Hitotsubashi Journal of Economics 46 (2005), pp.1-27.

RECENT CHANGES IN LINKAGES OF ASIAN CURRENCIES WITH THE U.S. DOLLAR*

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Received July 2004; Accepted November 2004

Abstract

The paper analyzes whether twelve Asian countries have dollar-linked or yen-linked currencies, by applying Kalman Filter techniques to the Haldane and Hall (1991) and Frankel and Wei (1994) models. The empirical results show that the weights to the major currency of the region are stable and not low in the cases of Korea, Thailand, Singapore, Indonesia, and Philippines, in comparison with the case of Europe in the 1970s. Furthermore, when trade weighted exchange rates are used instead of yen/$ exchange rates, the East Asian currencies give higher weight to the currency basket than the dollar. In contrast to McKinnon (2000), the paper suggests that there is room to strengthen the regional currency cooperation on the basis of the yen or an East Asian currency basket.

Keywords: linkage, Kalman Filter technique, time varying parameter, trade weighted exchange rates
JEL Classification: F31, F33

I. *Introduction*

There has been much research on a yen bloc or a regional currency union, as a means of preventing excess volatility of Asian exchange rates after the 1997 currency crisis. This issue has gained greater interest in Korea and Japan since the circulation of euro cash in 2002. However, European currencies have slowly and continuously become linked to the Deutsche mark (DM) over the past thirty years, whereas Asian currencies have co-moved with the Japanese yen since the mid-1990s. Therefore, our arguments for a currency union first of all focus on whether Asian currencies have become synchronized with the Japanese yen to a greater or lesser degree. This topic must also be viewed from political and historical, as well as economic, aspects, because Asian countries have a strong antipathy against Japan for its past colonialism and the U.S. is against a regional currency union. Nevertheless, Asian countries

* I am grateful to the Forum on East Asian Financial and Monetary Cooperation for research support, and to an anonymous referee for insightful suggestions and comments.

have common economic incentives to organize a regional currency union. For example, like Germany, Japan wants to insulate its economy from dollar shock. Therefore, its exchange risk would be greatly reduced if the other Asian countries were to peg collectively to the yen instead of the dollar. The economies of Indonesia, Korea, Malaysia, Philippines, and Thailand suffered from the 1997 crisis and also need to prevent domestic currency from being exposed to excessive volatility through short-term, hot money flows and contagion or spillover effects. In a realistic aspect, as a currency union keeps member countries' currencies fixed at a common currency, it is important to scrutinize whether the countries which will potentially participate in a currency union have dollar-linked or yen-linked currencies. Therefore, this paper first examines to what extent Asian currencies have tracked the dollar or yen, using the methods of Haldane and Hall (1991) and Frankel and Wei (1994). Second, for comparative analysis it examines how closely the European currencies were synchronized with the DM from 1973 to 1998. Finally, it compares the European and Asian cases and explores the possibility of regional currency cooperation in East Asia.

The rest of the paper is organized as follows. In Section II, the Haldane and Hall (1991) model is estimated by Kalman Filter techniques to examine whether or not the yen is dominant in Asia. In Section III, in order to investigate the validity of the empirical results in Section II, the Kalman Filter technique is applied to the Frankel and Wei (1994) model under the assumption that the stochastic parameters are time varying. Section IV uses the Haldane and Hall Kalman Filter technique to consider the link of European currencies to the DM. Section V compares the European and Asian cases and discusses the possibility of currency cooperation in Asia. Section VI summarizes and concludes the paper.

II. *The Haldane and Hall (1991) Method*

As capital and financial markets have undergone continual liberation and opening in recent years, international financial markets are now more closely linked than before [e. g., Lin et al. (1994), Karolyi and Stulz (1995), Longin and Solnik (1995), Serletis and King (1997), Engle (2002), Manning (2002)]. For example, the co-movement of financial markets is generally estimated by cross-market correlation coefficients and the correlation analysis of stock returns indicates that the Korean stock returns as indicated by the KOSPI index move together with the Dow Jones and Nikkei 225 Indexes. In addition, the Korean won is more closely linked to the yen than before capital market liberalization or the currency crisis. Southeast Asian currencies also have similar patterns. In order to examine the link of these currencies to the yen, this paper uses the methods of Haldane and Hall (1991) and Frankel and Wei (1994).

First, the Haldane and Hall (1991) method is considered as follows:[1]

$$\ln S_{₩/\$,t} = \alpha_0 + \alpha_t \ln S_{¥/\$,t} + \varepsilon_t \tag{1}$$

$$\ln S_{₩/¥,t} = \beta_0 + \beta_t \ln S_{¥/\$,t} + \eta_t \tag{2}$$

where $S_{₩/\$,t}$ is the won-dollar exchange rate, $S_{₩/¥,t}$ the won-yen exchange rate, and $S_{¥/\$,t}$ the

[1] See also Hall et al. (1992), Caporale and Pittis (1993), Serletis and King (1997), and Manning (2002) for other applications.

yen-dollar exchange rate. α_0 and β_0 are stochastic constants and partial out all systematic influences on the won-dollar and won-yen exchange rates other than those resulting from movements in the yen-dollar exchange rate. The time varying parameters α_t and β_t measure the won's temporal relationship with the dollar and yen, respectively. For instance, when the won is perfectly pegged to the dollar, the won-dollar exchange rate is independent of the yen-dollar exchange rate, while the won-yen exchange rate is perfectly negatively correlated with the yen-dollar exchange rate. In this case, if the won-dollar and won-yen exchange rates are separately regressed against the yen-dollar exchange rate, the estimates of parameters are 0 and -1, respectively. On the other hand, when the won is perfectly linked to the yen, the won-yen exchange rate is independent of the yen-dollar exchange rate, while the won-dollar exchange rate is perfectly positively correlated with the yen-dollar exchange rate. In this case, the estimates of parameters are 1 and 0, respectively. If we estimate equations (1) and (2) by ordinary least squares, the problem of spurious regression arises. The co-movement of exchange rates is also time varying. Therefore, this study estimates the following equations, using the time varying parameter methodology proposed by Haldane and Hall (1991):

$$\ln S_{LC/\$,t} = \alpha_0 + \alpha_t \ln S_{¥/\$,t} + \varepsilon_t \tag{3}$$

$$\alpha_t = \phi_0 + \phi_1 \alpha_{t-1} + \upsilon_t \tag{4}$$

$$E(\varepsilon_t \varepsilon_\tau') = \begin{cases} R, & t=\tau \\ 0, & t\neq\tau \end{cases} \tag{5}$$

$$E(\upsilon_t \upsilon_\tau') = \begin{cases} Q, & t=\tau \\ 0, & t\neq\tau \end{cases} \tag{6}$$

$$E(\varepsilon_t \upsilon_\tau') = 0, \ t = 1, 2, \ldots, \tau = 1, 2, \ldots \tag{7}$$

where equations (3) and (4) are known as observation and state equations, respectively. LC represents the local currency and $S_{LC/\$}$ implies local currency-dollar exchange rates. The disturbances ε_t and υ_t are white noises and are assumed to be uncorrelated at all lags. The parameters α_0, ϕ_0, and ϕ_1 are estimated by the maximum likelihood estimation method under the normal distribution assumption [e. g., Hamilton (1994)].

Twelve Asian countries are considered in this study — Korea, Thailand, Taiwan, Singapore, India, China, Hong Kong, Malaysia, Indonesia, Philippines, Pakistan, and Sri Lanka. The daily exchange rate data were obtained from Pacific Exchange Rate Service in the University of British Columbia.[2] The sample period was from January 4, 1993 to December 31, 2003 for eight of the currencies, but from November 16, 1995 to December 31, 2003 in the cases of Indonesia, Philippines, and Pakistan, and from December 1, 1995 to December 31, 2003 for Sri Lanka. Figures 1 and 2 depict local currency-dollar exchange rates, with a base rate of 100 at the beginning of the sample period. After the Thai baht was floated on July 2, 1997, almost all Asian exchange rates began to depreciate steeply. Especially, the baht, rupiah, and won depreciated approximately 63, 81, and 64 percent, respectively, against the U.S. dollar between June and December 1997. Indonesian rupiah-dollar exchange rates are excluded for convenience's sake, because of problems with vertical scaling.

When equations (3) and (4) are estimated by the Kalman Filter method, the estimates of

[2] See http://fx.sauder.ubc.ca.

parameters ϕ_0 and ϕ_1 are almost 0 and 1, respectively. Therefore, as in other papers [Haldane and Hall (1991), Serletis and King (1997), Manning (2002)], the present study reports the estimation results under the following assumption that the stochastic parameter α_t follows a random walk:

$$\alpha_t = \alpha_{t-1} + \upsilon_t \tag{8}$$

Table 1 reports the estimation results for the whole sample period. As mentioned above, the sample period is shorter for Indonesia, Philippines, Pakistan, and Sri Lanka. As Malaysia declared a fixed exchange rate of 3.8 ringgits to the dollar in September 1998, the sample period considered in this case is from January 1993 to August 1998. In Tables 1 and 2, Ave. of α_t implies the average of α_t estimates over the sample period. As depicted in Figures 3 and 4, the estimates of parameters are time varying.

Figure 3 shows that the coefficient estimate of Korea ranged between 0.02 and 0.03 in the early 1990s, but began to increase after mid-1996 and peaked above 0.2 at the end of 1997 when the currency crisis occurred in Korea. The α_t coefficient persistently declined to 0.09 in the aftermath of the currency crisis, but again started to increase from the end of 2000. Rising up to 0.14, it then trended slightly downward again. In other words, the won was essentially a dollar-linked currency until the early 1990s, but began to co-move with the yen after the mid-1990s. The linkage between the won and yen increased dramatically during the 1997 currency crisis and then fell away slightly. Nevertheless, the won's association with the yen is significantly stronger than before the currency crisis. The average of α_t estimates over the sample period is 0.081, as shown in Table 1. The other crisis economies of Thailand, Malaysia, and Indonesia all had similar patterns to that of Korea, even though the absolute values of estimates were different. In short, the link between the crisis economies' local currencies and the yen increased markedly during the 1997 crisis and then declined moderately thereafter. Nevertheless, the estimates of α_t after the currency crisis were larger than before the currency crisis. All three countries assigned higher weight to the yen than Korea did for the whole period. The strength of Indonesian rupiah's relationship with the yen was weaker than that of the Thai baht and the Malaysian ringgit before the currency crisis, but was stronger after the currency crisis. Even in the cases of the noncrisis economies of Taiwan and Singapore, the co-movement of their currencies with the yen gained strength after the currency crisis, but not to the extent of the four crisis economies. Furthermore, the weight on the yen didn't increase markedly during the Asian crisis, as it had for the four crisis economies.[3]

Figure 4 shows that the strength of Philippines peso's relationship with the yen significantly increased during the Asian currency crisis and continued to strengthen throughout the sample period. The same is not true of the other countries in Figure 4 and the countries in Figure 3 experienced some decline in the yen link after the crisis. The Indian, Pakistani, and Sri Lankan rupees approximately tracked the dollar before the currency crisis and therefore the estimates of α_t were close to 0. However, after the crisis the estimates of α_t increased over 0.04 as the influence of the yen began to appear weakly. The Chinese renminbi and the Hong

[3] In the cases of weekly and monthly data, the estimates of α_t essentially have the similar pattern. However, the absolute values of the estimates are bigger. In the appendix, Figures A1 and A2 show estimates of α_t in the cases of the average of five daily exchange rates, while Figures A3 and A4 show estimates in the cases of the average of twenty daily exchange rates. McKinnon (2000) suggested that the low frequency data understate the degree to which the East dollar standard has been resurrected.

