Convergence of Interest Rates around the Pacific Rim

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Abstract

This paper applies panel data unit root tests to examine the interest rate convergence of small-open Asian countries with major financial centers. With monthly data from 1988M1 to 1997M6, we find that the nominal interest rates of these countries converge to the U.S. rates rather than to Japan's. This finding is consistent with the view that the monetary authorities of non-Japan Asian economies arrange their monetary policies overwhelmingly to follow the monetary policy of the U.S. rather than that of Japan before the financial crisis of 1997.

Keywords: Interest rates; Convergence; Panel data unit-root tests; Bootstrap.

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1. Introduction

The last two decades have witnessed the fact that many small-open Pacific Rim countries have instituted a continuing policy of financial market liberalization, resulting in the globalization of their financial markets. It is also well known that non-Japan Asian economies have formed strong economic ties to the United States through international trade. At the same time, Japan has become one of the most important financial intermediaries with signs that her economic influence is increasing. As a result, the question for small-open Asian countries who are candidates to form a yen block is: Has Japan been gaining financial influence at the expense of the U.S. in this region over the past decade?

A traditional way to answer this question is to investigate whether the financial markets of non-Japan Asian countries are integrated with the U.S. or Japanese financial market. Empirically, most in the literature examine the cointegration between the domestic interest rate and the U.S. rate or the Japanese rate. (see, for example, Bhoocha-Oom and Stansell, 1990, Reisen and Yeches, 1993, Dooley and Mathieson, 1994, and Chinn and Frankel, 1994, 1995). The rationale of the previous approach is rooted in the uncovered interest parity condition (UIP). Assuming that exchange rate changes and the risk premium are both stationary, interest rates must be found to converge for two economies with perfectly capital mobility.

De Brouwer (1999) argues that if countries have open financial markets, then arbitrage occurs with all pairs of interest rates, and not just one of them. Thus, it will be difficult to interpret a case where one finds that countries are financially integrated with the U.S., but not with Japan. Furthermore, de Brouwer (1999) points out that what interest-rate parity tests do is to reveal information about which exchange rate a monetary authority targets over a specific period.[1] If the domestic interest rate converges to country A's rate rather than country B's, then it indicates that the domestic monetary authorities have directed monetary policy to peg against country A's currency rather than B's currency. Given this kind of monetary convergence, it is natural to say that country A has financial influence on the domestic country.[2] Therefore, testing the convergence of interest rates can help us to tell which country, the U.S. or Japan, has a direct financial influence on the non-Japan Asian countries. Moreover, interest rate convergence can be empirically examined by testing the stationarity of the interest differential as explained in the following section.

The purpose of our paper is to empirically examine the interest rate convergence between the (non-Japan) Asia economies and the U.S. or Japan. The paper differs from existing literature in its econometric method, sample period, and variables of interest. First, Chinn and Frankel (1995) and Glick and Hutchison (1990) investigate the existence of financial market integration between some Pacific Basin countries and the U.S. by testing stationarity of real interest rate differentials implied by the real interest-rate parity. Testing the real interest-rate parity is equivalent to examining the joint hypothesis of UIP and purchasing power parity (PPP). However,

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[1] Svennson (1991) and Hurley (1998) show that placing limits on the movement of economic fundamentals in order to maintain an exchange rate within a target zone also puts a target zone around interest rate differentials.

[2] The convergence of interest rates itself is one of the key indicators for monetary convergence, which has been discussed a lot for the European Monetary Union but seldom for Asian economies. See Groeneveld, Koedijk, and Kool (1998) for an investigation into the credibility of European economic convergence as an example.

Wu and Chen (1999) find evidence to reject PPP for Pacific Basin countries while Taylor (2001) provides a reason why a linear test for PPP with low-frequency data can so easily be rejected.[3] Since studies about PPP remain controversial and our interest is in the convergence of interest rates rather than the integration of commodity markets, this paper tests only the hypothesis of UIP in order to keep things simple.

Second, to consider the small-open Asian countries as a whole, we have to make sure that there is no significant difference in the extent of financial liberalization among countries over the sample period. Among the various Asian countries, Hong Kong was the first to liberalize her financial systems in 1973. Singapore deregulated her financial markets and foreign exchange markets in 1978, while Malaysia, the Philippines, and Indonesia undertook their financial reforms in the early 1980s. Among them, Indonesia, cited by the World Bank's 1992 report, is one of the bank's greatest success stories overall in the 1980s. More recently, Thailand, South Korea, and Taiwan began their movement toward liberalization in the mid-1980s. Although the timing and extent of liberalization have varied across countries, all countries in the region allowed domestic and foreign market forces to play a greater role in their financial markets by the late 1980s (Glick and Hutchison, 1990). Therefore, our sample period begins in 1988M1 in order to lower the influence of an uneven deregulation degree among these countries.

