies and Singapore Eurocurrency market are not integrated too. Hsieh et al. (1999) extended the former study by examining the relationships among returns on each of five major world currencies in three different geographic markets (the US domestic market, the Euromarket in London, and the Euromarket in Singapore). The results showed that, whatever the currency will be, there always exist long-run equilibrium relationships among currency returns in three separate geographic markets.

Some basic theories have been applied for international diversified portfolio. After Markowitz (1952) developed the models of diversified portfolio, many studies continued the research of this area on international portfolio. The examples were Agmon (1972), Lessard (1974), Makridakis and Wheelwright (1974). The main topics included international capital market structure, the correlation coefficients of stock market indices of different countries, comovements among financial markets, and welfare gain from internationally diversified portfolio.

Among three different theories of international capital market structure: segmentation, integration, and weakly segmentation, Agmon (1972) presented the approach of integration theory. It considered the international capital markets as one multinational perfect capital market—an integrated market. Under this hypothesis, the different national markets were highly correlated with one another and were affected by one international factor. In the new equilibrium, the returns of investing securities in different countries were standing on the same capital market line. It is impossible to attain potential welfare gain from international diversification; that is, the strategy of any investment in this ‘‘integrated’’ international market could not earn additional return through international diversification. This ‘‘one market’’ hypothesis has the advantage of being consistent with much of the accepted ‘‘cointegration’’ economic theory.

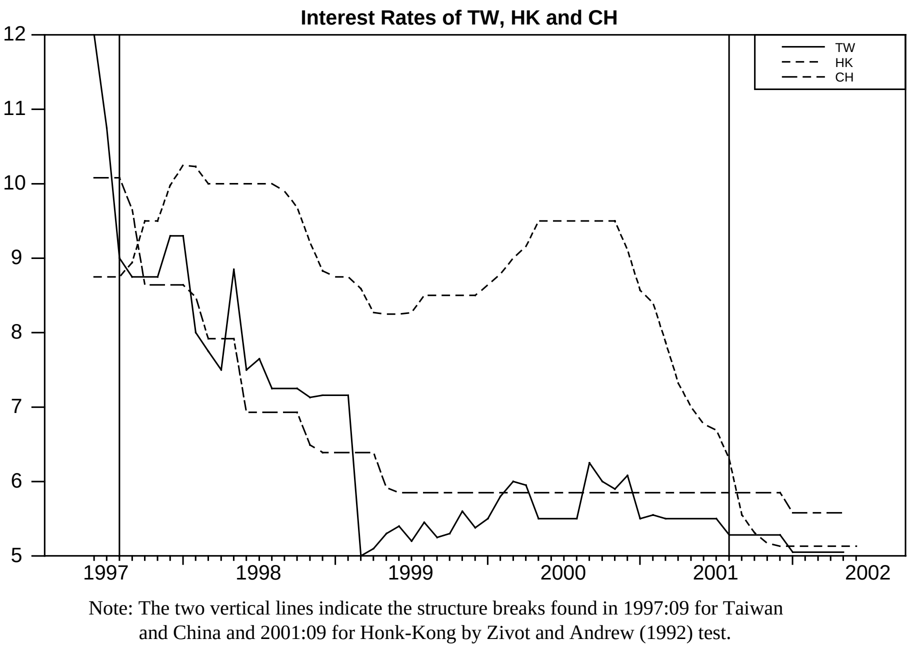
For the international transmission of interest rates, Bremnes et al. (2001) applied Johansen’s (1988) multivariate cointegration test, combining with the impulse response and variance decomposition, to analyze the relationships among short-term and long-term interest rates in the United States, Germany and Norway. The analyses illustrate that the US interest rates have a significant influence on both German and Norwegian interest rates, while the reverse effect is modest. Karfakis and Moschos (1990) and Katsimbris and Miller (1993) investigated the interest rate convergence among Exchange Rate Mechanism (ERM) member countries in the European Monetary System (EMS). Fountas and Wu (1998) further considered the existence of structure break testing for the interest rate convergence in EMS. Among Asian countries, Chinn and Frankel (1995) investigated the relative influence of US and Japanese real interest rates in the determination of local Pacific Rim rates by applying the Johansen’s (1988) multivariate cointegration test. The results indicate that Hong Kong, Malaysian and Taiwanese interest rates are cointegrated with both US and Japanese real rates, while only the Singapore rate is solely cointegrated with the US rate. Phylaktis (1999) examined the extent of capital market integration in a group of Pacific Basin countries (Singapore, Malaysia, Hong Kong, Korea, Taiwan and Japan) following the deregulation of their markets. The results showed that there existed long-run equilibrium relationships among capital markets in Asia-Pacific region and those Pacific Basin countries after deregulation process had greater capital market integration with Japan than with the US. The financial influence of Japan in the region had overtaken that of the US. The study of ‘‘Financial Integration in East Asia’’ by De Brouwer (1999) also pointed out that the integration process in East Asia is based on the sensitivity of domestic interest rates—rather than capital flows—to changes in world interest rates.