Kong dollar are both pegged to the dollar and the estimates of α_t were close to 0, although that of China jumped in January 1994 when the renminbi was devalued against the dollar.

It is well known that such estimation results can be distorted by large currency crises. As we are therefore more interested in recent exchange rate variations since the currency crisis of 1997, estimates were made with the same model for the sample period from January 4, 1999 to December 31, 2003. The empirical results are represented in Table 2 and Figures 5 and 6. First, as shown in Figure 5, the estimates of α_t for Korea, Thailand, and Singapore ranged in a similar fashion between 0.16 and 0.22. The average of α_t estimates was 0.192, 0.190, and 0.191 for Korea, Thailand, and Singapore, respectively (Table 2). The coefficient estimate of Korea had a similar pattern, but the absolute value was larger, compared with that in Figure 3 for the whole sample period. In Figure 3, the estimate of α_t ranged between 0.10 and 0.15 during the sub-sample period from 1999 to 2003, whereas in Figure 5 it ranged between 0.17 and 0.22 during the same period. This result indicates that Korea gave greater weight to the yen in the latter case. However, the result was reversed for Thailand and Singapore, in which the estimates of α_t in Figure 5 were smaller than those in Figure 3 and had a very similar pattern to that of Korea. The case of Indonesian rupiah did not differ from these two currencies. The Indonesian coefficient estimate fluctuated between 0.30 and 0.40 during the sub-period from 1999 to 2003, when the model was estimated over the whole sample period, but ranged between 0.06 and 0.118 when only the sub-sample period from 1999 to 2003 was considered, which was smaller than even that for Korea, Thailand, and Singapore. The currencies of these four countries were more strongly linked to the yen in 2001 than in 1999 or 2003, which implies that the linkage between East Asian local currencies and the yen has not strengthened continuously. The case for Taiwan was similar. The estimates of α_t for Philippines, India, and Pakistan were smaller in the sub-period of Figure 6 than those in the whole period of Figure 4. However, for the Sri Lankan rupee, there was no difference between Figures 4 and 6.

III. *An Application of the Frankel and Wei (1994) Method*

As already mentioned above, the estimates of α_t in Figures 5 and 6 show the same pattern as those in Figures 3 and 4, but the absolute values of the estimates are different. In order to check the empirical results again, the Frankel and Wei (1994) method was also used for the sub-sample period from January 1999 to December 2003. Before Frankel and Wei (1993, 1994), Giavazzi and Giovannini (1990) regressed local currency-DM exchange rate changes on local currency-dollar exchange rate changes for European countries and found that the link of Exchange Rate Mechanism (ERM) bilateral rates to the dollar has declined rapidly since the European Monetary System (EMS) was established. McKinnon (2000), Ogawa (2001), and Baig (2001) have subsequently used the Frankel and Wei (1994) method with ordinary least squares. However, the present paper applies the Kalman Filter method to the Frankel and Wei (1994) model under the assumption that the stochastic parameters are time varying. This leads to the following model:

$$e_{LC/SF,t} = \gamma_0 + \gamma_t' \, e_t + \zeta_t \tag{9}$$

$$\gamma_t = \lambda_0 + \lambda_1 \gamma_{t-1} + \omega_t \tag{10}$$

$$E(\zeta_t \zeta_\tau') = \begin{cases} \Gamma, & t=\tau \\ 0, & t \neq \tau \end{cases} \tag{11}$$

$$E(\omega_t \omega_\tau') = \begin{cases} \Psi, & t=\tau \\ 0, & t \neq \tau \end{cases} \tag{12}$$

$$E(\zeta_t \omega_{i,\tau}') = 0, \qquad t = 1, 2, \ldots, \tau = 1, 2, \ldots \tag{13}$$

$$\gamma_t = \{\gamma_{\$/SF,t}, \gamma_{¥/SF,t}, \gamma_{EU/SF,t}\}'$$

$$e_t = \{e_{\$/SF,t}, e_{¥/SF,t}, e_{EU/SF,t}\}'$$

$$\lambda_0 = \{\lambda_{\$/SF,0}, \lambda_{¥/SF,0}, \lambda_{EU/SF,0}\}'$$

$$\omega_t = \{\omega_{\$/SF,t}, \omega_{¥/SF,t}, \omega_{EU/SF,t}\}'$$

where LC is the local currency, SF is the Swiss franc, and EU is the euro. $S_{LC/SF,t}$ implies local currency-Swiss franc exchange rates and $e_{LC/SF,t}$ is calculated as $100\times(\ln S_{LC/SF,t} - \ln S_{LC/SF,t-1})$. λ_1 in equation (10) is the 3×3 diagonal matrix which is composed of $\lambda_{\$/SF,1}$, $\lambda_{¥/SF,1}$, and $\lambda_{EU/SF,1}$. In the same way, Ψ in equation (12) is the 3×3 diagonal matrix which is composed of $\Psi_{\$/SF}$, $\Psi_{¥/SF}$, and $\Psi_{EU/SF}$. The disturbances ζ_t, $\omega_{\$/SF,t}$, $\omega_{¥/SF,t}$, and $\omega_{EU/SF,t}$ are white noises and they are assumed to be uncorrelated at all lags. In earlier research on Asian currencies, the Swiss franc and SDR were used as numeraires to measure the value of currencies [e. g., Frankel (1992), Frankel and Wei (1993, 1994), McKinnon (2000), Baig (2001)], with the Swiss franc being preferred because it was known to be comparatively independent of other currencies.[4] Therefore, this paper also selects the Swiss franc as numeraire. If the local currency is fixed against the dollar, then $\gamma_{\$/SF,t}$ should be close to 1, while $\gamma_{¥/SF,t}$ and $\gamma_{EU/SF,t}$ are close to 0. If it is pegged to the yen, then $\gamma_{¥/SF,t}$ should be close to 1 and the others close to 0. The maximum likelihood estimation method was used to estimate γ_0, $\lambda_{\$/SF,0}$, $\lambda_{¥/SF,0}$, $\lambda_{EU/SF,0}$, $\lambda_{\$/SF,1}$, $\lambda_{¥/SF,1}$, $\lambda_{EU/SF,1}$, Γ, $\Psi_{\$/SF}$, $\Psi_{¥/SF}$, and $\Psi_{EU/SF}$ over the sample period from January 4, 1999 to December 31, 2003.

Table 3 presents summary statistics for local currency-Swiss franc exchange rate changes, $e_{LC/SF,t}$. During the sample period, all local currencies depreciated against the Swiss franc, but the depreciation was statistically significant at the 10 percent level only in the case of Sri Lanka. The standard deviation indicates that the rupiah was most volatile. In the cases of Indonesia, Philippines, and Pakistan, the skewness was negative. Kurtosis was greater than 3 (kurtosis of normal distribution) in all cases, and particularly was greater than 10 for EU, Indonesia, Philippines, and Pakistan. The maximum and minimum changes in this sample were relatively big for Indonesia, Philippines, and Pakistan. Q(10) indicates Ljung-Box statistics for 10^{th} order correlation in $e_{LC/SF,t}$. The test results show that local currency-Swiss franc exchange rate changes were not serially correlated, with the exceptions of Indonesia and Pakistan.

First, local currency-Swiss franc exchange rate changes were regressed on dollar-Swiss franc and yen-Swiss franc exchange rates, in order to compare the estimation results from the two methods. The estimation results are reported in Table 4. Ave. of $\gamma_{\$/SF,t}$ ($\gamma_{¥/SF,t}$) represents

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[4] However, as mentioned below, the Swiss franc had co-moved with the DM from the 1970s. Therefore, it is problematic to use the Swiss franc as an independent currency in the case of the DM. Using the SDR as numeraire also has the shortcoming that the SDR is itself a basket of five major currencies.

the average of $\gamma_{\$/SF,t}$ ($\gamma_{¥/SF,t}$) over the sample period. Figure 7 also shows to what extent six East Asian local currencies have tracked the yen. The other six local currencies were not considered because they were pegged strongly to the dollar and empirical results were not different from those from the Haldane and Hall (1991) method. For Korea, the estimate of $\gamma_{¥/SF,t}$ was not basically different from the estimate of α_t in the Haldane and Hall (1991) method. The estimate of $\gamma_{¥/SF,t}$ ranged between 0.16 and 0.22 and was bigger in the mid-sample period than in the beginning or end of the sample period. The coefficient estimates of Thailand and Singapore were similar to that of Korea. However, in Thailand the estimate of $\gamma_{¥/SF,t}$ was constant rather than time varying, which was different from the estimate of α_t in the Haldane and Hall (1991) method. In the cases of Indonesia and Taiwan, although the estimation results were indistinguishable between the two methods, the estimates of $\gamma_{¥/SF,t}$ obtained from the Frankel and Wei (1994) method changed more rapidly because exchange rate changes were used instead of level. The estimate of $\gamma_{¥/SF,t}$ seem to be constant for Philippines. Table 4 reports that the averages of $\gamma_{¥/SF,t}$ estimates were very close to 0 for Malaysia, China, Hong Kong, India, and Sri Lanka. Figure 8 depicts the dollar coefficients. All six countries had estimates of $\gamma_{\$/SF,t}$ above 0.7 on average, which suggests that Asian currencies are linked closer to the dollar than to the yen. In the case of Philippines, the estimate of $\gamma_{\$/SF,t}$ was constant rather than time varying. The other six countries had $\gamma_{\$/SF,t}$ coefficients close to 1.

Table 5 reports the empirical results of regressing local currency-Swiss franc exchange rate changes on three variables — dollar-Swiss franc, yen-Swiss franc, and euro-Swiss franc exchange rate changes. The movements of $\gamma_{¥/SF,t}$ coefficients are described in Figure 9. In the case of Korea, the estimate of $\gamma_{¥/SF,t}$ was more volatile than that in the two-variable model shown in Figure 7. The estimate of $\gamma_{¥/SF,t}$ was fixed at approximately 0.1 in the beginning of the sample period, then increased to 0.5 in the first half of 2001 and then decreased to range between 0.2 and 0.3. For Thailand, the estimate of $\gamma_{¥/SF,t}$ was almost constant at around 0.18 in Figure 7. In the case of Singapore, the estimate of $\gamma_{¥/SF,t}$ was greater after 2001 than in the beginning of the sample period. Indonesia, Taiwan, and Philippines had almost the same $\gamma_{¥/SF,t}$ coefficients as those in the two-variable model. Figure 10 describes each Asian currency's relationship with the dollar. Although the estimates of $\gamma_{\$/SF,t}$ were similar to those in Figure 8, the Korean $\gamma_{\$/SF,t}$ coefficient was more volatile and smaller in the end of the sample period, compared with the two-variable model, and the Singaporean coefficient was also more volatile. Figure 11 depicts how Asian currencies responded to the euro. The estimates of $\gamma_{EU/SF,t}$ were time varying only for Singapore, while they seemed to be constant for the other five countries.