Many of the Asian countries which encountered the 1997 Asian financial crisis, on the other hand, saw that the crisis had a different influence on their monetary indicators. Given current suspicions that some monetary authorities of Asian

[3] Frankel (1986) points out that the major reason for rejecting the real interest-rate parity for industrialized countries is the failure of PPP.

economies have turned to peg the yen rather than the dollar and that there is growing popularity of a control in capital flow, studies in the changes of the monetary policy direction are obviously interesting. Nonetheless, these changes cannot easily be identified due to the short time span of data (4 years from the end of the crisis). In addition, some countries' currencies were seriously attacked by the 1997 crisis while some others were only mildly influenced. It would be hard to interpret the meanings of the results from the full sample panel test. We therefore end this paper's sample period in 1997M6 so as to isolate the impact of the 1997 Asian financial crisis.

As for the econometric method, the main feature of Chinn and Frankel (1995) and Phylaktis (1999) is the consideration of nonstationary data that is neglected in other studies. They apply the conventional ADF test and cointegration test to Pacific Rim rates. It is well known that the power of the conventional ADF test is low when the sample span is short (Shiller and Perron, 1985). The panel data unit-root test of Levin, Lin and Chu (2002) has thus recently been widely applied in empirical literature due to its high power relative to conventional single-equation based tests (abbreviated to the LLC test hereafter). However, the LLC test has been criticized for assuming the same long-run multiplier across countries under the alternative hypothesis. To address this limitation, Im, Pesaran, and Shin (2003) propose a new panel data unit-root test based on the mean group approach (abbreviated to the IPS test hereafter), and find that their test achieves a more accurate size and higher power relative to the LLC test by allowing for a greater degree of heterogeneity across individuals.

However, Taylor and Sarno (1998) and Karlsson and Lothgren (2000) point out a potential pitfall in applying panel unit root tests. The pitfall is: panel unit root tests may lead to a very high probability of rejection of the joint unit-root null when there is a single stationary series in a system of an otherwise unit-root process. To avoid this pitfall, Taylor and Sarno (1998) suggest using the JLR test as a complement to panel unit root tests. The null hypothesis of the JLR test is that there is at least one non-stationary series in the panel. This null is only violated if all of the series are stationary. Therefore, the rejection of the null with the JLR test implies that all series are stationary. This paper thus employs both IPS test and JLR test to re-examine the interest-rate convergence of Pacific Rim countries.

We focus our analysis on the interest-rate convergence of small-open Asia countries as a whole vs. the U.S. and Japan, respectively. Therefore, our empirical strategy benefits from using cross-sectional information, and is in sharp contrast to that in existing literature, in which several pairs of countries are analyzed individually. By using more powerful tests, a more satisfying sample period, and more appropriate variables, this paper finds that the interest rates of non-Japan Asian countries converge relatively to the U.S. interest rate instead of Japan's.

The remainder of the paper is organized as follows. Section 2 briefly describes the econometric strategy. Section 3 reports empirical results from the IPS test and the JLR test. Critical values of the IPS test are then simulated by Bootstrap to correct for a small sample bias and to take into account the presence of serial correlation and contemporaneous correlation. Conclusions are summarized in the final section.

2. Model and Econometric Methodology

If the capital mobility between domestic and foreign countries is perfect, then UIP will hold as follows:

$$\begin{aligned} i_t - i_t^* &= E[s_{t+1} - s_t \mid I_t] + \eta_t, \\ &= \Delta s_{t+1} + \varepsilon_t + \eta_t, \end{aligned}$$

where i_t and i_t^* are the domestic and foreign nominal interest rate, respectively; s_t is the logarithmic nominal exchange rate; $E(\cdot | I_t)$ is the conditional expectation based on the information set at time t, I_t; η_t is a time-varying risk premium; Δs_t is the nominal exchange rate change; and ε_t is the forecast error which is a white noise under the assumption of rational expectation. It is well known that s_t is I(1) and hence $s_{t+1} - s_t$ is I(0) (see Meese and Singleton, 1982).[4] As long as η_t is a covariance stationary process, under the assumptions of rational expectations and that the exchange rate is difference-stationary, the interest rate differential $i_t - i_t^*$ is thus stationary.[5] This has a natural implication for defining interest rate convergence.

