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ESSAYS ON EMPIRICAL EXCHANGE RATE ECONOMICS: INTERVENTION, STERILIZATION, AND EXCHANGE RATE POLICY

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Ph.D. Dissertation at Osaka University

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1. Overview

This dissertation presents empirical analyses of recent exchange rate regimes and policies. It consists of four essays dealing with (1) the Louvre Accord and central bank intervention in Japan, (2) the exchange rate policies of East Asian countries, (3) the sterilization of capital inflows in East Asia, and (4) risk premiums and exchange rate expectations in East Asia. The present chapter presents an overview of this dissertation.

1.1. The Louvre Accord and Central Bank Intervention: Was There a Target Zone?

The first essay (found in Chapter 2) presents an empirical analysis of central bank intervention following the Louvre Accord of February 1987 in order to determine whether or not a target zone was indeed adopted and, if so, to characterize the nature of that target zone. Over the past decade, a considerable number of theoretical studies have analyzed the exchange rate target zone (*e.g.*, Krugman (1991), Bertola and Caballero (1992) and Bertola and Svensson (1993)). However, comparatively little attention has so far been paid to the target zone which is believed by some to have been adopted following the Louvre Accord.

On February 22, 1987, major industrial countries agreed that they would coordinate macroeconomic policies to stabilize exchange rates at "around current levels" in what became known as the Louvre Accord. Although the details of the agreement were not made public, it is suggested in the popular literature that the countries adopted target zones as a way of maintaining exchange rate stability (Funabashi (1989)). It should be noted, however, that no official statement ever confirmed the adoption of a target zone. Thus, we first verify whether or not the Bank of Japan and the US Federal Reserve indeed adopted the target zone arrangement to stabilize the yen-dollar exchange rate, by using daily foreign exchange intervention data. We find that, at least operationally, the central banks did adopt a target zone during the period, to the extent that they intervened to stabilize the exchange rate at the presumed Louvre target levels and frequently coordinated their intervention operations at the lower edge of the band for the US dollar.

The exchange rate would not be stable in a target zone, unless the zone were credible, as showed by Bertola and Caballero (1992) and Bertola and Svensson (1993). Thus, we then estimate the expected future exchange rate and the expected rate of devaluation to examine how credible the target zone was during the period immediately following the Louvre Accord, by assuming that a target zone was indeed put in place. It is shown that the target zone for the yen-dollar exchange rate was less credible than that for the deutsche mark-dollar exchange rate, because (1) the yen's expected future exchange rate was frequently outside the presumed exchange rate band and (2) the expected rate of devaluation was more volatile for the yen-dollar rate than that for the deutsche mark-dollar rate.

The contribution of this essay lies in showing empirically that the Bank of Japan and the US Federal Reserve did adopt a target zone arrangement following the Louvre Accord of February 1987, but that the target zone for the yen-dollar exchange rate was not credible.

1.2. Was It Really a Dollar Peg?: The Exchange Rate Policies of East Asian Countries, 1980-97

The second essay (found in Chapter 3) presents a quantitative analysis of the exchange rate policies of six East Asian countries (*i.e.,* Thailand, Korea, Singapore, Malaysia, Indonesia and the Philippines) during the 17-year period preceding the currency crisis of 1997 and examine whether the East Asian currencies were actually pegged to the US dollar. Subsequent to the dramatic fall of the Thai baht on July 2, 1997, other East Asian currencies were also subjected to speculative pressure and the authorities were forced to allow them to depreciate sharply. In the context of this Asian currency crisis of 1997, it is often claimed that their currencies were effectively pegged to the US dollar.

This common argument linking the supposed dollar peg with the currency crisis of 1997, however, is curious for at least two reasons. First, at least officially, all of the East Asian countries or regions, except Hong Kong, had claimed to have a relatively flexible exchange rate policy during the period of at least 10 years leading up to the currency crisis. Second, the pioneering works of Frankel (1992) and Frankel and Wei (1994) indicate that the implicit weight of the dollar was large, but not unity, in the determination of the nominal values of these currencies. Moreover, the literature also suggests that the weight of the dollar was by no means fixed and that the weight of the yen did seem to increase over time (Kwan (1995)). It is thus possible that the exchange rate policies of East Asian countries were too involved to be characterized as a simple dollar peg or possibly even as a de facto dollar peg.