McKinnon (2000) and McKinnon and Schnabl (2004a) similarly used the original Frankel and Wei (1994) model for nine Asian exchange rates. They estimated equation (9) by OLS for three periods: pre-crisis, crisis, and post-crisis. They didn't employ the Kalman Filter estimation method. But the estimation results of linear models are often not credible, because they are easily distorted by the existence of outliers. For example, the linkage between the won and yen increased dramatically during the 1997 currency crisis in this paper. On the contrary, McKinnon (2000) and McKinnon and Schnabl (2004a) showed that the Korean won was pegged more strongly to the dollar during the currency crisis. The results are exceptional, even if they are compared with their results for the other Asian countries. Their estimation results for the other Asian countries are not largely different from those of this paper. The crisis period may be arbitrarily chosen. Furthermore, as already shown above, estimates of parameters are not constant, but time varying. In order to confirm this, McKinnon (2000) and

McKinnon and Schnabl (2004a, 2004b) employed 30- and 130-trading day rolling regressions, respectively. In these cases, however, the results were also arbitrary because they depended on how many trading days were chosen. As the number of trading days included in the rolling regressions was increased, the parameter estimates become smoother. Their estimation results showed that 30- and 130-trading day rolling coefficient estimates for dollar-Swiss franc exchange rate changes were different. The results in the present paper are relatively similar to those in the 130-trading day rolling regressions. However, the 130-trading day rolling coefficient estimates were often greater than 1 and were more volatile than those in this paper. In their studies, DM-Swiss franc exchange rate changes are included as one of the independent variables, despite the fact that the Swiss franc has co-moved with the DM since the 1970s.

In short, under the assumption that the stochastic parameters are time varying, the empirical results obtained from the Haldane and Hall (1991) and Frankel and Wei (1994) methods are summarized as follows. The strength of the relationship of the Asian currencies with the yen was greater than that before the Asian currency crisis, with the exceptions of China and Hong Kong, and of Malaysia since 1998, whose currencies are fixed against the dollar. Particularly, in the crisis economies of Korea, Thailand, Indonesia, and Malaysia, the synchronization of their currencies with the yen strengthened more dramatically during the crisis period. When only the post-crisis period from 1999 to 2003 is considered, each currency's association with the yen was stronger in the mid-sample period than in the beginning or end of the sample period. Nevertheless, the Asian currencies remained linked more closely to the dollar than the yen, even after the currency crisis. The euro coefficients were relatively smaller and close to 0.

IV. *Comparison with Europe*

Following the above analysis of how closely the Asian currencies have been linked to the dollar or yen since the early 1990s, using the Haldane and Hall (1991) and Frankel and Wei (1994) methods, this section applies the Haldane and Hall (1991) method to the daily exchange rates of 14 European currencies and then compares these results with those for Asian countries. The Haldane and Hall (1991) method is chosen instead of the Frankel and Wei (1994) method, because the focus is on the DM-dollar exchange rates and the Swiss franc and SDR have limitations as numeraires, as already described above. The sample period from January 2, 1973 to December 31, 1998 was selected, corresponding to the floating of the main currencies in March 1973 and the debut of the euro on January 1, 1999. Fourteen European countries — France, Italy, the United Kingdom, the Netherlands, Belgium, Norway, Sweden, Switzerland, Austria, Denmark, Finland, Ireland, Portugal, and Spain — are analyzed. The data were obtained from the Federal Reserve Board database. Figures 12 and 13 show the daily exchange rates of the 14 European countries, plus Germany, converted to a base of 100 at the start of the period. The currencies were certainly less linked to the DM in the 1970s than in the 1980s. The linkage between the local currencies and the DM also temporarily trended downward in the pre- and post-1993 periods when the Maastricht Treaty became effective. The DM has been stable since the mid-1980s.

All assumptions are the same as those in the former section. DM-dollar exchange rates are used instead of yen-dollar exchange rates in equation (3). The estimation results are reported in Table 6. In Figures 14 and 15, as well as in Table 6, it is not easy to uniformly judge whether the European currencies were linked to the dollar or DM for the sample period from 1973 to 1979 in which the EMS was established. The currencies of the countries neighboring Germany, i.e., Netherlands, Switzerland, and Austria were more closely linked to the DM than the dollar from the early 1970s. In Figures 14 and 15, when the coefficient estimates are greater than 1 or less than 0, they are assumed to be equal to 1 or 0, respectively. On the other hand, the Italian lira, Swedish krona, Spanish peseta, and Portuguese escudo had close relationships with the dollar for the same period. For France and the United Kingdom, the estimates of α_t were less than 0.5 and therefore were linked more closely to the dollar. Nevertheless, it is clear from Figures 14 and 15 that the European currencies began to rapidly co-move with the DM since the beginning of the 1980s. After the mid-1980s, the British pound, Italian lira, Swedish krona, Spanish peseta, and Finnish markka were less linked to the DM than before. Especially, after West Germany and East Germany were united in 1990, the DM appreciated because of fiscal expansion combined with tight monetary policy. This forced the other European currencies to depreciate rapidly with the consequence that the U.K. and Italy resigned from the ERM. From this point, these currencies were fast linked to the DM. There was a speculative attack against the French franc in mid-1993. As a result, the EMS officials allowed a greater fluctuation of exchange rates among member countries from ± 2.25 percent to ± 15 percent. Figure 14 depicts that the French franc's association with the DM declined temporarily in mid-1993. Nevertheless, the European currencies became very closely pegged to the DM after the Maastricht Treaty came into effect in November 1993. In the cases of Sweden, Norway, Ireland, and especially the U.K., the estimates of α_t suddenly decreased from the mid-1990s.[5]

V. *A Policy Implication*

The above section examined the changing strength of the relationship of fourteen European currencies with the DM from 1973 to 1998. In contrast with Europe, Asian countries have never experienced any regional currency union. Therefore, if we want to test the currency-bloc hypothesis in two regions, it is first of all necessary to compare Asia in recent years with Europe in the 1970s. In the case of Europe before 1980, the estimates of α_t were not similar among the fourteen countries. Switzerland and Austria had α_t coefficients greater than 0.90, while Sweden and Spain had α_t coefficients less than 0.10. The estimates of α_t were less than 0.45 in the same period for France, the U.K., and Italy. In Table 6, Ave. of $\alpha_t^{\blacklozenge}$ represents the average of α_t estimates for the sample period from January 2, 1973 to February 28, 1979. During this period, the averages of α_t estimates were also dissimilar among the European countries.

On the other hand, for Korea, Thailand, and Singapore in Asia, the estimates of α_t moved slightly between 0.18 and 0.20. Indonesia and Philippines also had stable α_t coefficients which were 0.12 on average. The weight on the yen, the major Asian currency was stable and not low in these five Asian countries, compared with Europe in the 1970s. However, China, Hong

[5] Even if the low frequency data are used, the results are not changed. In the appendix, Figures A5 and A6 show estimates of α_t in the cases of the average of five daily exchange rates. Similarly, Figures A7 and A8 plot estimates of α_t in the cases of the average of twenty daily exchange rates.

Kong, and Malaysia had α_t coefficients very close to 0 because their currencies have been perfectly pegged to the dollar. Particularly in the case of Malaysia, its currency was fixed against the dollar on the way to its synchronization with the yen. Therefore, if Malaysia abandons its fixed exchange rate, the Malaysian ringgit seems likely to be fast linked to the yen from the viewpoint of economic fundamentals. However, there are political and historical constraints on the yen becoming a common or central currency in the Asian region. As already mentioned above, Asian currencies remain more closely linked to the dollar than the yen.

In these respects, we also need to consider the currency basket, which is derived from multiplying each Asian currency by its total trade weight or GDP weight in the Asian region. For example, it is assumed that the weighted exchange rates (WER) are calculated as follows:

$$\mathrm{WER}_t = \sum_{i=1}^{13} W_{i,t}(100 \times \frac{LC_{i,t}}{\$_t} / \frac{LC_{i,1}}{\$_1}) \tag{14}$$

$$W_i = \kappa_i / \sum_{i=1}^{13} \kappa_i \tag{15}$$

where κ_i implies each Asian country's total trade weight in the U.S., which is obtained from the FRB database. The total trade weights of Pakistan and Sri Lanka in the U.S. are small and are therefore not described in the FRB database. So κ_i is assumed to be 0 for these two countries.

First, the paper analyzes whether or not each Asian currency closely links itself to the trade-weighted currency, instead of the yen. The estimation results from the Haldane and Hall (1991) method are reported in Figure 16. As shown in Figure 16, Japan and Indonesia had α_t coefficients greater than 1.[6] Korea, Thailand, Singapore, and Philippines had estimates of α_t greater than 0.5. The empirical results suggest that these countries' currencies were more closely linked to the trade-weighted currency (WC) than to the dollar and that the movements of α_t coefficients were stable. In the case of Taiwan, the estimate of α_t was bigger than 0.2, but was about 0.04 when yen-dollar exchange rates were used.

Figures 17 and 18 show the regression of local currency-Swiss franc exchange rates on dollar-Swiss franc and weighted currency-Swiss franc exchange rates. Figure 17 plots $\lambda_{WC/SF}$ weighted currency coefficients. Japan and Indonesia had $\lambda_{WC/SF}$ coefficients greater than 1, and Thailand and Singapore greater than 0.6. The Korean estimate of $\lambda_{WC/SF}$ was in the neighborhood of 1. The empirical results were not different from those obtained from the Haldane and Hall (1991) method. Figure 18 describes the movements of the dollar $\lambda_{\$/SF}$ coefficients. For Japan, Indonesia, Korea, Thailand, and Singapore, the estimates of $\lambda_{\$/SF}$ were less than 0.5 and therefore their currencies were more closely linked to the weighted currency than the dollar.

In summary, for a currency basket derived from multiplying each Asian currency by its total trade weight in the U.S., the estimation results from the two methods suggest that the currencies of Japan, Indonesia, Korea, Thailand, Singapore, and Philippines had closer relations with this weighted currency than with the dollar, and that their relations were very stable. Furthermore, from the viewpoint of the past European and Malaysian experiences, and based on the present political and economic situations, if China, Hong Kong, and Malaysia

[6] In the figures plotted to now, coefficient estimates greater than 1 or less than 0 are assumed to be equal to 1 or 0, respectively.

abandon their fixed exchange rates, their currencies will become largely linked to the yen or the basket currency. As they are economically and geographically independent of East Asia, the currencies of India, Pakistan, and Sri Lanka are almost unlinked to the yen. Hence, they don't seem to participate in any regional currency cooperation based on the major currency of Asia in the near future. Their currencies, however, have started to co-move with the yen recently. This paper has focused on daily data. However, when weekly or monthly data are used, the strength of the relationship of the East Asian currencies with the yen is significantly increased. To conclude the empirical results here, in contrast to McKinnon (2002), suggests that regional currency cooperation on the basis of the yen or a currency basket can be expanded, at least in East Asia.[7]

VI. *Conclusions*

This paper analyzes the closeness of the relationships between twelve Asian currencies and the major currencies — dollar, yen, and euro — by applying Kalman Filter methods to the Haldane and Hall (1991) and Frankel and Wei (1994) models. It also compares the empirical results with those for fourteen European currencies and examines the possibility of regional currency cooperation in Asia.

The estimation results from the two methods indicate that the currencies of most of the Asian countries, except for China, Hong Kong, and Malaysia, are more closely linked to the yen after the Asian currency crisis than they had been before the crisis. Particularly, for the crisis economies of Korea, Thailand, Indonesia, and Malaysia, the strength of each currency's relationship with the yen increased rapidly during the crisis. For the sample period from January 1999 to December 2003, their synchronization with the yen was slightly stronger in the mid-sample period than in the beginning or end of the sample period. Furthermore, when trade weighted exchange rates were used instead of yen/$ exchange rates, the East Asian currencies gave higher weight to the currency basket than the dollar.