For a pool of countries, the definition of interest-rate convergence can be similarly defined as in the case of a two-country world. Consider regions 1 to N+1. Let $i_{k,t}$ denote the nominal interest rate in country k at time t. Also, let $\delta_{k,t} \equiv i_{N+1,t} - i_{k,t}$ (for k = 1, …, N) be the interest differential between country N+1 and k.

Definition: Interest differentials between country N+1 *and* k, *for* k = 1, …, N, *converge if* $\lim_{j \to \infty} E(\delta_{k,t+j} | I_t) = c_k$, *where* c_k, k = 1, …, N, *is a constant.*

This definition is modified from Bernard and Durlauf (1995), which leads us to ask whether the long-run forecasts of interest differentials tend to be constant when the forecasting horizon increases infinitely. It is quite straightforward to see that the

[4] Evidence supporting that nominal exchange rates are I(1) for Pacific Basin countries can be found in Chen and Wu (1997), for example.

[5] The assumption of a stationary risk premium can be found in Edison and Paul (1993) and Wu (1999).

definition of convergence will be satisfied when interest differentials are stationary. Consider the following specification:

$$\delta_{k,t} = \mu_k + \rho_k \delta_{k,t-1} + e_{k,t}, \qquad (1)$$

where μ_k and ρ_k are constants and $e_{k,t}$ is a zero-mean stationary process. Solving equation (1) recursively, we derive

$$\delta_{k,\,t+j} = \sum_{i=0}^{j-1} \rho_k^i \mu + \rho_k^i \delta_{k,\,t} + \sum_{i=0}^{j-1} \rho_k^i e_{k,\,t+j-i}\,.$$

Hence, $\lim_{j\to\infty} E(\delta_{k,t+j} | I_t) = \mu_k/(1-\rho_k)$, if $0 \le |\rho_k| < 1$.

We therefore employ the IPS test to examine the stationarity of interest differentials. Suppose that there are N interest differentials in a panel and they have the following autoregressive representation:

$$\delta_{it} = \mu_i + \rho_i \delta_{it-1} + \eta_{it}, \;\; i = 1,\ldots, N.$$

The IPS test examines the null hypothesis: [6]

$$H_0: \; \rho_1 = \rho_2 = \ldots = \rho_N = 1,$$

against

$$H_a: \; \rho_i < 1, \text{ for some i.}$$

The possible serial correlation in η_{it} can be corrected using the ADF method. Hence, the model to be estimated is given as follows:

$$\Delta\delta_{it} = \mu_i + \phi_i \delta_{it-1} + \sum_{j=1}^{w_i} \kappa_{ij} \Delta\delta_{it-j} + \zeta_{it}, \qquad (2)$$

where $\Delta\delta_{it} = \delta_{it} - \delta_{it-1}$, and ζ_{it} is assumed to be uncorrelated over time.

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[6] The null hypothesis in Levin, Lin and Chu (2002) is $\rho_1 = \rho_2 = \ldots = \rho = 1$, while the alternative hypothesis is $\rho_1 = \rho_2 = \ldots = \rho < 1$. The alternative hypothesis is implausible in empirical analysis, because it means that each interest rate reverts to its respective unconditional mean over time at the same rate. The same criticism can also be applied to Abuaf and Jorion (1990).

Now define the t statistic of $\hat{\phi}_i = 0$ based on T observations, in (2) as $t_{iT}(w_i)$. The IPS's t-bar statistic is defined as follows:

$$W_{tbar} = (\bar{t}_{NT} - a_{NT}) \Big/ \sqrt{b_{NT}/N}\,, \tag{3}$$

where $\bar{t}_{NT} = (1/N)\sum_{i=1}^{N} t_{iT}(w_i)$, $a_{NT} = (1/N)\sum_{i=1}^{N} E[t_{iT}(w_i)]$, and

$b_{NT} = (1/N)\sum_{i=1}^{N} V[t_{iT}(w_i)]$, while $E[t_{iT}(w_i)]$ and $V[t_{iT}(w_i)]$ are the mean and variance of $t_{iT}(w_i)$, respectively. The limiting distribution of W_{tbar} here is a standard normal one. The adjustment terms in W_{tbar}, a_{NT} and b_{NT}, depend on the selection of the lag length w_i and the time span T, and they are generated by simulation through 100,000 iterations.