3. Data

This paper aims at figuring out the extent of regional economic integration after the Asian Financial Crisis and the interrelationships among the money markets of the three economies mentioned above through employing time series methodologies. The research is done on the basis of the interest rates of the three money markets considered to analyze the extent of comovement and causal relations among them. Data are downloaded from the AREMOS Statistical Data Bank of Ministry of Education, Taiwan. Monthly data runs from July 1997 to June 2002. A total of 60 observations are obtained for each rate. The interest rates selected are all major rates which highly reflect the financial environment of these three region's money markets. They are inter-bank rate for Taiwan and discount rates for Hong Kong and China. We choose inter-bank rates and discount rates because of that changes in monetary policies of countries considered have immediate effects on these rates. Besides, changes in these rates will adequately reflect underlying changes in the marginal cost of borrowing from the banking sector, the leading source of short-term finance to business. Fig. 1 shows the movement of these rates during sample period selected.

4. Methodologies

This paper employed a variety of time series methodologies to fully investigate the interrelationships among three rates considered.



Note: The two vertical lines indicate the structure breaks found in 1997:09 for Taiwan and China and 2001:09 for Honk-Kong by Zivot and Andrew (1992) test.

Fig. 1. Interest rates movement of Taiwan, Hong Kong and China.

Among various testing strategies, this paper tests for ''stationarity'' of each interest rate by first applying Dickey and Fuller's (1981) ADF three-model tests since, according to Schwert (1989), the ADF test with long lags is superior to the others.[1] The ADF three models are express as the following forms:

$$\Delta y_t = \phi y_{t-1} + \sum_{i=1}^{p-1} \beta_i \Delta y_{t-i} + \varepsilon_t \tag{1}$$

$$\Delta y_t = \alpha + \phi y_{t-1} + \sum_{i=1}^{p-1} \beta_i \Delta y_{t-i} + \varepsilon_t \tag{2}$$

$$\Delta y_t = \alpha + \phi y_{t-1} + \gamma t + \sum_{i=1}^{p-1} \beta_i \Delta y_{t-i} + \varepsilon_t \tag{3}$$

Model (1) is a pure random walk with the lag terms. Model (2) possesses a drift. Model (3) includes a drift and a time trend. The null hypothesis for ADF test is: H_0: $\phi = 0$, with the alternative H_1: $-2 < \phi < 0$.

Hamilton (1994) and Elder and Kennedy (2001) both argued that a strategy is necessary to determine which of the three ADF models should be employed in conducting the unit-root test.[2] In this paper, we follow the determining rule by Doldado, Jenkinson, and Sosvilla-Rivero (DJS) (1990) to determine the appropriate model for each rate.[3] Moreover, since the estimation might be biased if the lag length is pre-designated without rigorous determination, this paper adopts newly developed Modified Akaike's information criterion (MAIC) suggested by Ng and Perro (2001) to select the optimal number of lags based on the ''principle of parsimony.''

Table 1 presents the results of the ADF tests, which shows that only the rate of Hong Kong has unit-root in the level and is rejected to be ''non-stationary'' in the first difference, which insures an $I(1)$ type series. However, the rates of Taiwan and China are both $I(0)$ series.

Even the post-AFC period is considered in this study, the suspected structure break may still exist in our sample period. Since there are a number of econometricians arguing that the standard ADF tests are not appropriate for testing for the stationarity of series encountering structure change, we thus further take structure break into account when employ the unit-root test.

Perron (1989) argued that the existence of structure changes tends to bias the finding from ADF tests and developed a model to test the hypothesis that a given series has unit-root with an exogenous structure break which occurs at time T_B. Further elaborated work

[1] Ayat and Burridge (2000) also argued that the ADF test has become the most popular of many competing tests in the literature.

[2] Elder and Kennedy (2001) also developed a student-friendly testing strategy in determine the appropriate ADF model by applying traditional OLS method.

[3] The determining rule by DJS (1990) is to test for the significance of trend coefficient in the third model first, followed by testing for the significance of the drift coefficient in the second model. If both outcomes result in insignificant, the first model is select.

Table 1
The results of ADF unit-root tests (period July 1997 to June 2002)

	Level			First difference		
	$\tau(0)$	$\tau_\mu(0)$	$\tau_\tau(0)$	$\tau(1)$	$\tau_\mu(1)$	$\tau_\tau(1)$
Taiwan	−1.637 [5]	**−2.781** [1]*	−6.40 [3]***	**−4.376** [6]***	−3.059 [8]**	−13.368 [0]***
Hong Kong	−0.978 [1]	−0.384 [1]	**−2.082** [1]	**−1.902** [1]*	−2.046 [4]*	−2.111 [4]
China	−1.998 [2]*	**−3.535** [3]**	−2.215 [4]	**−12.191** [0]***	−12.434 [0]***	−12.263 [0]***

Note. $\tau(0)$, $\tau_\mu(0)$, and $\tau_\tau(0)$ are the test statistics for a unit-root in the level without constant, with constant, and with both constant and trend, respectively. $\tau(1)$, $\tau_\mu(1)$, and $\tau_\tau(1)$ are the test statistics for a unit-root in the difference without constant, with constant, and with both constant and trend, respectively. The critical values for the ADF t statistics are from the Mackinnon (1991) table. The numbers within the square bracket are the appropriate lag lengths for each interest rate based on MAIC suggested by Ng and Perro (2001). The bold numbers indicate the appropriate model determined by DJS (1990).

* Significant at 1% level.