As for the European currencies, their synchronization with the DM was volatile and there was little apparent similarity among the member countries before 1979 when the EMS was established. On the other hand, in the cases of Korea, Thailand, Singapore, and Indonesia, even if the weights to the major currency of the region were weaker than those of France, the U.K., and Italy, their movements were more stable. In addition, from the viewpoint of the past European and Malaysian experiences, and on the basis of the present political and economic situations, if China, Hong Kong, and Malaysia allow their exchange rates to move freely, their currencies will become more closely linked to the yen or a currency basket based on the trade and GDP weights in the Asian region. India and Pakistan haven't had relatively close relations with East Asia until now. Nevertheless, their currencies have started to become weakly linked to the yen in recent years. In this paper, daily data were used, but the link of East Asian currencies to the yen is strengthened when low-frequency data are chosen. These findings support the promotion of some form of regional currency cooperation based on the yen or an

[7] Baek and Song (2001) suggested that Japan, Korea, Thailand, Singapore, Indonesia, Taiwan, Malaysia, Hong Kong, and China are plausible candidates, from the viewpoint of similarities of intraregional trade, export structures, and openness.

East Asian currency basket.

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TABLE 1. THE HALDANE AND HALL (1991) ESTIMATION METHOD (Whole Sample Period)

	α_0	Q	R	ln L	Ave. of α_t
Korea	6.549**	exp(−12.303)**	exp(−25.504)	8824.916	0.081
Thailand	2.488**	exp(−12.889)**	exp(−24.663)	9639.889	0.216
Taiwan	3.097**	exp(−14.891)**	exp(−15.693)	12364.202	0.066
Singapore	−0.584**	exp(−14.440)**	exp(−35.081)	11792.300	0.225
India	3.606**	exp(−14.689)**	exp(−14.034)	11925.380	0.013
China	1.996**	exp(−12.820)**	exp(−18.636)	9543.104	0.019
Hong Kong	2.032**	exp(−19.763)**	exp(−17.022)	18110.853	0.004
Malaysia	0.011	exp(−12.861)**	exp(−23.577)	4993.595	0.211
Indonesia	7.407**	exp(−10.671)**	exp(−26.039)	4799.552	0.285
Philippines	2.905**	exp(−13.005)**	exp(−26.213)	7178.269	0.167
Pakistan	3.848**	exp(−13.612)**	exp(−11.373)	7240.159	0.014
Sri Lanka	4.372**	exp(−14.434)**	exp(−31.376)	5396.520	0.015

Notes: 1) ln L implies the maximum value of log likelihood function.
2) ** denotes significant at the 1% level.
3) Ave. of α_t implies the average of α_t estimates over the sample period.

TABLE 2. THE HALDANE AND HALL (1991) ESTIMATION METHOD (1999.1.4–2003.12.31)

	α_0	Q	R	ln L	Ave. of α_t
Korea	6.184**	exp(−14.064)**	exp(−13.060)**	4946.507	0.192
Thailand	2.821**	exp(−14.302)**	exp(−15.364)	5206.280	0.190
Taiwan	3.296**	exp(−15.442)**	exp(−14.898)	5852.747	0.044
Singapore	−0.348**	exp(−15.164)**	exp(−21.335)	5764.819	0.191
India	3.834**	exp(−16.564)**	exp(−14.450)**	6327.991	−0.000
China	2.113**	exp(−26.237)	exp(−16.794)**	8712.740	0.001
Hong Kong	2.040**	exp(−19.769)**	exp(−17.709)**	8348.112	0.003
Malaysia	1.336**	exp(−26.927)	exp(−15.830)**	8123.413	0.001
Indonesia	8.512**	exp(−11.673)**	exp(−17.868)	3577.407	0.120
Philippines	3.311**	exp(−13.594)**	exp(−12.321)**	4618.418	0.117
Pakistan	3.997**	exp(−14.529)**	exp(−11.540)**	4808.121	0.009
Sri Lanka	4.373**	exp(−14.421)**	exp(−27.908)	5299.573	0.016

Notes: 1) ln L implies the maximum value of log likelihood function.
2) ** denotes significant at the 1% level.
3) Ave. of α_t implies the average of α_t estimates over the sample period.

TABLE 3. SUMMARY STATISTICS FOR $e_{LC/SF,t}$

	Mean	Stand. Dev.	Skewness	Kurtosis	Maximum	Minimum	Q(10)
U.S.	0.008	0.666	0.095	4.010	2.728	−2.897	6.871
Japan	0.004	0.787	0.039	4.567	3.652	−3.296	3.490
EU	0.003	0.239	0.656	11.305	2.141	−1.402	9.947
Korea	0.008	0.812	0.224	3.756	3.178	−2.514	12.139
Thailand	0.015	0.725	0.095	4.668	3.897	−3.682	11.019
Taiwan	0.012	0.679	0.098	3.858	2.789	−2.269	11.175
Singapore	0.010	0.647	0.153	4.003	2.846	−2.232	9.624
India	0.013	0.685	0.139	4.096	2.887	−2.790	7.437
China	0.008	0.665	0.091	4.011	2.713	−2.889	6.735
Hong Kong	0.008	0.666	0.097	4.036	2.727	−2.897	6.395
Malaysia	0.008	0.669	0.099	3.971	2.720	−2.910	7.451
Indonesia	0.014	1.543	−0.207	10.901	9.844	−10.404	37.323**
Philippines	0.037	0.894	−2.176	37.325	4.569	−12.640	12.442
Pakistan	0.017	0.864	−0.186	18.420	8.330	−8.551	29.681**
Sri Lanka	0.035+	0.745	0.279	6.019	4.865	−3.592	8.704

Notes: 1) Q(10) implies Ljung−box statistics for 10th order correlation in $e_{LC/SF,t}$.
2) + and ** denote significant at the 10% and 1% levels, respectively.

TABLE 4. THE FRANKEL AND WEI (1994) METHOD: KALMAN FILTER ESTIMATION (1999.1.4–2003.12.31)

	Korea	Thailand	Taiwan	Singapore	India	China
γ_0	−0.001	0.006	0.006	0.001	0.003	−0.000
$\lambda_{\$/SF,0}$	0.314+	0.382	0.109	0.582**	1.317**	0.940**
$\lambda_{¥/SF,0}$	0.100**	0.178**	0.035**	0.057*	−0.002	−0.000
$\lambda_{\$/SF,1}$	0.628**	0.516+	0.884**	0.237	−0.317*	0.058
$\lambda_{¥/SF,1}$	0.477**	0.011	−0.081	0.685**	0.046	−0.015
$\Psi_{\$/SF}$	exp(−3.722)	exp(−3.911)	exp(−6.749)	exp(−4.684)*	exp(−4.520)**	exp(−24.097)
$\Psi_{¥/SF}$	exp(−5.163)	exp(−4.269)+	exp(−4.256)**	exp(−6.114)	exp(−20.009)	exp(−8.411)**
Γ	exp(−1.635)**	exp(−2.124)**	exp(−3.302)**	exp(−2.998)**	exp(−3.897)**	exp(−7.169)**
Ln L	−814.131	−535.652	153.576	25.832	543.841	2626.059
Ave. of $\gamma_{\$/SF,t}$	0.844	0.790	0.946	0.762	1.000	0.998
Ave. of $\gamma_{¥/SF,t}$	0.192	0.180	0.032	0.180	−0.002	−0.000

	Hong Kong	Malaysia	Indonesia	Philippines	Pakistan	Sri Lanka
γ_0	0.000	−0.000	−0.000	0.031	0.004	0.023*
$\lambda_{\$/SF,0}$	1.037**	0.989**	1.781**	0.883**	1.219+	0.559**
$\lambda_{¥/SF,0}$	0.002	−0.002	0.090	0.021	0.008	−0.010
$\lambda_{\$/SF,1}$	−0.040	0.013	−0.886**	0.033	−0.223	0.436**
$\lambda_{¥/SF,1}$	0.009	−0.003	0.105	0.752**	0.380**	0.037
$\Psi_{\$/SF}$	exp(−25.584)	exp(−22.815)	exp(−3.967)	exp(−18.767)	exp(−15.901)	exp(−2.864)**
$\Psi_{¥/SF}$	exp(−24.537)	exp(−22.453)	exp(−1.057)**	exp(−31.173)	exp(−2.045)**	exp(−23.784)
Γ	exp(−6.931)**	exp(−5.918)**	exp(0.496)**	exp(−1.020)**	exp(−1.527)**	exp(−2.416)**
Ln L	2564.312	1929.670	−2168.420	−1139.910	−989.946	−410.506
Ave. of $\gamma_{\$/SF,t}$	0.997	1.002	0.944	0.914	0.997	0.992
Ave. of $\gamma_{¥/SF,t}$	0.002	−0.002	0.101	0.088	0.014	−0.010

Notes: 1) ln L implies the maximum value of log likelihood function.
2) +, *, and ** denote significant at the 10%, 5%, and 1% levels, respectively.
3) Ave. of $\gamma_{\$/SF,t}$ ($\gamma_{¥/SF,t}$) implies the average of $\gamma_{\$/SF,t}$ ($\gamma_{¥/SF,t}$) estimates over the sample period.

TABLE 5. THE FRANKEL AND WEI (1994) METHOD: KALMAN FILTER ESTIMATION (1999.1.4–2003.12.31)

	Korea	Thailand	Taiwan	Singapore	India	China
γ_0	−0.003	0.006	0.006	−0.001	0.002	−0.000
$\lambda_{\$/SF,0}$	0.025	0.387^{+}	0.110	0.322	1.358	1.169
$\lambda_{¥/SF,0}$	0.003	0.175*	0.033**	0.003	−0.002	−0.000
$\lambda_{EU/SF,0}$	0.039	0.083	0.039	0.062^{+}	0.005	−0.004
$\lambda_{\$/SF,1}$	0.970**	0.504^{+}	0.883**	0.571*	−0.365	−0.172
$\lambda_{¥/SF,1}$	0.983**	0.036	−0.014	0.983	0.462	−0.384*
$\lambda_{EU/SF,1}$	−0.944**	0.016	0.176	0.348	0.865	−0.394^{+}
$\Psi_{\$/SF}$	exp(−6.603)	exp(−3.937)$^{+}$	exp(−6.762)	exp(−5.197)	exp(−4.571)**	exp(−24.176)
$\Psi_{¥/SF}$	exp(−7.300)	exp(−4.223)**	exp(−4.230)**	exp(−9.206)	exp(−22.164)**	exp(−8.530)**
$\Psi_{EU/SF}$	exp(−19.870)	exp(−16.875)	exp(−24.059)	exp(−3.110)*	exp(−5.271)**	exp(−26.908)
Γ	exp(−1.609)**	exp(−2.127)**	exp(−3.307)**	exp(−3.029)**	exp(−3.948)**	exp(−7.175)**
Ln L	−808.273	−534.061	155.198	36.039	549.325	2626.323
Ave. of $\gamma_{\$/SF,t}$	0.836	0.780	0.942	0.751	0.995	0.998
Ave. of $\gamma_{¥/SF,t}$	0.207	0.181	0.033	0.186	−0.002	−0.000
Ave. of $\gamma_{EU/SF,t}$	0.020	0.085	0.048	0.095	0.038	−0.003