If the ζ_{it}'s in (2) are contemporaneously correlated, then the limiting distribution of $\bar{z}_{1NT}$ is unknown. The finite sample critical values of W_{tbar} are therefore simulated by a bootstrap in which residuals are allowed to have both serial and contemporaneous correlation. A detailed discussion of our bootstrap procedure is given in Appendix 1.

Note that the alternative hypothesis of the IPS test is "some series in the panel are stationary." In the case of rejection, we are not sure whether all of the series in the panel are stationary or some of them are non-stationary series. Taylor and Sarno (1998) thus provide a JLR test with the null hypothesis that at least one of the series in the panel is non-stationary as a complement. The JLR test is a special application of Johansen's (1988) maximum likelihood procedure for testing for the number of cointegrating vectors in a system. Under the null hypothesis that at least one of the series is a realization of a unit root process, the JLR test has a known limiting χ^2 distribution with one degree of freedom. To correct for small sample bias, we follow the method proposed by Taylor and Sarno (1998) to construct the 10% finite sample critical values. The discussion of the JLR test is not described here, but can be found in Taylor and Sarno (1998).

3. Empirical Investigation

The empirical period begins in 1988M1 and ends in 1997M6. We start from 1988, because our sample's small-open Asian countries liberalized their financial markets before 1988. The data plotted in Figure 1 include monthly observations of nominal interest rates, which is clearly described in Appendix 2. We include the United States (US) and Japan (JP), and 8 interesting Asian countries in our sample: Hong Kong (HK), Indonesia (IN), South Korea (KR), Malaysia (ML), the Philippines (PH), Singapore (SG), Thailand (TH), and Taiwan (TW).

The conventional ADF test is applied in order to examine the hypothesis of a unit root without a time trend for nominal interest rates. Considerable evidence exists that data-dependent methods for selecting the value of the lag order k in the ADF regressions are superior to choosing a fixed k a priori. We thus follow the recursive t-statistic procedure suggested by Campbell and Perron (1991) to determine the model's optimal lag order.[7]

In Table 1 we report the model's optimal lag order, the ADF test t-statistics, and the P-values of these statistics. The unit-root hypothesis is rejected at the 5% level of

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[7] Start with an upper bound, $k_{max} = 12$, on k. If the last included lag is significant, then choose k= k_{max}, otherwise reduce k by one until the last lag becomes significant. If no lags are significant, then set k=0. The 10% value of the asymptotic normal distribution, 1.645, is used to assess the significance of the last lag. Ng and Perron (1995) discuss the advantages of this method over an alternative procedure where k is chosen to minimize the Akaike Information Criterion.

significance if the P-value of the ADF test statistic is less than 0.05. Findings from Table 1 point out that the unit-root hypothesis of interest rates cannot be rejected for all countries in the panel at the 5% level by the ADF test. In addition, except for the Thailand-Japan interest differential, the conventional ADF test fails to reject the unit-root hypothesis of interest differentials at the 5% level of significance, regardless of whether the base country is the U.S. or Japan.

One possible reason for the ADF test's failure to reject the unit-root hypothesis may be due to the data's short time span (Shiller and Perron, 1985). In order to investigate this possibility, our paper applies the IPS test to re-examine the stationarity of the interest rates. We correct for possible serial correlation in residuals according to the value chosen by the single equation model in Table 1. To adjust for small sample bias and to allow for the presence of serial correlation and contemporaneous correlation in innovations, this paper also simulates the test's finite-sample critical values with a non-parametric Bootstrap.[8]

The empirical results of the IPS test are listed in Table 2. We first present the results with all countries in the panel. Findings from panel A of Table 2 provide no evidence of rejecting the unit-root hypothesis of nominal interest rates at the 5% level of significance, confirming that the nominal interest rates in the panel are nonstationary. Moreover, the findings from panel A of Table 2 point out that the stationarity of interest differentials is sensitive to the selected base countries. If Japan is used as the base country, then the unit-root hypothesis of the interest differentials cannot be rejected at the 5% level of significance. However, the unit-root hypothesis

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[8] We adopt the procedure of block-sampling the residuals and set the block size to be 12. However, our empirical findings are not significantly affected by using a different block size such as 6 and 18.

is indeed rejected at the conventional significance level when the U.S. serves as the base country. Evidence indicates that interest rates of the non-Japan Asian countries do not converge to the Japanese rate, but they appear to converge more to the U.S. rate.