** Significant at 5% level.

*** Significant at 10% level.

by Zivot and Andrew (1992) (hereafter ZA) overthrew the presumed exogenous break point and developed a unit-root test with endogenous structure break, which has been regarded as a more suitable test to test for the order of integration of series. The three models of ZA tests are expressed as following equations:

$$\text{Model A}:\ \Delta Y_t = \mu_1^{\mathrm{A}} + \gamma_1^{\mathrm{A}} t + \mu_2^{\mathrm{A}} DU_t(\lambda) + \alpha^{\mathrm{A}} Y_{t-1} + \sum_{j=1}^{k-1} \beta_j \Delta Y_{t-j} + \varepsilon_t \tag{4}$$

$$\text{Model B}:\ \Delta Y_t = \mu_1^{\mathrm{B}} + \gamma_1^{\mathrm{B}} t + \gamma_2^{\mathrm{A}} DT_t^*(\lambda) + \alpha^{\mathrm{B}} Y_{t-1} + \sum_{j=1}^{k-1} \beta_j \Delta Y_{t-j} + \varepsilon_t \tag{5}$$

$$\text{Model C}:\ \Delta Y_t = \mu_1^{\mathrm{C}} + \gamma_1^{\mathrm{C}} t + \mu_2^{\mathrm{C}} DU_t(\lambda) + \gamma_2^{C} DT_t^*(\lambda) + \alpha^{\mathrm{C}} Y_{t-1} + \sum_{j=1}^{k-1} \beta_j \Delta Y_{t-j} + \varepsilon_t \tag{6}$$

where $DU_t(\lambda)$ is 1 and $DT_t^*(\lambda) = t - T\lambda$ if $t > T\lambda$, 0 otherwise. $\lambda = T_{\mathrm{B}}/T$ and T_{B} represents a possible break point. Model A allows for a change in the level of the series, Model B allows for a change in the slope of trend of a series, while Model C combines both changes in the level and the slope of trend. Since the appropriate model and the optimal lag lengths are crucial in interpreting the results of the tests, we adopt the findings from the ADF tests to select the model and the lag lengths for the ZA tests.

The results of the ZA tests are presented in Table 2 and shown in Fig. 2. All three rates carry unit-root in the level and reject the null of "non-stationarity" in the first difference. This insures the $I(1)$ type series for all the interest rates considered.

Various methods of estimating cointegration have been applied to capture the long-run equilibrium relationships among variables. Among those, Johanson methodology based on the likelihood ratio with non-standard asymptotic distributions involving integrals of

Table 2
The results of ZA unit-root tests with structural break (period July 1997 to June 2002)

	Break	Level ($t(\hat{\lambda}_{\inf})$)	First difference ($t(\hat{\lambda}_{\inf})$)
Taiwan	September 1997	−3.532 (B)	−6.074 (A)***
Hong Kong	September 2001	−4.065 (C)	−5.304 (A)**
China	September 1997	−2.834 (B)	−12.170 (A)***

Note. The critical value for (1, 5, and 10%) levels are (−5.34, −4.80, and −4.58), (−4.93, −4.42, and −4.11), and (−5.57, −5.08, and −4.82) for Model A, Model B, and Model C, respectively, from Zivot and Andrew (1992). The characters within the parenthesis indicate the appropriate model according to the results from ADF test.

** Significant at 5% level.

*** Significant at 10% level.

Brownian motions is found to be the best method to proceed cointegration estimation by Gonzalo (1994).[4]

The elaborate works developed by Johansen (1988, 1994) and Johansen and Juselius (1990) are summarized into five VAR models with ECM, which are presented in the following forms:

$$H_0(r):\ \Delta X_t = \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_{k-1}\Delta X_{t-(k-1)} + \alpha\beta' X_{t-1} + \Psi D_t + \varepsilon_t \quad (1988) \tag{7}$$

$$\mathrm{H}_1^*(r):\ \Delta X_t = \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_{k-1}\Delta X_{t-(k-1)} + \alpha(\beta', \beta_0)(X'_{t-1}, 1)' + \Psi D_t + \varepsilon_t \quad (1990) \tag{8}$$

$$\mathrm{H}_1(r):\ \Delta X_t = \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_{k-1}\Delta X_{t-(k-1)} + \alpha\beta' X_{t-1} + \mu_0 + \Psi D_t + \varepsilon_t \quad (1990) \tag{9}$$

$$\mathrm{H}_2^*(r):\ \Delta X_t = \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_{k-1}\Delta X_{t-(k-1)} + \alpha(\beta', \beta_1)(X'_{t-1}, t)'\mu_0 + \Psi D_t + \varepsilon_t \quad (1994) \tag{10}$$

$$\mathrm{H}_2(r):\ \Delta X_t = \Gamma_1 \Delta X_{t-1} + \cdots + \Gamma_{k-1}\Delta X_{t-(k-1)} + \alpha\beta' X_{t-1} + \mu_0 + \mu_1 t + \Psi D_t + \varepsilon_t \quad (1994) \tag{11}$$

To analyze the deterministic term, Johansen decomposed the parameters μ_0 and μ_1 in the directions of α and $\alpha_\perp$ as $\mu_i = \alpha\beta_i + \alpha_\perp\gamma_i$, thus we have $\beta_i = (\alpha'\alpha)^{-1}\ \alpha'_\perp\mu_i$ and $\gamma_i = (\alpha'_\perp\alpha_\perp)^{-1}\ \alpha'_\perp\mu_i$. The nested sub-models of the general model of null hypothesis $\Pi = \alpha\beta'$ are, therefore, defined as:

$\mathrm{H}_0(r):\ Y = 0$
$\mathrm{H}_1^*(r):\ Y = \alpha\beta_0$
$\mathrm{H}_1(r):\ Y = \alpha\beta_0 + \alpha_\perp\gamma_0$

[4] Gonzalo (1994) compared several methods of estimating cointegration, which include ordinary least squares, nonlinear least squares, maximum likelihood in an error correction model, principle components, and canonical correlations.