Accordingly, we estimate the implicit weights of foreign currencies in the nominal exchange rate determination of the East Asian currencies by means of a time-varying parameter model to ascertain whether they were actually pegged to the US dollar. It is found that the authorities of Korea and Malaysia significantly raised the weight of the Japanese yen, when it depreciated sharply against the US dollar (May 1995 - April 1997). On the other hand, the authorities of Singapore significantly raised the weight of the Japanese yen, when it appreciated sharply against the US dollar. These observations suggest that the important objectives of exchange rate policies were export promotion in Korea and Malaysia and price stability in Singapore.

It is also shown that the weight of the US dollar was large in the exchange rate policies of all countries. In the case of Indonesia and the Philippines, the weight was virtually 100 percent, so that it may rightly be said that the rupiah and the peso were effectively pegged to the US dollar. For the other currencies, however, the weight of the yen did increase in the 1990s. The weight was by no means insignificant in Thailand (almost 12 percent), Korea (almost 14 percent), Singapore (about 14 percent) and Malaysia (around 10 percent) when the currency crisis struck. In the crisis countries of Thailand, Korea and Malaysia, moreover, the weight was steadily rising during the preceding two years along with the depreciation of the yen against the US dollar. Thus, it is incorrect to presume that the exchange rate policies of the East Asian countries were characterized as a simple dollar peg or possibly even as a de facto dollar peg.

The contribution of this essay lies in using a time-varying parameter model to estimate the implicit weights of foreign currencies in the nominal exchange rate determination of the East Asian currencies and in showing that they were not simply pegged to the US dollar during the period immediately preceding the Asian crisis.

1.3. Sterilization and the Capital Inflow Problem in East Asia, 1987-97

The third essay (found in Chapter 4) presents an empirical analysis of the effects of sterilization of capital inflows on the growth of monetary aggregates in East Asia. The East Asian countries of Indonesia, Korea, Malaysia, the Philippines and Thailand received large volumes of capital inflows from the end of the 1980s through early 1997. The cumulative inflows were massive indeed, amounting to 50 percent of GDP in Malaysia and Thailand, over 20 percent in the Philippines and about 10 percent in Indonesia and Korea. Because of the potential risks they entail, these capital inflows were almost from the inception considered as posing a serious challenge for macroeconomic management, leading the profession to coin the term "the capital inflow problem".

Against the surge in capital inflows, East Asian countries took various policies, including capital controls, trade liberalization, greater exchange rate flexibility, fiscal contraction and a variety of monetary measures (Montiel (1998), Reinhart and Reinhart (1998) and Villanueva and Seng (1999)). In particular, the monetary authorities of East Asia took various monetary measures, including the conventional form of sterilization intervention (designed to offset the effect of reserve inflows on the monetary base by open market sales of domestic securities, defined as sterilization in the narrow sense), increases in reserve requirement (designed to limit the impact of reserve inflows on the growth of monetary aggregates by reducing the money multiplier), shifting of government deposits from commercial banks to the central bank, an increase in the discount rate or otherwise a greater limit on the discount window, moral suasion and credit controls (defined as sterilization in the broader sense).

However, only few empirical studies have so far been made to examine the effectiveness of sterilization (in the broader sense) in East Asia. Thus, we indirectly test whether the set of various sterilization measures pursued were effective in limiting the growth of narrow and broad money by using both time-series and structural approaches. Econometric tests based on quarterly data suggest that the set of various sterilization measures were effective in limiting the growth of narrow and broad money. This may have promoted additional capital inflows by keeping the level of domestic interest rates high, or caused disintermediation and expanded the volume of assets in the poorly supervised nonbank financial sector. It can be surmised that, in either case, the potential risk of capital inflows in East Asia was likely magnified by the active policy of sterilization.

The contribution of this essay lies in being the first to examine the effects of sterilization of capital inflows on the growth of monetary aggregates in East Asia and in showing that the capital inflow problem leading up to the crisis of 1997 was made more serious by the active and persistent policy of sterilization.