Taylor and Sarno (1998) point out that panel unit-root tests may lead to a very high probability of rejecting the joint unit-root hypothesis when there is a single stationary process in a system of otherwise unit-root processes. To check the robustness of our finding from the IPS test, we exclude those stationary variables (based on the ADF test) from the panel. Since the unit-root hypothesis of Indonesia's interest rate is rejected at the 10% level of significance by the ADF test, we therefore exclude it from the panel and then apply the IPS test to re-examine the unit-root hypothesis of interest rates. As shown in panel B of Table 2, the P-value of the IPS statistic is 0.179, which fails to reject the unit-root hypothesis of interest rates at the conventional level.

We also exclude Thailand from the panel where Japan is the base country since the ADF test rejects the unit-root hypothesis of the Thailand-Japan rate at the 10% level. As shown from panel B of Table 2, the P-value of the IPS test is 0.654, indicating the non-convergence of the non-Japan Asian rates to the Japanese rate. As for the panel where U.S. is the base country, we exclude Hong Kong and Indonesia from the panel since Hong Kong-US and Indonesia-US interest differentials are stationary at the 10% level. Findings from panel B of Table 2 indicate the rejection of the unit-root hypothesis at the 10% level, which confirm that at least some of the interest rates of the non-Japan Asian countries converge to the U.S. rate.[9]

Taylor and Sarno (1998) construct the JLR test, which has the null hypothesis that there is at least one non-stationary series in a panel, as a complement to panel unit root tests. Therefore, a rejection of the null hypothesis with the JLR test implies that all series are stationary. The JLR statistic has a χ^2 distribution with one degree of freedom. The results of JLR test are reported in Table 3, which point out that the hypothesis of at least one non-stationary interest rate differential of the non-Japan Asian rates to the U.S. rate is rejected at the 10% level of significance.[10] By contrast, the hypothesis of at least one non-stationary interest rate differential of the non-Japan Asian rates to the Japanese rate fails to be rejected at the 10% level of significance. In sum, our empirical evidence strongly supports that the interest rates of the non-Japan Asian countries converge to the U.S. rate, rather than the Japanese rate.

To compare with the finding in Chinn and Frankel (1995), we also conduct the IPS test with a block size of eight,[11] and with the same sample period (1982Q3-1992Q1) as in Chinn and Frankel (1995) for the real interest rates of the eight

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[9] Moreover, we also find that interest rates for the four little dragons, Hong Kong, South Korea, Singapore, and Taiwan, converge relatively to the U.S. rate, rather than the Japanese rate.

[10] The lag length of the VAR system is selected according to the whiteness of the estimated residuals. We also correct for the JLR statistic's small sample bias based on the method proposed by Taylor and Sarno (1998).

[11] The empirical findings are not significantly affected when using a different block size such as 4 and 12.

countries commonly used by those two authors and us.[12] Note that the sample period in Chinn and Frankel (1995) contains an uneven deregulation degree of financial liberalization for the Asian countries under consideration. In addition, testing the stationarity of real interest differentials is equivalent to examining the joint hypothesis of interest-rate parity and PPP. Keeping the previous two points in mind, the findings here confirm that the real interest rates of the non-Japan Asian countries converge to the U.S. rate rather than Japan's.[13] These findings are also compatible with those in Chinn and Frankel (1995). However, if we change the sample period to the one used in this paper (1988Q1 to 1997Q2), then these real interest rates themselves become stationary and thus no longer qualify for any unit root test of interest differentials.[14] Nonetheless, these results show no evidence against our findings of the convergence of nominal interest rates.

Our finding that the interest rates of the non-Japan Asian countries converge relatively to the U.S. rate is indeed distinct. This is in contrast to Karfakis and Moschos (1990) and Katsimbris and Miller (1993), who fail to support the hypothesis of nominal interest rate convergence among industrial countries based on

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[12] The panel for the real interest rate includes Japan, the U.S., and six Asian countries, Hong Kong, South Korea, Malaysia, Singapore, Thailand and Taiwan. Indonesia is excluded due to some missing data in 1986.

[13] The $\bar{z}_{1NT}$ statistic for the interest differentials between Japan and the six countries is –0.891 with a P-value of 0.189, while the statistic for the interest differentials between the U.S. and the six countries is –2.697 with a P-value of 0.001. Furthermore, these results are unchanged after excluding the countries with a stationary interest rate differential.