	Hong Kong	Malaysia	Indonesia	Philippines	Pakistan	Sri Lanka
γ_0	0.000	0.000	−0.001	0.031	0.004	0.017^{+}
$\lambda_{\$/SF,0}$	1.022	1.996	1.777**	0.887	1.136	0.527**
$\lambda_{¥/SF,0}$	0.002	−0.002	0.090	0.020	0.008	−0.008
$\lambda_{EU/SF,0}$	0.004	0.005	0.008	0.042	0.057	0.047
$\lambda_{\$/SF,1}$	−0.026	−0.992	−0.887**	0.024	−0.149	0.466**
$\lambda_{¥/SF,1}$	0.098	0.003	0.110	0.760	0.407**	0.283
$\lambda_{EU/SF,1}$	0.155	−0.001	0.567	−0.229	0.215	0.480**
$\Psi_{\$/SF}$	exp(−43.587)	exp(−48.831)	exp(−3.994)	exp(−66.389)	exp(−15.350)	exp(−2.766)**
$\Psi_{¥/SF}$	exp(−42.439)	exp(−49.902)	exp(−1.062)**	exp(−90.810)	exp(−2.088)**	exp(−22.945)
$\Psi_{EU/SF}$	exp(−22.958)	exp(−19.834)	exp(−9.209)	exp(−17.034)	exp(−3.256)	exp(−1.217)
Γ	exp(−6.932)**	exp(−5.919)**	exp(0.496)**	exp(−1.021)**	exp(−1.532)**	exp(−2.655)**
Ln L	2565.159	1926.422	−2168.352	−1139.481	−989.206	−394.822
Ave. of $\gamma_{\$/SF,t}$	0.997	1.002	0.942	0.910	0.989	0.988
Ave. of $\gamma_{¥/SF,t}$	0.002	−0.002	0.101	0.089	0.015	−0.011
Ave. of $\gamma_{EU/SF,t}$	0.005	0.005	0.020	0.034	0.073	0.091

Notes: 1) ln L implies the maximum value of log likelihood function.
2) $^{+}$, *, and ** denote significant at the 10%, 5%, and 1% levels, respectively.
3) Ave. of $\gamma_{\$/SF,t}$ ($\gamma_{¥/SF,t}$, $\gamma_{EU/SF,t}$) implies the average of $\gamma_{\$/SF,t}$ ($\gamma_{¥/SF,t}$, $\gamma_{EU/SF,t}$) estimates over the sample period.

TABLE 6. THE HALDANE AND HALL (1991) ESTIMATION METHOD (1973.1.2–1998.12.31)

	α_0	Q	R	ln L	Ave. of α_t	Ave. of $\alpha_t^\blacklozenge$
France	1.236**	exp(−11.002)**	exp(−14.385)**	28798.809	0.740	0.337
Italy	6.903**	exp(−9.175)**	exp(−28.584)	23365.131	0.350	−0.338
UK	−0.844**	exp(−9.905)**	exp(−12.636)**	24839.951	0.464	0.150
Netherlands	0.126**	exp(−12.749)**	exp(−14.400)**	33038.400	0.964	0.913
Belgium	3.029**	exp(−10.981)**	exp(−26.792)	29256.200	0.869	0.645
Norway	1.470**	exp(−10.174)**	exp(−14.149)**	26691.058	0.601	0.248
Sweden	1.553**	exp(−9.215)**	exp(−13.893)**	23903.129	0.399	−0.089
Switzerland	−0.603**	exp(−7.820)**	exp(−8.290)**	17953.709	1.758	1.738
Austria	1.981**	exp(−10.268)**	exp(−24.819)	26930.418	0.961	0.991
Denmark	1.374**	exp(−10.486)**	exp(−14.735)	28025.104	0.788	0.455
Finland	1.126**	exp(−10.254)**	exp(−11.853)**	24831.402	0.594	0.268
Ireland	0.005**	exp(−9.614)**	exp(−13.436)**	24875.266	0.598	0.165
Portugal	4.498**	exp(−8.996)**	Exp(−28.199)	22779.082	0.194	−1.216
Spain	4.376**	exp(−9.518)**	Exp(−29.461)	24483.597	0.453	−0.222

Notes: 1) ln L implies the maximum value of log likelihood function.
2) ** denotes significant at the 1% level.
3) Ave. of α_t implies the average of α_t estimates over the whole sample period.
4) Ave. of $\alpha_t^\blacklozenge$ implies the average of α_t for the sample period from January 2, 1973 to February 28, 1979.