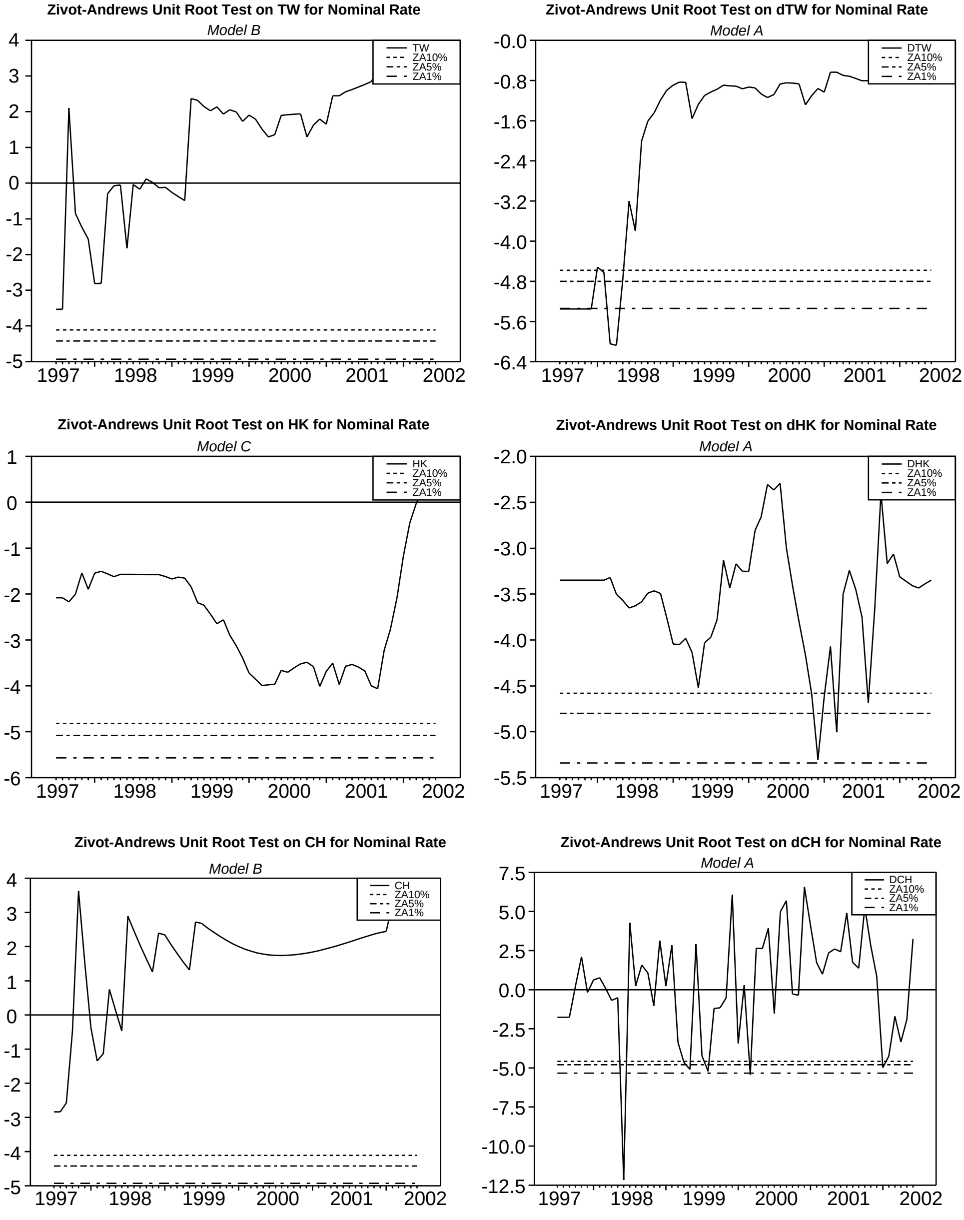


Fig. 2. Plots of $t(\hat{\lambda}_{\text{inf}})$ for level and difference of Zivot and Andrew unit-root test.

$$\text{H}_2^*(r): \; Y = \alpha\beta_0 + \alpha_{\perp}\gamma_0 + \alpha\beta_1 t$$
$$\text{H}_2(r): \; Y = \alpha\beta_0 + \alpha_{\perp}\gamma_0 + (\alpha\beta_1 + \alpha_{\perp}\gamma_1)t$$

Johansen (1994) emphasized the role of the deterministic term, $Y = \mu_0 + \mu_1 t$, which includes constant and linear terms in the Gaussian VAR. Following Nieh and Lee (2001), a

Table 3
Determination of cointegration rank in the presence of a linear trend and a quadratic trend

Rank	Model 1 (H_0)		Model 2 (H_1^*)		Model 3 (H_1)		Model 4 (H_2^*)		Model 5 (H_2)	
	$T_0(r)$	$C_0(5\%)$	$T_1^*(r)$	$C_1^*(5\%)$	$T_1(r)$	$C_1(5\%)$	$T_2^*(r)$	$C_2^*(5\%)$	$T_2(r)$	$C_2(5\%)$
$r = 0$	33.66	24.31	51.31	34.91	50.01	29.68	62.64	42.44	57.90	34.55
$r \leq 1$	**<u>3.28</u>**	12.53	16.26	19.96	15.58	15.41	25.87	25.32	21.20	18.17
$r \leq 2$	0.00	3.84	2.40	9.24	1.91	3.76	7.82	12.25	3.93	3.74