[14] The detailed results are available from the authors upon request.

conventional ADF tests. Our results indicate that domestic monetary authorities of the non-Japan countries have directed a monetary policy to follow the U.S. policies rather than the Japanese ones. Taking into account the fact that the countries under study are all small relative to the U.S. and Japan, the above finding implies that U.S. monetary polices have more direct influences than those of Japan on these countries.

4. Conclusion

The convergence of interest rates reveals information about which major country's monetary policy the monetary authority of a small-open economy targets. Empirically, interest rate convergence can be examined by testing the stationarity of nominal interest differentials. The conventional finding based on the ADF test points out that interest rates and interest differentials are nonstationary for non-Japan Asian countries. This paper increases the data's span by jointly testing for a unit root across a large number of interest rates.

By employing panel unit-root tests, we investigate the convergence of non-Japan Asian countries' interest rates versus the rates in the U.S. and Japan. This paper carefully chooses its sample period in order to avoid a rejection that may be due to an uneven degree of financial liberalization of the countries under consideration. The empirical evidence that the interest rates of the non-Japan Asian countries converge relatively to the U.S. rate rather than Japan's in 1988-1997 indicates that domestic monetary authorities of the non-Japan Asian countries have directed a monetary policy to follow the U.S.'s policy rather than Japanese one in this period. The policy implication of this finding is that the U.S. may have a direct financial influence on these small-open countries before the currency crisis of 1997.

Appendix 1

This appendix provides a detailed description of the bootstrap procedure.

1. We obtain the bootstrap sample of the error term $\varepsilon_t^0 = [\varepsilon_{1t}^0, \varepsilon_{2t}^0, \cdots, \varepsilon_{Nt}^0]$ by estimating the following system equations using the iterative seemingly-unrelated regression (SUR) method:

$$\delta_{it} = \alpha_i + \beta_i \delta_{it-1} + \sum_{j=1}^{h_i} \gamma_{ij} \Delta\delta_{it-j} + \varepsilon_{it}^0, \quad i = 1, \ldots, N,$$

where δ_{it} has a unit root under the null hypothesis of $\beta_i = 1$.

2. The block resampling procedure as described in Berkowitz and Kilian (1996) is applied to generate residuals for simulation. That is, we divide $\varepsilon^0 = [\varepsilon_1^0, \ldots, \varepsilon_T^0]'$ into T-k overlapping blocks with length k+1 and randomly select a block for replacement, where $\varepsilon_j^0 = [\varepsilon_{1j}^0, \ldots, \varepsilon_{Nj}^0]$, j=1,…, T. We first generate a pseudo-random number from the U(0,1) distribution and then use it to generate a random number integer h that takes on the value 1,…, T-k with equal probability. Once h is generated, we draw a block of fitted residuals, $\bar{\varepsilon}_1 = [\varepsilon_h^0, \ldots, \varepsilon_{h+k}^0]'$, to obtain $\bar{\varepsilon}_1^*$. Repeating this operation m=T/(k+1) times yields a complete bootstrap sample of the error terms $\varepsilon^* = [\bar{\varepsilon}_1^*, \ldots, \bar{\varepsilon}_m^*]'$. The bootstrap sample $\delta_{i,t}^*$ for $\delta_{i,t}$ is generated as

$$\Delta\delta_{it}^* = \sum_{j=1}^{h_i} \hat{\gamma}_{ij} \Delta\delta_{it\text{-}j}^* + \varepsilon_{it}^*,$$

where $\Delta\delta_{it}^* = \delta_{it}^* - \delta_{it-1}^*$, and where the $\hat{\gamma}_{ij}$'s are the SUR estimates obtained from step 1. The initial values of δ_i^{*0} are obtained by block re-sampling. That is, we divide δ_{it} into T-k overlapping blocks and randomly select a block with a replacement for δ_i^{*0}.

3. Compute the t values of ψ_i for each group separately:

$$\Delta\delta_{it}^* = \upsilon_i + \psi_i \delta_{it-1}^* + \sum_{j=1}^{w_i} \varpi_{ij} \Delta\delta_{it-j}^* + \text{residual}.$$

4. Construct the t-bar statistic, W_{tbar}, based on (3).

5. Repeat steps 2 to 4 five thousand times to derive the empirical distribution of W_{tbar}.

Appendix 2

This appendix provides a table giving a summary description of the nominal interest rates used in the paper. Data, except for the Taiwanese rates, are obtained from either the International Financial Statistic (IFS) tape or the database constructed by Wharton Econometric Forecasting Associates (INTLINE), depending on the availability of data. Nominal interest rates for Taiwan are obtained from Financial Statistics Monthly, Taiwan District, R.O.C., (FSM).