1.4. Risk Premiums and Exchange Rate Expectations: A Reassessment of the So-Called Dollar Peg Policies of Crisis East Asian Countries, 1994-97

The fourth essay (found in Chapter 5) presents a statistical analysis of foreign exchange risk premiums and exchange rate expectations in the East Asian countries of Indonesia, Korea, Malaysia and Thailand during the 42-month period immediately preceding the onset of the East Asian currency crisis in July 1997 to assess the credibility of the so-called dollar peg policies. It is well known that, from the end of the 1980s through early 1997, the emerging market economies of East Asia received a large volume of capital inflows. While responsible for this surge of capital inflows were both internal (or "pull") and external (or "push") factors of various types, the substantial interest rate differentials that existed in favor of assets denominated in East Asian currencies over those denominated in major industrial country currencies were undoubtedly an important contributing factor. Moreover, adjusted for actual exchange rate changes, the average excess returns over US dollar-denominated instruments remained substantial, amounting to 2.5-6.1 percent per year.

In this connection, it is important to remember that these large positive excess returns on East Asian currency assets were observed against the background of the so-called dollar peg policies, the exchange rate policies of maintaining relative stability against the US dollar (Ito, Ogawa and Sasaki (1998), Takagi (1999) and Chapter 3 of this thesis). The presence of interest rate differentials favoring East Asian currency assets in this environment means that there were risk premiums, expected depreciation, or some combination of both. Equivalently, the presence of positive ex post excess returns means that there were risk premiums, unexpected appreciation, or some combination of both on the part of East Asian currency assets. If the exchange rate policies had been credible in the sense that market participants expected the US dollar exchange rates to remain stable, one would have observed the expected rate of currency depreciation to be small. Then, most of the interest rate differentials would be explained by foreign exchange risk premiums. On the other hand, lack of credibility in the dollar peg policies would have meant that an important component of the interest rate differentials would reflect expected exchange rate change. An important task, therefore, is to decompose the observed interest rate differentials into risk premiums and expected rates of depreciation, which are both unobservable.

In this essay, we accordingly use an unobserved components model to extract foreign exchange risk premiums from the ex post excess returns of East Asian currency assets over the US dollar assets and derive the implied expected future spot rates of East Asian currencies against the US dollar. It is found that, during the period under consideration, risk premiums were substantial and time-varying, suggesting that East Asian currency assets and US dollar assets were imperfect substitutes. This evidence indicates that sterilization of capital inflows was effective, as documented by Chapter 4 for these and other East Asian countries during 1987-97. It is shown that market participants consistently formed expectations of either appreciation or depreciation, suggesting that the so-called dollar peg policies were not credible.

The contribution of this essay lies in directly estimating the risk premiums of crisis East Asian currency assets as a way of driving market participants' exchange rate expectations and in thereby showing that the so-called dollar peg policies were not credible.

2. The Louvre Accord and Central Bank Intervention: Was There a Target Zone? [1]

2.1. Introduction

Over the past decade, a considerable number of theoretical studies have analyzed the exchange rate target zone (*e.g.*, Krugman (1991), Bertola and Caballero (1992) and Bertola and Svensson (1993)). However, comparatively little attention has so far been paid to the target zone which is believed by some to have been adopted following the Louvre Accord.[2] Accordingly, this chapter will present an empirical analysis of central bank intervention following the Louvre Accord of February 1987 in order to determine whether or not a target zone was indeed adopted and, if so, to characterize the nature of that target zone. In particular, this chapter will examine whether the Bank of Japan and the US Federal Reserve adopted a target zone in order to stabilize the yen-dollar exchange rate during the 10-month period following the Louvre Accord. If such a target zone is found to have existed, moreover, the chapter will further analyze how credible the zone might have been by estimating the expected future exchange rate and the expected rate of devaluation.

On February 22, 1987, major industrial countries agreed that they would coordinate macroeconomic policies to stabilize exchange rates at "around current levels" in what became known as the Louvre Accord. Although the details of the agreement were not made public, it is suggested in the popular literature that the

[1] This chapter is largely based on Esaka (2000).

[2] As notable empirical studies, Flood, Rose and Mathieson (1990) and Lindberg and Söderlind

countries adopted target zones as a way of maintaining exchange rate stability (Funabashi (1989)). Because they announced neither the central rates nor the bands for the exchange rate, it may be said that these were unofficial target zones, if the target zone arrangement was adopted at all.