Note. $T_0(r)$, $T_1^*(r)$, $T_1(r)$, $T_2^*(r)$, and $T_2(r)$ denote the LR test statistics for all the null of H(r) vs. the alternative of $H(p)$ of Johansen's five models. $C_0(5\%)$, $C_1^*(5\%)$, $C_1(5\%)$, $C_2^*(5\%)$ and $C_2(5\%)$ are 5% LR critical values for Johansen's five models, which are extracted from Osterwald-Lenum (1992). The model selection follows Nieh and Lee's (2001) decision procedure, diagnosing models one by one until the model that cannot be rejected for the null. The bold number with underline indicates the selection of the rank in the presence of linear trend and quadratic trend. VAR length is 5 for all the models selected based on the smallest number of SBC.

decision procedure among the hypotheses H(r) and H*(r) for five different models is presented in the following procedure:[5]

$$\begin{aligned} &\mathrm{H}_0(0) \rightarrow \mathrm{H}_1^*(0) \rightarrow \mathrm{H}_1(0) \rightarrow \mathrm{H}_2^*(0) \rightarrow \mathrm{H}_2(0) \rightarrow \mathrm{H}_0(1) \rightarrow \mathrm{H}_1^*(1) \\ &\quad \rightarrow \mathrm{H}_1(1) \rightarrow \mathrm{H}_2^*(1) \rightarrow \mathrm{H}_2(1) \rightarrow \cdots \rightarrow \cdots \rightarrow \mathrm{H}_0(p-1) \rightarrow \mathrm{H}_1^*(p-1) \\ &\quad \rightarrow \mathrm{H}_1(p-1) \rightarrow \mathrm{H}_2^*(p-1) \rightarrow \mathrm{H}_2(p-1) \end{aligned}$$

We diagnose models one by one until the model that cannot be rejected for the null. The SBC criterion is adopted in selecting the appropriate lag length for our cointegration test.

The empirical findings for the long-run relationship among China, Hong Kong, and Taiwan money markets are presented in Table 3. It shows that one cointegrating vector exists in these three rates. Based on the decision procedure described above, it presents in Johansen's first model with no linear and quadratic trend.

Even though a long-run equilibrium relationship is found in the comovement of three rates considered, it does not show the pairwise relationships between each pair of these rates. For the suspected relationships between each two-region pair, we further carry out cointegration test for the two-region settings. All the results are presented in Table 4. A surprising finding is obtained that no cointegration relationship exists between rates in every two-region setting. This finding may be interpreted as a phenomenon that, in the Chinese economic society, the money markets of Taiwan, Hong Kong and China have to be mutually linked in order to ensure the cointegration relationship. Once one of the rates of these three regions is ignored, the long-run equilibrium relationship disappears.

The Engle and Granger's (1987) "Granger representation theorem" argued that error correction and cointegration are equivalent representations. For the relationship without the existence of cointegration, the unrestricted VAR can be applied to test for the causal relation between variables. Considering two series, X_t and Y_t, the Granger's (1969)

[5] Johansen (1992, 1994) developed a testing procedure based on the ideas developed by Pantula (1989) to determine the number of cointegrating rank in the presence of linear trend (Johansen, 1992) and quadratic trend (Johansen, 1994). Nieh and Lee (2001) further reorganized them in a more friendly way.

Table 4
Cointegration test for the two-region settings

Rank	Model 1 (H_0)		Model 2 (H_1^*)		Model 3 (H_1)		Model 4 (H_2^*)		Model 5 (H_2)	
	$T_0(r)$	$C_0(5\%)$	$T_1^*(r)$	$C_1^*(5\%)$	$T_1(r)$	$C_1(5\%)$	$T_2^*(r)$	$C_2^*(5\%)$	$T_2(r)$	$C_2(5\%)$
Panel A: Taiwan and Hong Kong										
r = 0	<u>11.08</u>	12.53	11.87	19.96	8.70	15.41	14.31	25.32	12.09	18.17
$r \leq 1$	2.44	3.84	2.52	9.24	0.54	3.76	4.86	12.25	4.82	3.74
Panel B: Taiwan and China										
$r = 0$	<u>7.12</u>	12.53	15.79	19.96	15.25	15.41	16.00	25.32	15.41	18.17
$r \leq 1$	0.18	3.84	6.86	9.24	6.38	3.76	7.06	12.25	6.99	3.74
Panel C: Hong Kong and China										
$r = 0$	<u>8.21</u>	12.53	13.88	19.96	11.92	15.41	17.26	25.32	12.83	18.17
$r \leq 1$	2.14	3.84	2.21	9.24	0.67	3.76	5.94	12.25	4.56	3.74

Note. $T_0(r)$, $T_1^*(r)$, $T_1(r)$, $T_2^*(r)$, and $T_2(r)$ denote the LR test statistics for all the null of H(r) vs. the alternative of $H(p)$ of Johansen's five models. $C_0(5\%)$, $C_1^*(5\%)$, $C_1(5\%)$, $C_2^*(5\%)$ and $C_2(5\%)$ are 5% LR critical values for Johansen's five models, which are extracted from Osterwald-Lenum (1992). The model selection follows Nieh and Lee's (2001) decision procedure, diagnosing models one by one until the model that cannot be rejected for the null. The bold number with underline indicates the selection of the rank in the presence of linear trend and quadratic trend. Lag length is 7, 5 and 4 for Panel A, Panel B and Panel C, respectively, by SBC.