Table A1. Nominal interest rates

Country Code	Country	Source	Description
HK	Hong Kong	INTLINE	Prime rate
JP	Japan	INTLINE	Eurocurrency rate -- 90 days
IN	Indonesia	IFS	Money market rate
KR	South Korea	INTLINE	Average yield on corporate bonds
ML	Malaysia	IFS	Money market rate
PH	Philippines	IFS	Treasury bill rate -- 91 days
SG	Singapore	INTLINE	Bank rate, fixed deposit, 3 months
TH	Thailand	INTLINE	Call money rate
TW	Taiwan	FSM	Money market rate on banker's acceptance -- 3 months
US	United States	INTLINE	Eurocurrency rate -- 3 months

Table 1. The ADF Test (1988M1-1997M6)

		HK	IN	JP	KR	ML	PH	SG	TH	TW	US
	k	1	8	4	12	10	11	2	12	8	7
i_j	τ_μ	-2.041	-2.653	-0.822	-1.655	-2.030	-0.906	-2.262	-2.220	-2.140	-2.447
	PV	0.271	0.076*	0.796	0.423	0.253	0.747	0.184	0.176	0.211	0.121
	k	3	8	---	10	3	8	2	0	8	1
$i_j - i_{JP}$	τ_μ	-1.036	-1.831	---	-2.146	-0.641	-1.768	-0.138	-3.057	-1.353	-0.583
	PV	0.730	0.344	---	0.208	0.848	0.374	0.937	0.034**	0.583	0.866
	k	6	8	1	12	4	8	7	12	8	---
$i_j - i_{US}$	τ_μ	-2.629	-2.740	-0.583	-2.171	-1.548	-1.026	-2.469	-2.411	-2.133	---
	PV	0.083*	0.062*	0.866	0.193	0.497	0.717	0.116	0.121	0.214	---

Note: 1. i_j is the nominal interest rate of country j, while $i_j - i_l$ (l = JP, US) is the interest rate differential of country j and l.

2. k is the lag order of the ADF test, which is selected based on the recursive procedure of Campbell and Perron (1991).

3. τ_μ is the t-statistic in the ADF test.

4. "PV" indicates the P-value of the ADF test.

5. ** and * denote the rejection of the unit-root hypothesis at the 5% and 10% level of significance, respectively.

Table 2. The Im-Pesaran-Shin Test (1988M1-1997M6)

	W_{tbar}	P-value	Critical Values 1%	5%	10%
A. IPS test results with the panel including all series					
i_j (all)	-1.631	0.113	-2.734	-2.105	-1.700
$i_j - i_{JP}$ (all)	-0.011	0.483	-2.701	-1.865	-1.429
$i_j - i_{US}$ (all)	-2.161	0.028**	-2.512	-1.849	-1.420
B. IPS test results with the panel excluding stationary series					
i_j (excluding IN)	-1.231	0.179	-2.657	-2.018	-1.662
$i_j - i_{JP}$ (excluding TH)	0.600	0.680	-2.799	-1.915	-1.482
$i_j - i_{US}$ (excluding HK and IN)	-1.378	0.085*	-2.254	-1.658	-1.278

Note: 1. $\bar{z}_{1NT}$ is the test statistic of Im, Pesaran and Shin (2003).

2. P-values are constructed based on the bootstrapped distribution.

3. The panel for the case of i_j(all) includes HK, IN, JP, KR, ML, PH, SG, TH, TW, and US, while the panel for the cases of $i_j - i_{JP}$ (all) and $i_j - i_{US}$ (all) both are HK, IN, KR, ML, PH, SG, TH, and TW.

4. The panel for the case of i_j (excluding IN) includes HK, JP, KR, ML, PH, SG, TH, TW, and US; for the case of $i_j - i_{JP}$ (excluding TH) the panel includes HK, IN, KR, ML, PH, SG, and TW; while the panel for the case of $i_j - i_{US}$(excluding HK and KR) includes IN, ML, PH, SG, TH, and TW.

5. ** and * denote the rejection of the unit-root hypothesis at the 5% and 10% level of significance, respectively.