According to Funabashi (1989), the central rates following the Louvre Accord were supposedly 153.50 yen and 1.825 marks per dollar with a band of ± 5 percent. In the case of the yen-dollar rate, it is also said that, on April 7, the central rate was rebased to 146 yen per dollar in order to reflect the new market conditions. It should be noted, however, that no official statement ever confirmed the adoption of a target zone. The first of our tasks in this chapter is thus to follow Lewis (1990) and to verify whether or not the central banks indeed adopted the target zone arrangement.

According to Bertola and Caballero (1992), Bertola and Svensson (1993) and others, the exchange rate would not be stable in a target zone, unless the zone were credible. The issue of target zone credibility in the context of the Exchange Rate Mechanism (ERM) of European Monetary System (EMS) and the Sweden krona has been analyzed by Bertola and Svensson (1993), Lindberg, Söderlind, and Svensson (1993) and Svensson (1993), by estimating the expected rate of devaluation. Thus, our second task is to follow their methodology to examine how credible the target zone was during the period immediately following the Louvre Accord, by assuming that a target zone was indeed put in place.

The chapter is organized as follows. Section 2.2 will present an overview of the Louvre Accord and post-Louvre intervention. Section 2.3 will ascertain whether or not the Bank of Japan and the US Federal Reserve adopted a target zone, by using a multinomial logit model. Section 2.4 will examine the credibility of the presumed

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(1994) have examined the post-Louvre (as well as pre-Louvre) target zone in the context of the target zone by estimating the expected future exchange rate and the expected rate of devaluation. Section 2.5 will present a summary and concluding remarks. Finally, the Appendix will outline the sources of data.

2.2. The Louvre Accord and Central Bank Intervention: An Overview

2.2.1. The Louvre Accord

Let us begin by presenting an overview of the Louvre Accord. The dollar had depreciated substantially against the other major currencies during the period following the Plaza Agreement of September 1985. In particular, the yen had appreciated from 240 yen to 155 yen to the dollar. The authorities of most major countries recognized that a further substantial shift in the value of their currencies could damage the prospect for economic growth, and agreed to stabilize the exchange rates at "around current levels". On February 22, 1987, this agreement was officially reached at the meeting of the Group of Five (G5) countries held at the Palais du Louvre. Although the details of the agreement were never made public, Funabashi (1989) and others in the press reported that major countries did in fact adopt a target zone at that time.

According to Funabashi (1989), the major elements of the Louvre Accord are as follows. First, the central rates were 153.50 yen and 1.825 deutche marks per dollar. Second, the margin of 2.5 percent from the central rate was determined as a first line of defense for mutual intervention. Third, intervention efforts were expected to intensify from the 2.5 percent margin to the 5 percent margin. Fourth, intervention was on a consultative basis rather than on an obligatory basis. Fifth, the central rates

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Exchange Rate Mechanism of the European Monetary System.

could be rebased by mutual agreement. Sixth, the agreements were to be kept strictly confidential and were provisional until the G5 meetings scheduled to be held in Washington in April. Seventh and finally, the central rates as well as the bands would not be made public.

Figure 2.1 shows the movements of the yen and the deutsche mark against the US dollar from the day after the Louvre Accord (February23, 1987) to the day before the stock market crash (October 18,1987), with the broken lines indicating the upper and lower bounds of the target zone, as reported by Funabashi (1989). From this figure, we observe that the yen and the deutsche mark were almost entirely inside the exchange rate bands, and that the deutsche mark was particularly close to the central rate.

2.2.2. Central Bank Intervention

Next, we investigate how the Bank of Japan and the Federal Reserve might have intervened to stabilize the exchange rate during the period.[3] If the central banks had adopted the target zone regime, we may conjecture that they must have intervened when the exchange rate was close to the upper or lower edge of the band. Although analysis of daily central bank intervention requires daily intervention data, the Japanese authorities do not make intervention data public. Thus, in what follows, we obtain an estimate of daily intervention by using the newspaper accounts of the *Nihon Keizai Shinbun*, Japan's leading daily business paper.