bivariate VAR model for the Granger causality test (GC test) is presented in the form as follows:

$$\Delta X_t = \alpha_1 + \sum_{i=1}^{n_1} \alpha_{11}(i)\Delta X_{t-i} + \sum_{j=1}^{m_1} \alpha_{12}(j)\Delta Y_{t-j} + \varepsilon_{Xt} \tag{12}$$

$$\Delta Y_t = \alpha_2 + \sum_{i=1}^{n_2} \alpha_{21}(i)\Delta X_{t-i} + \sum_{j=1}^{m_2} \alpha_{22}(j)\Delta Y_{t-j} + \varepsilon_{Yt} \tag{13}$$

where ε_{Xt} and ε_{Yt} are stationary random processes intended to capture other pertinent information not contained in lagged values of X_t and Y_t. The lag lengths, n and m, are again decided by SBC. The series Y_t fails to Granger cause X_t if all $\alpha_{12}(j) = 0$ ($j = 1, 2, 3, m_1$) and the series X_t fails to cause Y_t. If all $\alpha_{21}(i) = 0$ ($i = 1, 2, 3, n_1$).

The GC test in this paper is to examine if there exists feedback (two-way), or one-way causality in every bivariate pair among the three interest rates. The results of the GC test are shown in Table 5 that there exists apparent two-way causality between rates of Taiwan and China. Among other causal relations, only an unidirectional causality running from the rate of China to the rate of Hong Kong is found. These results dictate that the rate of China provides a leading indicator for the movement of other two rates and the rate of Taiwan shows a feedback effect to the rate of China. The exogeneity ordering of rates of three countries can be arranged as China, Taiwan, and Hong Kong from the relative value of the F statistic.

The more recent researches have largely applied the impulse response functions (IRF) and variance decompositions (VDC) to conquer the difficulty of interpreting the estimated coefficients of a VAR model. An IRF traces the response of one of the innovations on current and future values of the endogenous variables to a one standard deviation shock. This shock to a variable directly affects itself, and is also transmitted to all of the

Table 5
Pairwise Granger causality test for three interest rates

Null hypothesis	F statistic	Probability
HK does not Granger cause TW	2.6072	0.1120
TW does not Granger cause HK	2.4427	0.1237
CH does not Granger cause TW	6.5094*	0.0135
TW does not Granger cause CH	5.0458*	0.0287
CH does not Granger cause HK	4.7856*	0.0329
HK does not Granger cause CH	0.0473	0.8285

Note. HK, TW and CH are the symbols for the interest rates of Hong Kong, Taiwan, and China, respectively. The null hypothesis, H_0, is for "no causal relation." Lag length is 1 selected by SBC.
* Significant at 1% level.

endogenous variables through the dynamic structure of the VAR. On the other hand, VDC decomposes forecast error variance (FEV) in an endogenous variable into percentage shocks to its own and other endogenous variables in the VAR, which in turn offers information about the relative importance (exogeneity ordering) of each random innovation to the variables.

Following Sims (1980) and Hamilton (1994), the reduced form of the structure VAR model: $Bx_t = \Gamma_0 + \Gamma_1 x_{t-1} + \varepsilon_t$, can be transformed to a standard form:

$$x_t = A_0 + A_1 x_{t-1} + e_t$$

where Γ_0 and $A_0 = B^{-1}\Gamma_0$ are vectors of constant; $\Gamma_1, A_1 = B^{-1}\Gamma_1$ and the back operator B are matrices; the white-noise, ε_t, and the disturbance $e_t = B^{-1}\varepsilon_t$ are vectors.

Further derivation can reach to a vector moving average (VMA) representation:

$$x_t = \alpha + \sum_{i=0}^{\infty} A_1^i e_{t-i}, \quad \left[\text{i.e., } x_t = (\mathrm{I} + A_1 + \cdots + A_1^n)A_0 + \sum_{i=0}^{n} A_1^i e_{t-i} + A_1^{n+1} x_{t-n-1}\right]$$

and then a form of white-noise disturbance:

$$x_t = \alpha + \sum_{i=0}^{\infty} \phi_{jk}(i)\varepsilon_{t-i} \tag{14}$$

where α is a vector of constants and all the elements of $\phi_{jk}(1)$, with $\phi_{jk}(0) = \mathrm{I}$, are the "impact multipliers," which examine the interaction over the entire path of three interest rates considered in this paper. Eq. (14) is the so-called impulse response function.

If the disturbance at all lags, ε_{t-1}, are absolute contemporaneously uncorrelated, we can easily find the percentage of the FEV that occurs in the VAR, and then judge the relative exorgeneity of all the presumed endogenous variables. However, it is not always the case.

When applying Choleski decomposition (i.e., multiply disturbance term, ε_{t-1} by a lower triangular matrix V, where $VV' = \mathrm{I}$), a VMA representation: $x_t = \alpha + \sum_{i=0}^{\infty} C_i \varepsilon_{t-i}$, can be transformed into:

$$x_t = \alpha + \sum_{i=0}^{\infty} C_i VV' \varepsilon_{t-i} = \alpha + \sum_{i=0}^{\infty} D_i \mu_{t-i} \tag{15}$$

where C_i is a matrix with, $D_i = C_i V$ and $\mu_{t-i} = V'\varepsilon_{t-i}$.