FIG. 1. ASIAN EXCHANGE RATES

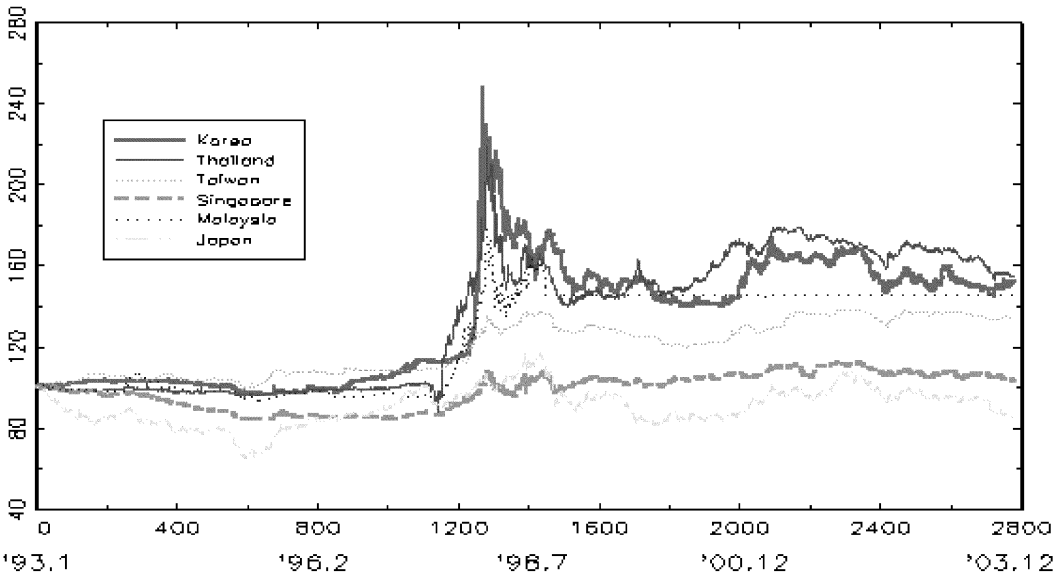

FIG. 2. ASIAN EXCHANGE RATES

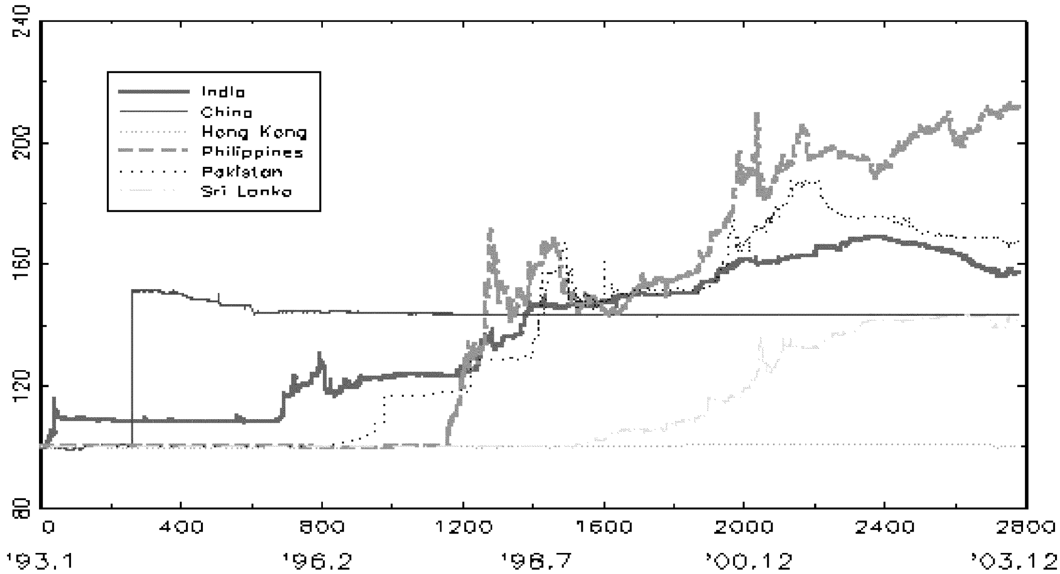

FIG. 3. yen / $ COEFFICIENTS

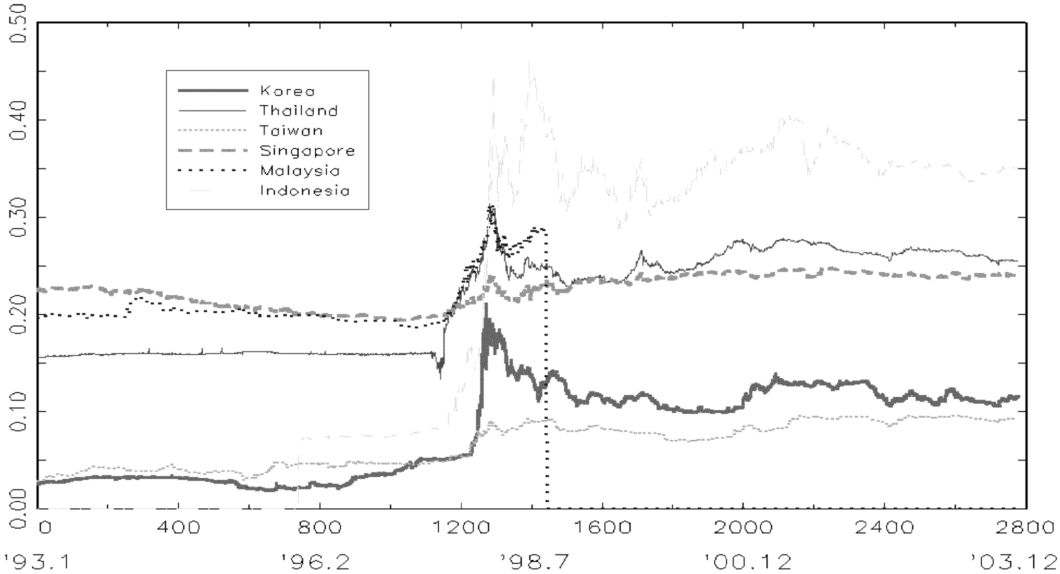

FIG. 4. yen / $ COEFFICIENTS

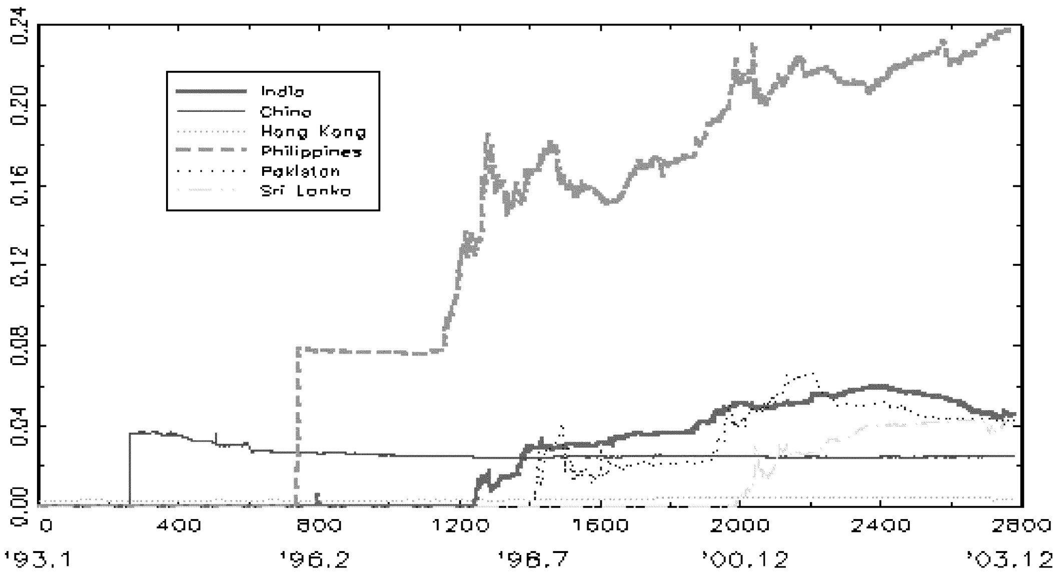

FIG. 5. yen / $ COEFFICIENTS

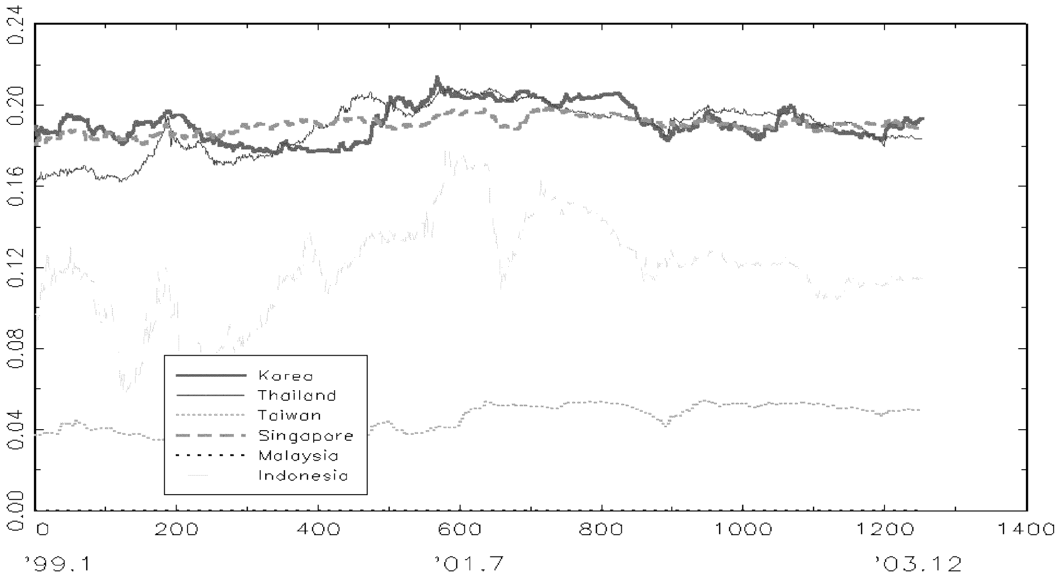

FIG. 6. yen / $ COEFFICIENTS

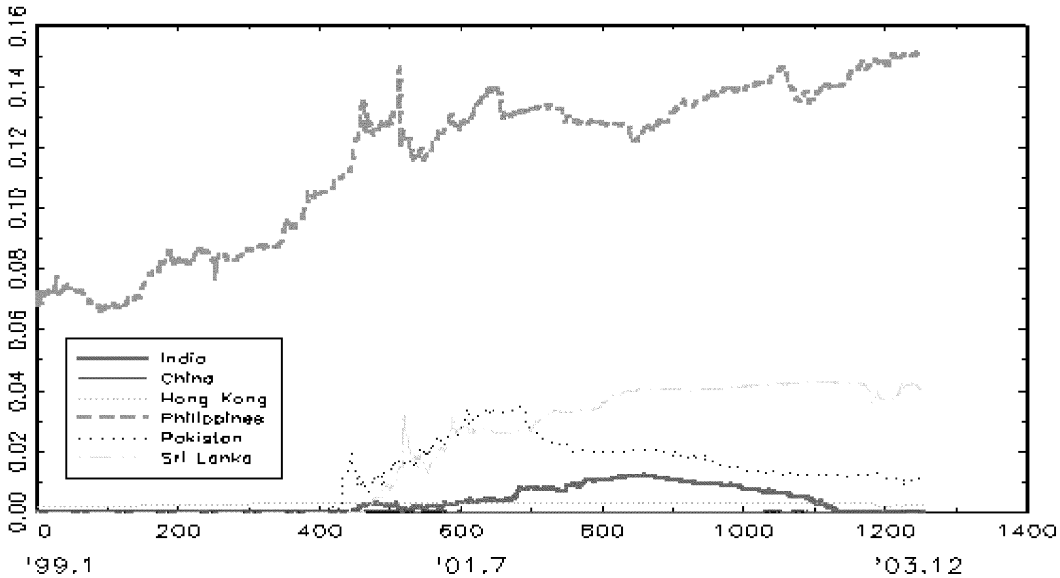

FIG. 7. yen / SF COEFFICIENTS

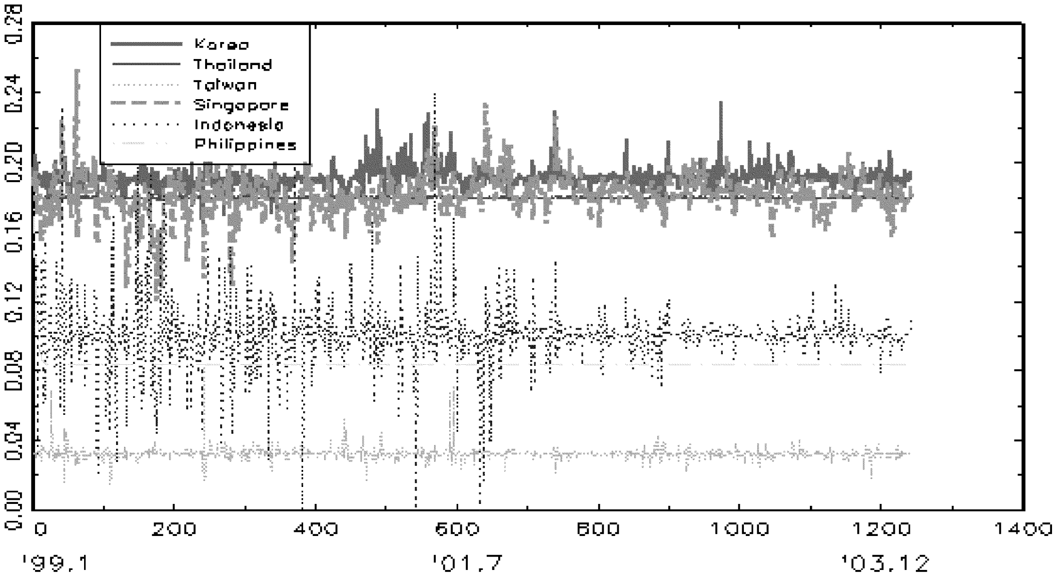

FIG. 8. $ / SF COEFFICIENTS

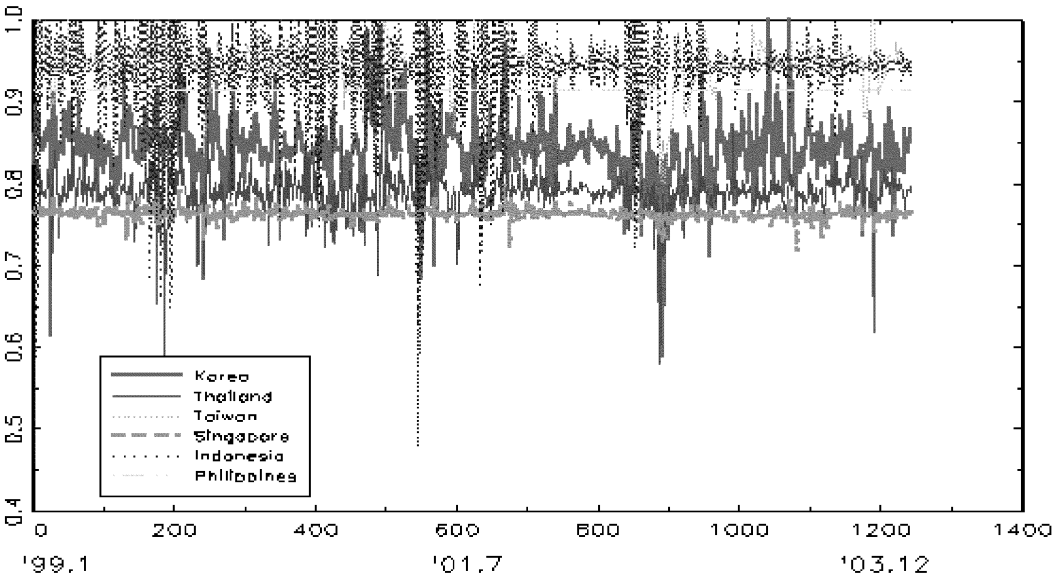

FIG. 9. yen / SF COEFFICIENTS

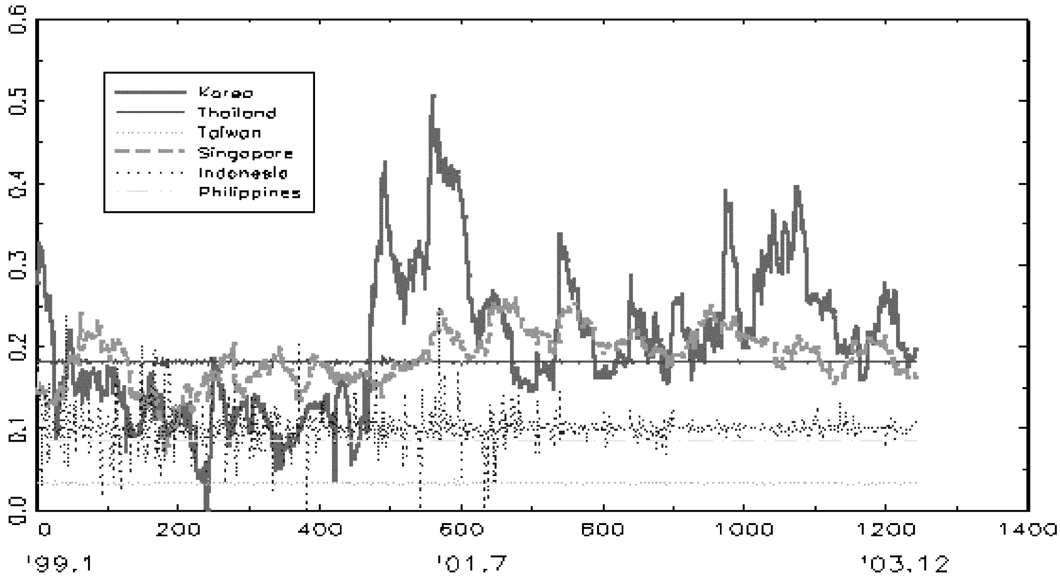

FIG. 10. $ / SF COEFFICIENTS