Figure 2.2 shows how the central banks might have intervened in the foreign exchange market on a daily basis, on the basis of accounts in the *Nihon Keizai*

[3] For empirical analyses of central bank intervention, see Dominguez (1990), Takagi (1991) and Domingues and Frankel (1993). Takagi (1991) shows that intervention by the Bank of Japan was almost completely sterilized and that the monetary authorities followed a "leaning against the wind" policy during much of the period of 1973-89.

Shinbun. Here, no intervention, dollar-selling intervention, and dollar-purchasing intervention are represented as 0, -1, and 1, respectively. From this figure, we observe that dollar-purchasing intervention was made by the Bank of Japan when the yen appreciated from the central rate; it was made most frequently when the yen was near the lower edge of the band. On the other hand, there apparently is no evidence to show that the Bank of Japan intervened when the yen depreciated from the central rate.

Similarly, we observe that dollar-purchasing intervention was made by the Federal Reserve when the yen appreciated from the central rate. On the other hand, dollar-selling intervention was sometimes made by the Federal Reserve when the yen depreciated from the central rate. It can be observed that dollar-purchasing intervention was made most frequently by the Federal Reserve when the yen was near the lower edge of the band. Hence, we may surmise that the Bank of Japan and the Federal Reserve coordinated their intervention operations when the yen appreciated from the central rate, but not when the yen depreciated.

2.3. The Econometric Evidence of a Target Zone: A Multinomial Logit Model

2.3.1. The Multinomial Logit Model of Intervention

Let us now examine how the central banks might have intervened to stabilize the exchange rates by using the multinomial logit model,[4] as in Lewis (1990, 1995). Lewis (1990, 1995), however, used the exchange rates at 7:00 a.m. Eastern Standard Time (EST) in London at date *t-1*, whereas the Bank of Japan intervened in Tokyo (at least 12 hours apart) and the Federal Reserve intervened in New York (at least 26

[4] For details of the multinomial logit model, see Greene (1993).

hours apart). Thus, we will improve on her data set by using more closely synchronized exchange rate data.

Because the foreign exchange market is open almost 24 hours across national borders, the exchange rate fluctuates considerably throughout the day. It seems reasonable to assume that the monetary authorities know the preceding developments in a foreign market before deciding to intervene in the home market. We assume that they take either one of the following three policy decisions: no intervention, dollar-selling intervention, and dollar-purchasing intervention, where the intervention dummy variable is defined as,

$I_t = 0$ for no intervention,

$I_t = -1$ for dollar-selling intervention,

and $I_t = 1$ for dollar-purchasing intervention.

The probability of intervention is defined as a logistic distribution and the probabilities are estimated as the function of a constant term and the exchange rate prevailing before the opening of the domestic foreign exchange market.[5] The probability of intervention can thus be represented as,

$$P\,(I_t = 0|S_{BO}) = \frac{\exp(\alpha_0 + \alpha_1 S_{BO})}{1 + \exp(\beta_0 + \beta_1 S_{BO}) + \exp(\alpha_0 + \alpha_1 S_{BO})}, \qquad \text{(1-a)}$$

$$P\,(I_t = -1|S_{BO}) = \frac{1}{1 + \exp(\beta_0 + \beta_1 S_{BO}) + \exp(\alpha_0 + \alpha_1 S_{BO})}, \qquad \text{(1-b)}$$

[5] Because the probability of intervention is defined as the function of the exchange rate prevailing before the opening of the domestic foreign exchange market, there is no problem of simultaneity between intervention and the level of the exchange rate.

and $P(I_t = 1|S_{BO}) = \dfrac{\exp(\beta_0 + \beta_1 S_{BO})}{1 + \exp(\beta_0 + \beta_1 S_{BO}) + \exp(\alpha_0 + \alpha_1 S_{BO})}$, (1-c)

where α_0, α_1, β_0 and β_1 are coefficients to be estimated, S_{BO} denotes the spot exchange rate prevailing before the opening of the domestic foreign exchange market (in units of the domestic currency per unit of the foreign currency) and $P(\cdots|S_{BO})$ denotes the conditional probability of intervention, given the level of the exchange rate prevailing before the opening of the domestic foreign exchange market. Taking the logarithms of equations (1-a), (1-b) and (1-c) and combining , we obtain,

$$\ln\ (P\ (I_t = 0) / P\ (I_t = -1)|S_{BO}) = \alpha_0 + \alpha_1 S_{BO}, \tag{2-a}$$

$$\ln\ (P\ (I_t = 1) / P\ (I_t = -1)|S_{BO}) = \beta_0 + \beta_1 S_{BO}, \tag{2-b}$$

and
$$\ln\ (P\ (I_t = 1) / P\ (I_t = 0)|S_{BO}) = (\beta_0 - \alpha_0) + (\beta_1 - \alpha_1) S_{BO}. \tag{2-c}$$