From Eq. (15), the k-step ahead forecast error of x_t is given by:

$$x_t - \hat{E}_{t-k}x_t = D_0\mu_t + D_1\mu_{t-1} + \cdots + D_{k-1}\mu_{t-k+1} \tag{16}$$

where $\hat{E}_{t-k}x_t = D[x_t|x_{t-k}, x_{t-k-a}, x_{t-k-2}, \ldots]$, implies that utilizing all the information set at period $t-k$ to forecast present value of x_t. The corresponding variance–covariance matrix of this k-step ahead forecast error is expressed as follows:

$$\begin{aligned} E(x_t - \hat{E}_{t-k}x_t)(x_t - \hat{E}_{t-k}x_t)' = D_0E(\mu_t\mu_t')D_0' + D_1E(\mu_t\mu_t')D_1' + \cdots \\ + D_{k-1}E(\mu_t\mu_t')D_{k-1}' \end{aligned} \tag{17}$$

Cooley and LeRoy (1985) criticize that orthogonalized IRF and VDC based on Cholesky decomposition are, in general, not meaningful. King et al. (1991) and Zhou (1996) point out that as there is more than one common trend in a model, different ordering of variables may significantly affect the results of IRF and VDC if the common trends are not absolutely uncorrelated. The newly developed Generalized VAR (the G-IRF and G-VDC) models by Pesaran and Shin (1998) overwhelm the shortcoming of orthogonalized IRF and VDC, which have the advantage of being invariant to the ordering of the variables. Dekker et al. (2001) compared traditional VAR with G-VAR in estimating linkages of the Asia-Pacific equity markets. It was found that the generalized approach significantly gives more realistic results, particularly for those markets with closest geographical and economic links.

Owing to the advantage of being invariant to the ordering of the variables the G-IRF and G-VDC are employed in this study. The mutual impacts of shocks among three rates are first exhibited in Fig. 3. These diagrams of G-IRF show that only self-responses exist for all three rates. Among these self-responses, the rates of Taiwan and China show transitory effects by their own innovations, whereas Hong Kong's rate has a permanent effect cause by one shock of itself.[6] These findings seem to correspond with the findings of break points from Zivot and Andrew (1992) unit-root test shown in Fig. 1 that the rates of Taiwan and China have earlier break points, whereas Hong Kong's rate has a later break point.

We interpret that the three rates are mutually independent. They will not be influenced by the innovations from any other rates. All three money markets are self-reliant and any shock from each money market is not easy to transmit to others.

The further finding of G-VAR is from the implementation of G-VDC. The results are presented in Table 6, which first show a similar finding with the G-IRF that all the FEVs of the three rates are explained mostly by each of themselves. This can be interpreted by the phenomenon that the fluctuations happened in all the rates are caused by their own variations. However, as time goes by, a moderate increasing influence power is found from the rates of China and Hong Kong to Taiwan's rate, but not conversely. As we observed from Table 6, after 1 year (12 months period), the explanatory proportion of the rate of China, describing 19.939% of the FEVs in the rate of Taiwan. It seems to be similar with

[6] For the transmission effects of IRF, Lutkepohl and Reimers (1992) distinguished between transitory and permanent effects by the argument that an effect of a one-time impulse on a variable is called transitory (permanent) if the variable (does not) return(s) to it's previous equilibrium value of zero after some period.