Table 3. The JLR Test (1988M1-1997M6)

	JLR Statistic	Critical Values	
		5%	10%
$i_j - i_{JP}$ (all)	0.164	3.061	4.330
$i_j - i_{US}$ (all)	3.348*	3.061	4.338

Note: 1. The panel for the cases of $i_j - i_{JP}$ (all) and $i_j - i_{US}$ (all) both are HK, IN, KR, ML, PH, SG, TH, and TW.

2. * denotes the rejection of the null hypothesis of at least one non-stationary series at the 10% level of significance.

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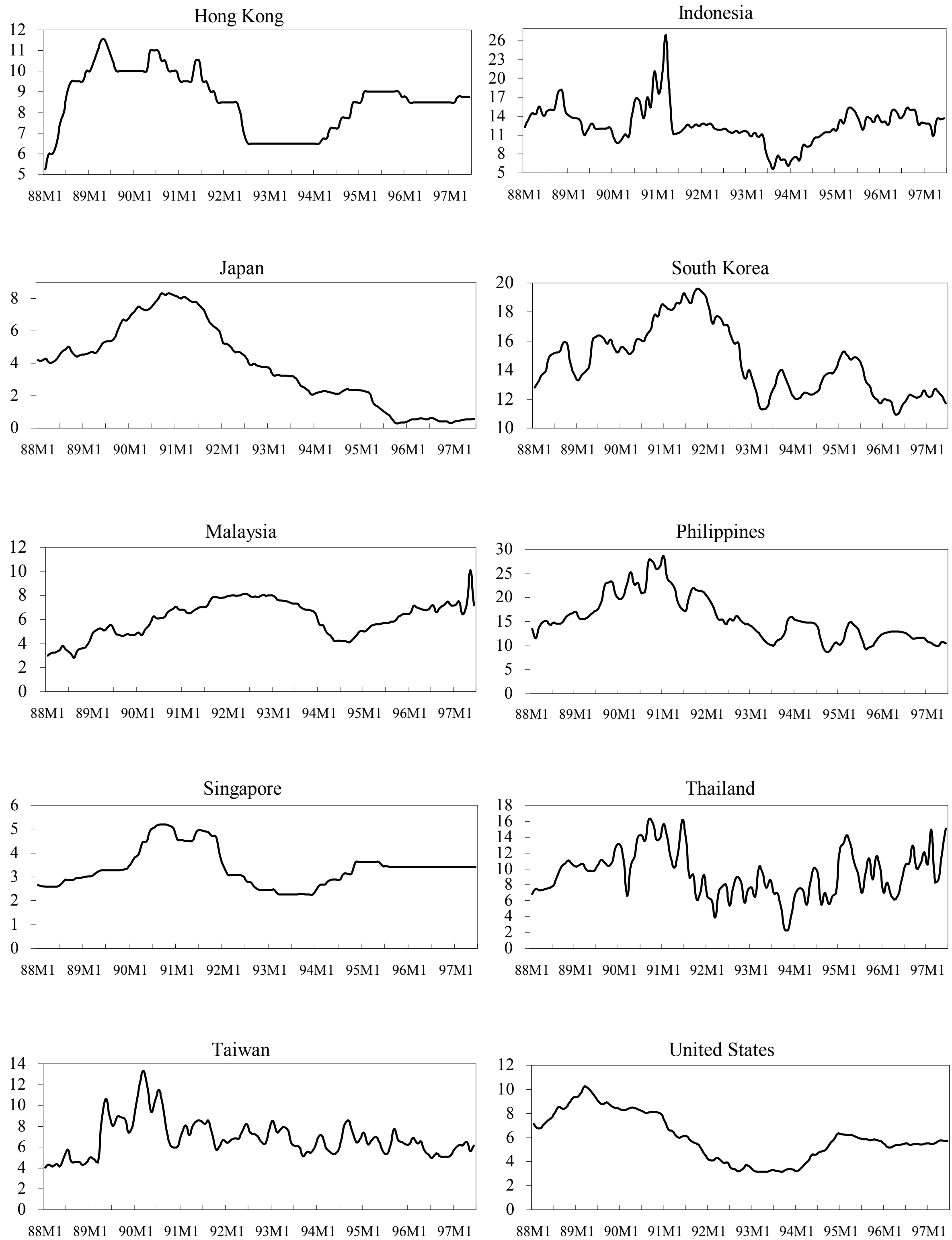

Figure 1. Nominal interest rates of ten countries (in %)