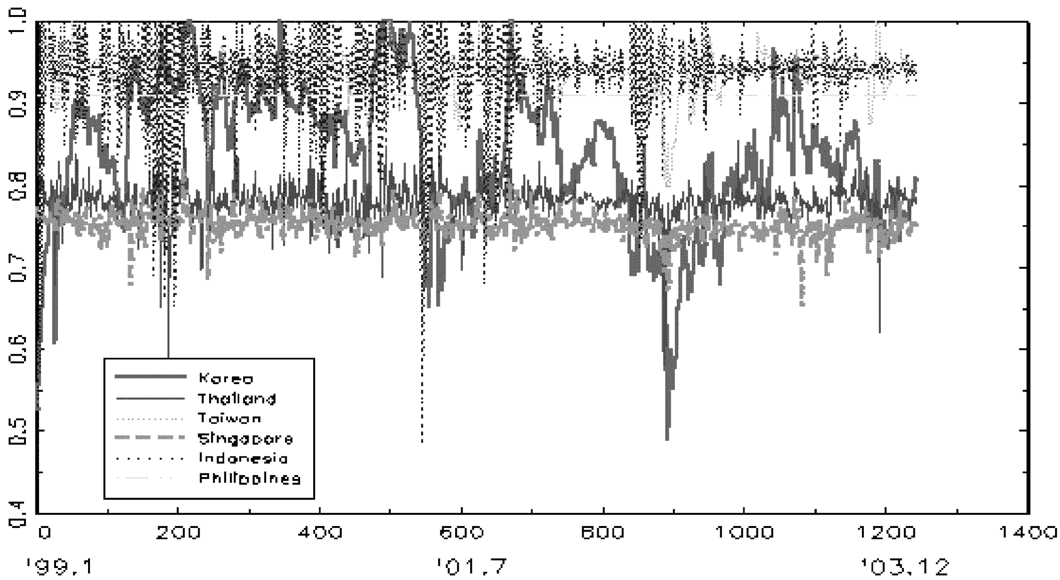

FIG. 11. euro / SF COEFFICIENTS

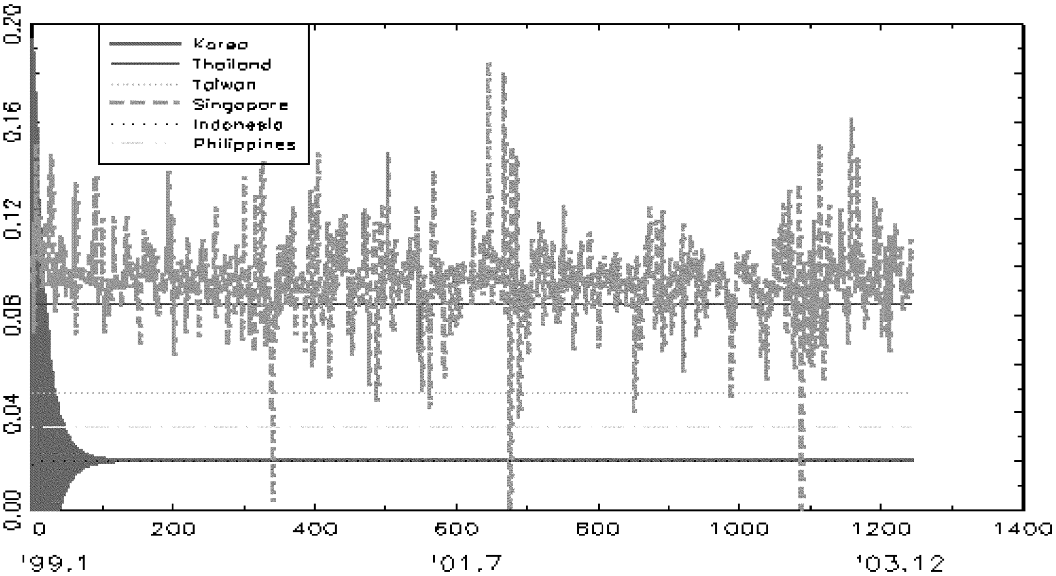

FIG. 12. EUROPEAN EXCHANGE RATES

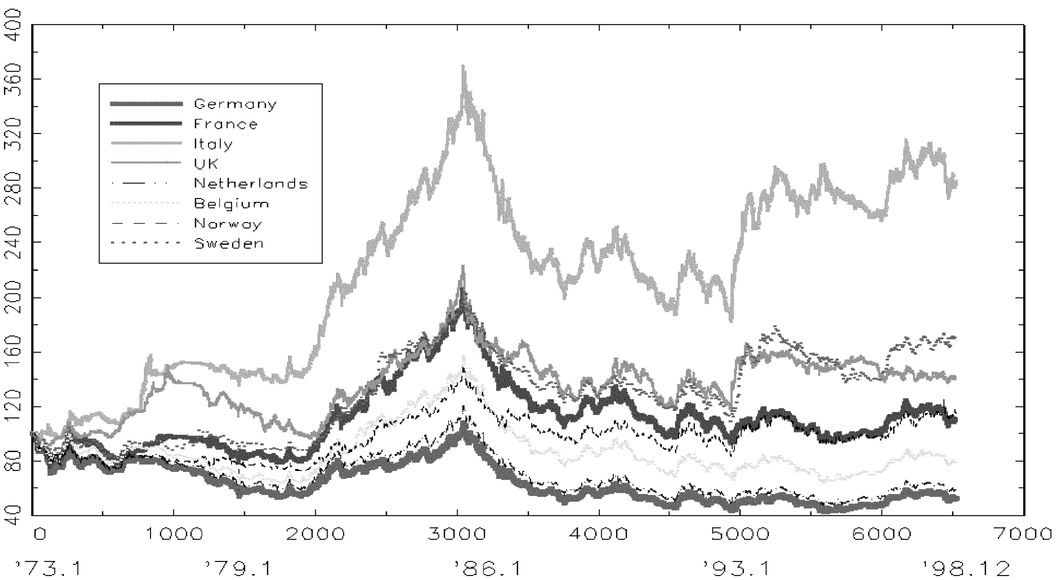

FIG. 13. EUROPEAN EXCHANGE RATES

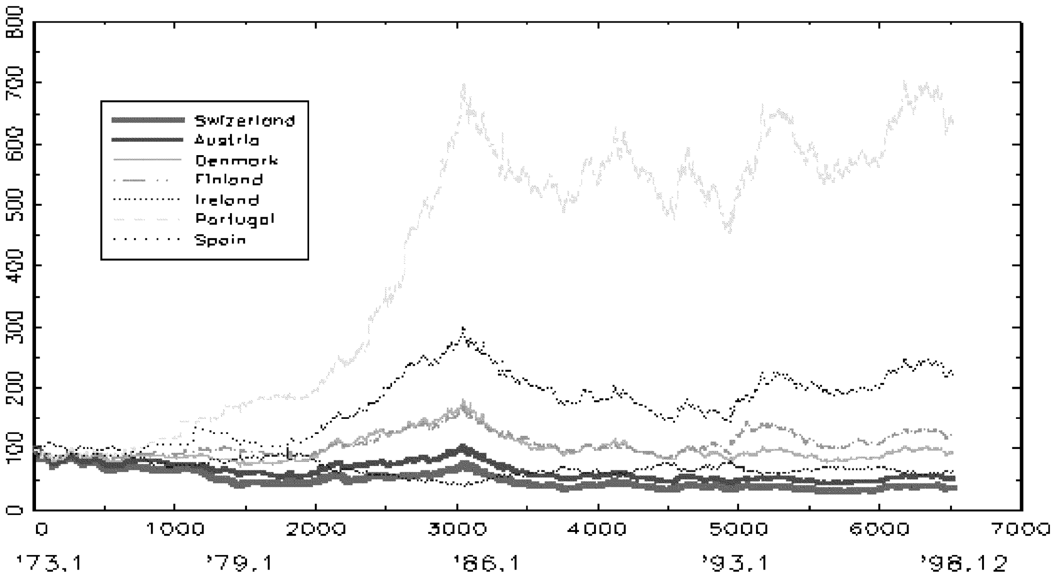

FIG. 14. DM / $ COEFFICIENTS

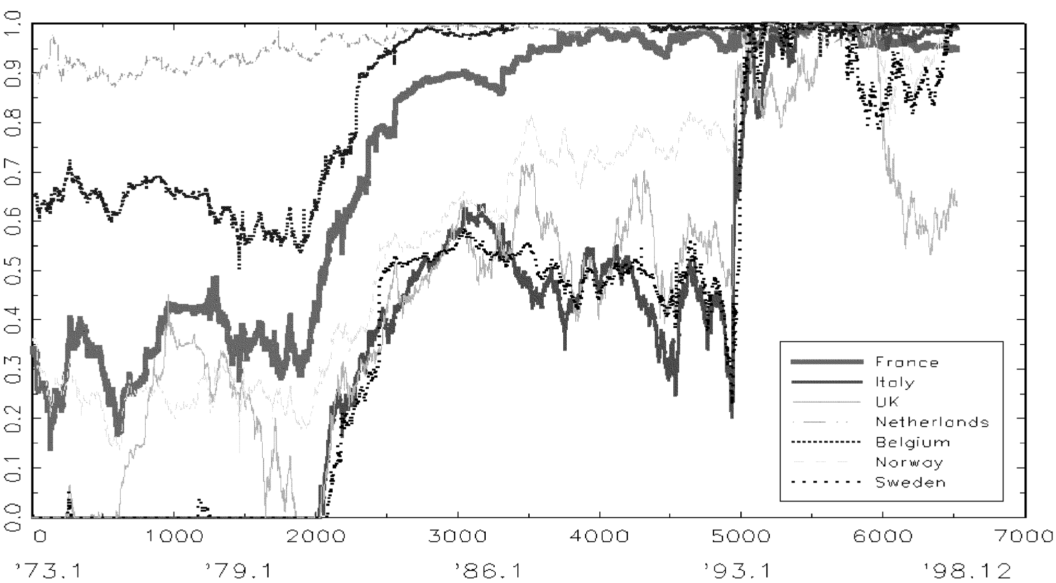

FIG. 15. DM / $ COEFFICIENTS

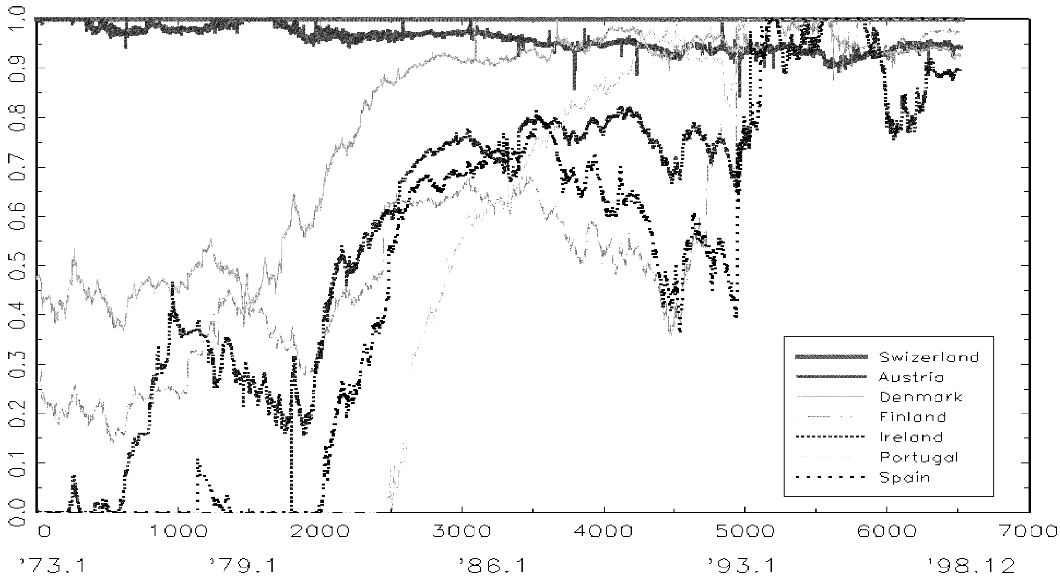

FIG. 16. WC / $ COEFFICIENTS

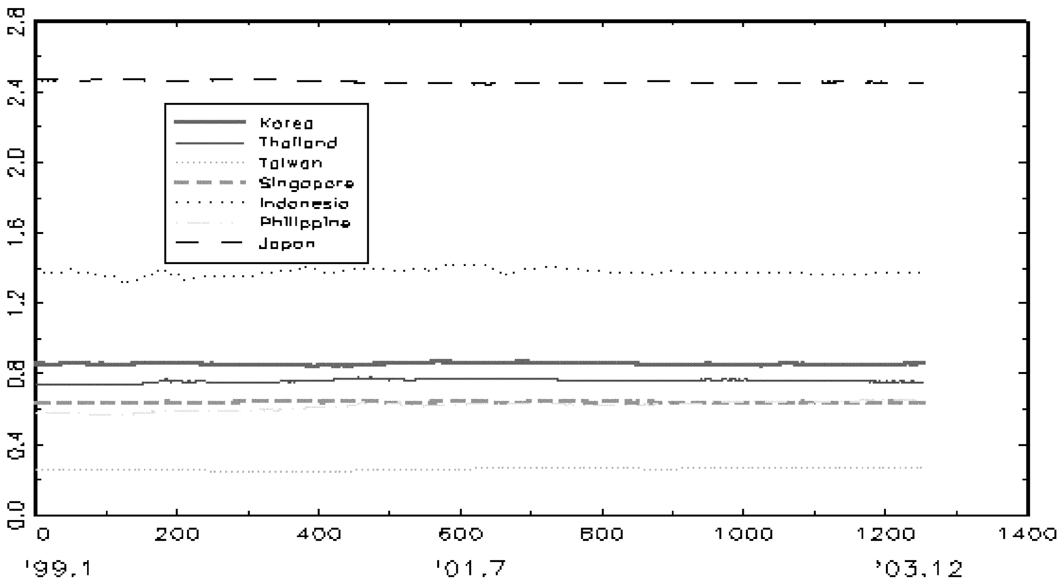

FIG. 17. WC / SF COEFFICIENTS

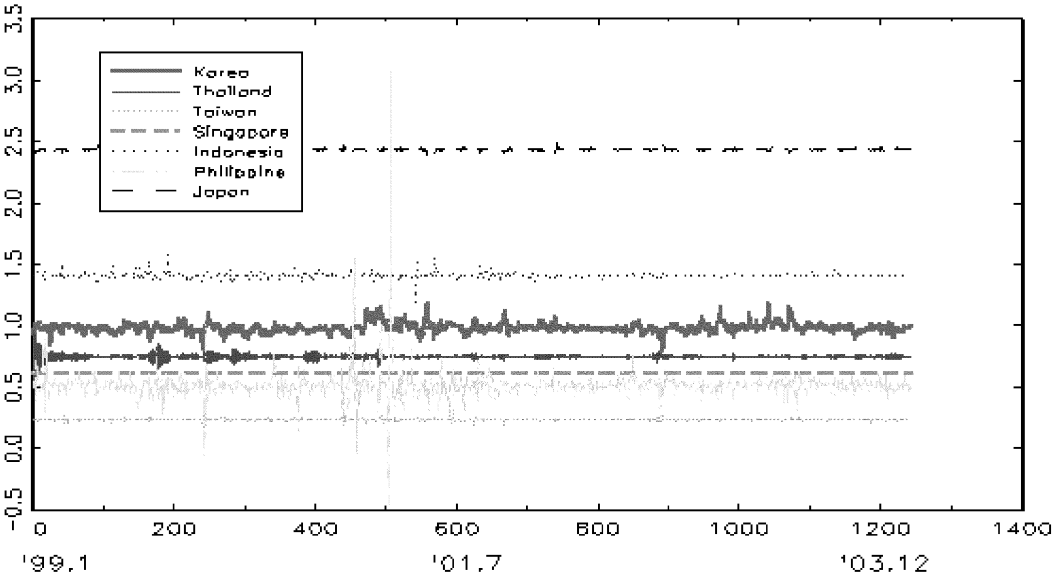

FIG. 18. $ / SF COEFFICIENTS

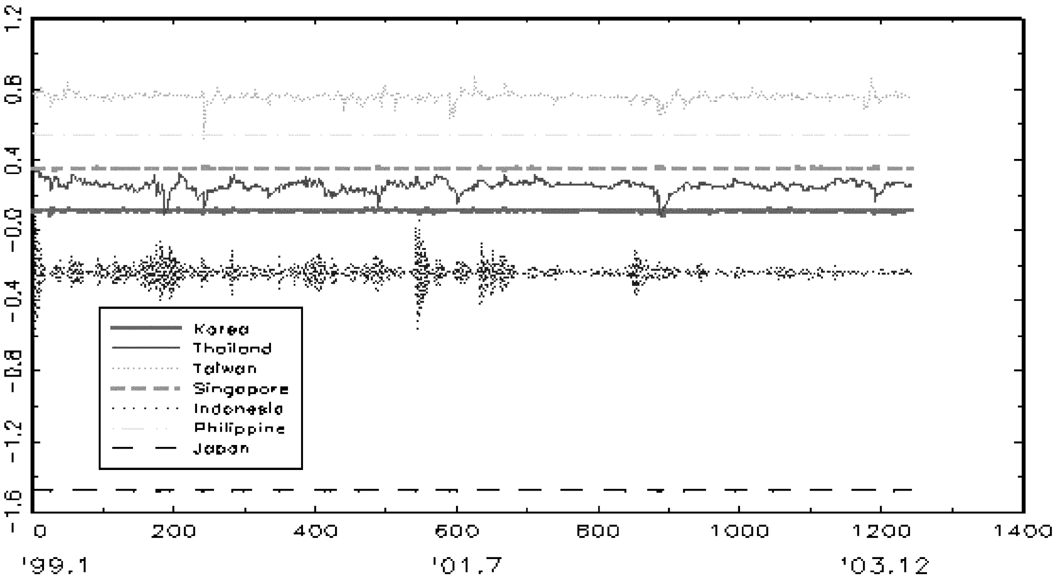

FIG. A1. yen / $ COEFFICIENTS (Average of 5 Days)