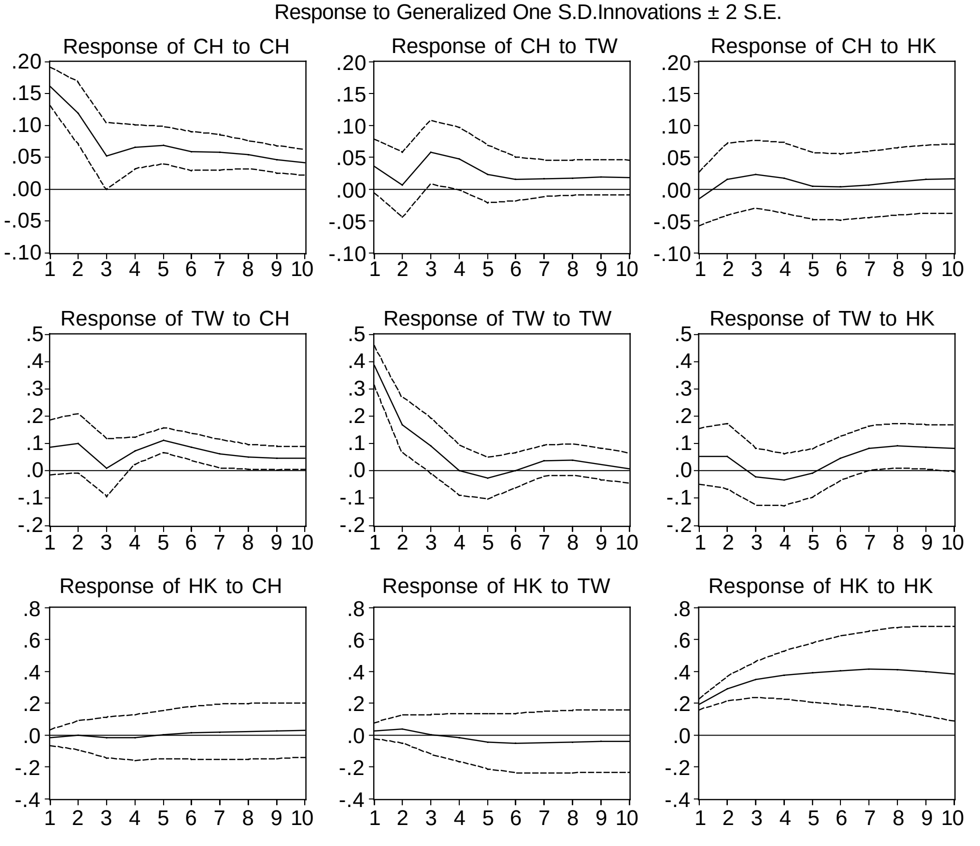


Fig. 3. Generalized-Impulse response to one S.D. innovation.

Table 6
The results of the generalized-forecast error variance decomposition

	Period	China	Taiwan	Hong Kong
China	1	100.000	0.000	0.000
	4	89.640	7.191	3.169
	8	90.935	5.764	3.301
	12	89.841	5.414	4.746
Taiwan	1	4.956	95.044	0.000
	4	11.458	86.792	1.750
	8	19.691	71.072	9.237
	12	**19.939**	**61.072**	**18.989**
Hong Kong	1	0.913	2.659	96.428
	4	0.240	0.678	99.081
	8	0.169	1.171	98.661
	12	0.361	1.278	98.361

Note. Each number is a percentage value. The value of the forecast error variance decomposition describes the explanatory proportion of the changes in each of the interest rate.

the rate of Hong Kong with 18.989% explanatory power in describing the FEVs in Taiwan's rate.

5. Conclusion

This paper aims at investigating the regional financial integration in CEA, mainly the regions of Taiwan, Hong Kong and China, after the Asian Financial Crisis. We employ various time series methodologies to fully investigate the dynamic relationships among the interest rates of these three regions. To be more persuasive, some deliberate methods including unit-root test with structural change and the generalized vector autoregressive (G-VAR) model are considered.

Some conclusions have been drawn from our empirical tests. First, though an ambiguous finding is obtained when testing for the stationarity of all the interest rates from the defective ADF unit-root test, we obtain that all the interest rates of China, Hong Kong, and Taiwan are type $I(1)$ series when we take account of structure break into our unit-root testing model. Second, Johansen maximum likelihood cointegration test shows that one cointegrating vector exists in these three rates, presenting in Johansen's first model with no linear and quadratic trend. However, a long-run equilibrium relationship in the comovement of three rates considered does not imply the pairwise relationships between each pair of these rates. When carry out cointegration test for the two-region settings, a surprising finding is obtained that no cointegration relationship exists between rates in every two-region setting. This outcome indicates that, in the Chinese economic society, the money markets of Taiwan, Hong Kong and China have to be mutually linked in order to ensure the cointegration relationship. Once one of the rates of these three regions is ignore, the long-run equilibrium relationship disappears. Third, from the GC test, an apparent two-way causality between rates of Taiwan and China and an unidirectional causality running from the rate of China to the rate of Hong Kong are found. These findings demonstrate that the rate of China provides a leading indicator for the movement of other two rates and the rate of Taiwan shows a feedback effect to the rate of China. These results describe the progress we have experienced for the past decade that China has successfully attracted Taiwanese investments to its shores after its opening of the trade gate and its booming in 1990s and Hong Kong has provided huge funds for investment in China, especially after 1997s took over. As we observed in recent year, more and more business people from Taiwan and Hong Kong has made their capital budgeting decision in Mainland China. Therefore, lots of funds has moved into China's market and the capital flow among the three regions has interacted more frequently than ever.

We further examine the influence of the ''impact innovation'' by G-IRF and G-VDC with the advantage of being invariant to the ordering of the variables. From G-IRF, three rates are found to be mutually independent. They will not be influenced by the innovations from any other rates. All three money markets are self-responded. However, the rates of Taiwan and China show transitory effects by their own shocks, whereas Hong Kong's rate has a permanent effect caused by one shock of itself. These seem to correspond with the findings of earlier break points for rates of Taiwan and China and a later break point for Hong Kong's rate by Zivot and Andrew (1992) unit-root test. Finally, the results of G-VDC indicate that all the FEVs of the three rates are explained mainly by their own innovations. However, as time goes by, a moderate increasing influence power is found from the rates of China and Hong Kong to Taiwan's rate, but not conversely. The rates of China and Hong Kong are shown to have similar explanatory power in describing the FEVs of the rate of Taiwan.

The overall conclusion can be drawn as the following summary. In the long-run, the money markets of Taiwan, Hong Kong and China have to be mutually linked in order to ensure the long-run equilibrium relationship in the Chinese economic society. In the shot-run, as expected, the China's interest rate has a significant impact on the interest rates of Taiwan and Hong Kong. Conversely, the interest rates of Taiwan and Hong Kong have much less influence on the China's money market. The China money market is more rigid due to the existence of market barriers and China's economic boom in 1990s. On the other hand, since Taiwan and Hong Kong are typical island-style economies with higher degree of market liberalization and their business people have invested huge amount in China market in recent years, they tend to be more sensitive to the innovations and the volatilities from Mainland China market. Moreover, the permanent self-responded effect of Hong Kong's rate reflects its persistently and relatively high interest rate.

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