In equation (2-a), the appreciation of the yen is shown to increase the probability of no intervention relative to dollar-selling intervention, such that $\alpha_1 < 0$. In equation (2-b), the appreciation of the yen is shown to increase the probability of dollar-purchasing intervention relative to dollar-selling intervention, such that $\beta_1 < 0$. Finally, in equation (2-c), the appreciation of the yen is shown to increase the probability of dollar-purchasing intervention relative to no intervention, such that $(\beta_1 - \alpha_1) < 0$, or $|\beta_1| > |\alpha_1|$.

In what follows, data on daily foreign exchange intervention will be compiled from the news paper accounts of the *Nihon Keizai Shinbun.* In order to correspond to our assumption, we use the closing yen-dollar rate in New York at date *t-1* in the case of the Bank of Japan and the yen-dollar rate at 7:00 a.m. EST (9:00 p.m. Japan Time) in London at date t in the case of the Federal Reserve.[6]

2.3.2. Estimating the Multinomial Logit Model of Intervention

Table 2.1 shows the result of estimating a multinomial logit model of central bank intervention during the period of February 23 - October 18, 1987 (sample period I) and the period of April 7- October 18, 1987 (sample period II).[7] In particular, we estimate separate multinomial logit models for intervention by the Bank of Japan, intervention by the Federal Reserve, coordinated intervention,[8] and combined intervention, by using the maximum likelihood (ML) method. Here, coordinated intervention is defined as an operation in which both of the central banks intervene in the same direction on the same day, whereas combined intervention is defined as an operation in which either of the central banks intervene.[9]

From this table, we note that α_1 is significantly negative for Bank of Japan intervention, and that both α_1 and β_1 are significantly negative, with $|\beta_1|>|\alpha_1|$, for Federal Reserve intervention. In the case of coordinated intervention, α_1 is significantly negative during period II, while α_1 is negative but not significant during period I.[10] In the case of combined intervention, both α_1 and β_1 are significantly negative, with $|\beta_1|>|\alpha_1|$.

From these results, it would be reasonable to suppose that the model is consistent

[6] The Tokyo foreign exchange market opens at date t (9:00 a.m. Japan Time), two and a half hours after the New York foreign exchange closes at date t-1 (4:30 p.m. EST). As the New York market opens at 9:00 a.m. EST, the business hours of London and New York overlap, so that we cannot use the closing yen-dollar rate in London for the Federal Reserve.

[7] Although we tried to estimate the same multinomial logit model for the period of February 23 - April 6, 1987, the coefficients were found not to converge to a finite value.

[8] Lewis (1990, 1995) did not consider what we call coordinated intervention.

[9] In the case of coordinated intervention and combined intervention, we use the closing yen-dollar rate in New York at date t-1 in order to avoid possible simultaneity between intervention and the level of the exchange rate.

[10] However, α_1 is significantly negative at the 15 percent level during period .

with the data, because both α_1 and β_1 are significantly negative. That is to say, it is statistically confirmed that dollar-purchasing intervention was made by the Bank of Japan when the yen appreciated from the target level; dollar-purchasing intervention was made by the Federal Reserve when the yen appreciated, and dollar-selling intervention was made when the yen depreciated. At least statistically, we can be reasonably sure that the central banks coordinated their intervention operations when the yen appreciated from the target level.

2.3.3. Estimating the Probability of Intervention and the Target Level

Figure 2.3 shows the estimated probabilities of different types of intervention from the multinomial logit model. From this figure, we observe that the probability of dollar-purchasing intervention by the Bank of Japan rose from the latter half of March to May, reaching the peak of 44 percent in the first half of the month. The average probability for the period was 23 percent. On the other hand, we cannot estimate the probability of dollar-selling intervention by the Bank of Japan, because no such intervention was observed.