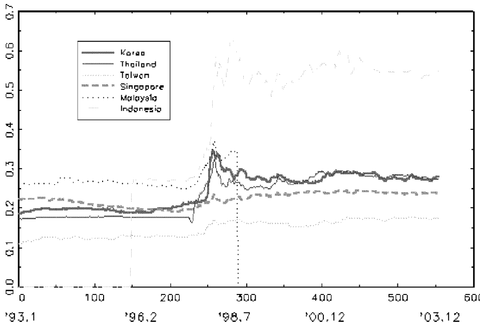

FIG. A2. yen / $ COEFFICIENTS (Average of 5 Days)

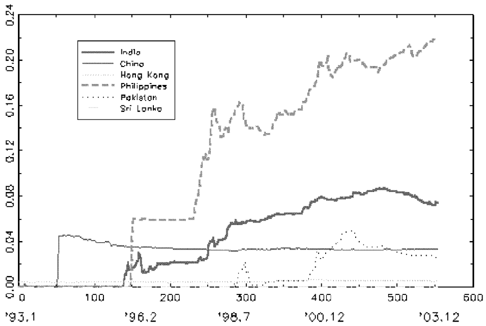

FIG. A3. yen / $ COEFFICIENTS (Average of 20 Days)

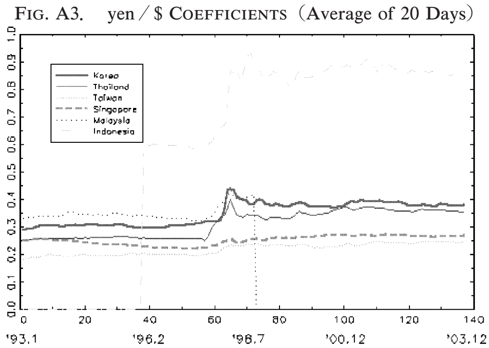

FIG. A4. yen / $ COEFFICIENTS (Average of 20 Days)

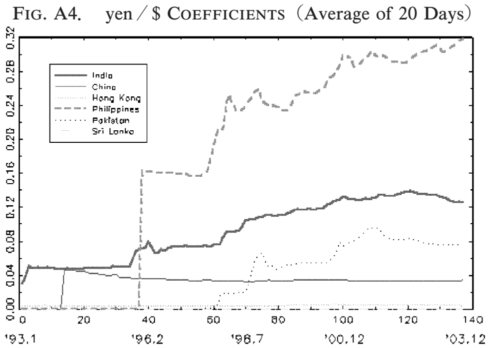

FIG. A5. DM / $ COEFFICIENTS (Average of 5 Days)

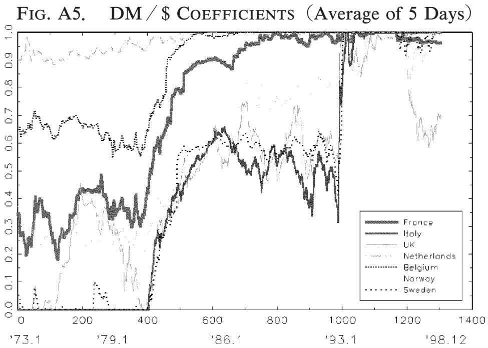

FIG. A6. DM / $ COEFFICIENTS (Average of 5 Days)

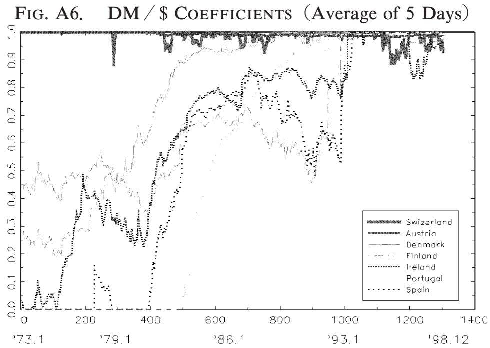

FIG. A7. DM / $ COEFFICIENTS (Average of 20 Days)

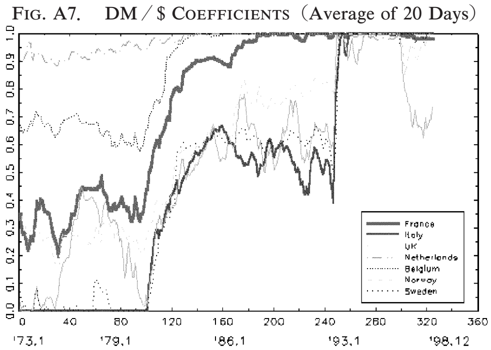

FIG. A8. DM / $ COEFFICIENTS (Average of 20 Days)