Likewise, we observe that the probability of dollar-purchasing intervention by the Federal Reserve rose from the latter half of March to April, reaching the peak of 25 percent. The average probability for the period was 12 percent, so that the probability of dollar-purchasing intervention by the Federal Reserve was not higher than that for the Bank of Japan. It can also be observed that the probability of dollar-selling intervention by the Federal Reserve rose sharply to 17 percent, when the yen became very close to the upper edge of the presumed band.

Here, the question is what target levels the central banks adopted during the period. If we define the target level as that level of the exchange rate at which the probability of dollar-purchasing intervention is equal to the probability of dollar-selling intervention, such that

$$\ln\ (P\ (I_t = 1) / P\ (I_t = -1)|S_{BO}) = \beta_0 + \beta_1 S_{BO} = 0, \tag{3}$$

the target level (S^*) can be estimated as,

$$S^* = -(\beta_0 / \beta_1)\text{, (\textit{target level (1)}).} \tag{4}$$

Following Lewis (1995), the target level can alternatively be defined as the level of the exchange rate at which the probability of intervention is minimized,

$$S^* = \frac{\ln\ (\alpha_1 / (\beta_1 - \alpha_1)) - \beta_0}{\beta_1}\text{, (\textit{target level (2)}).} \tag{5}$$

In what follows, we will use both of these definitions.

Table 2.2 reports the estimated target levels. From this table, we can see that the target level (2) for the Federal Reserve was estimated as 148.07 yen per dollar and that the target level (2) for the central banks combined was estimated as 150.08 yen during the period I. We can also see that the target level (2) for the Federal Reserve was 146.68 yen and that the target level (2) for the central banks combined was 148.02 yen during period II. Hence, we can reasonably confirm that the estimated target levels were closer to the central rates reported by Funabashi (1989) than the target levels estimated by Lewis (1995) for the period of February 22 to October 18,1987, possibly reflecting that fact that Lewis (1990, 1995) did not consider the debasing of the target on April 7 and used a less synchronized data set. For both sample periods, the target level (2) is closer to the Funabashi figure than the target level (1). At any rate, although no official statements have ever been made, it seems reasonable to conclude that the central banks of Japan and the United States effectively adopted a target zone regime, in the sense that they intervened when the yen was near the upper and the lower edges of a definite band.

2.4. Target Zone Credibility

2.4.1. The Methodology of Estimating the Expected Rate of Devaluation

According to Bertola and Caballero (1992), Bertola and Svensson (1993) and others, the exchange rate would not be stable in a target zone, unless the zone were credible. This issue of target zone credibility can be examined by estimating the expected future exchange rate and the expected rate of devaluation.

Following Bertola and Svensson (1992), we now present a methodology of estimating the expected rate of devaluation as a way of examining target zone credibility. In order to do so, we first assume uncovered interest parity,[11] such that the interest rate differential between the domestic and foreign currency interest rates equal the expected rate of currency depreciation, *i.e.,*

$$i_t - i_t^* = \delta_t = E_t(\Delta s_{t+\tau}), \tag{6}$$

[11] To be sure, uncovered interest parity has been rejected in a variety of empirical tests (see Froot and Thaler (1990)). However, the standard test of whether the forward exchange rate is an unbiased predictor of the future exchange rate may not be appropriate in a target zone context because of the peso problem, which causes a skew in the distribution of forecast errors. Svensson (1992) indicates that the foreign exchange risk premium is likely to be small for exchange rate target zones, and that uncovered interest parity should thus be a good approximation.

where i_t and i_t^* are the domestic and foreign currency interest rates on deposits of the same default risk and maturity (τ >0); δ_t is the interest rate differential at time t; s_t is the logarithm of the spot exchange rate measured in units of the domestic currency per unit of the foreign currency; $E_t(\cdot)$ is the expectations operator conditional on the information available in period t; and Δ is a difference operator.

We express the logarithm of the exchange rate as,

$$s_t \equiv c_t + x_t, \tag{7}$$

where c_t is the logarithm of the central rate and x_t is the logarithm of the deviation from the central rate. Then, x_t can be defined as the exchange rate within the band. It follows that the expected rate of currency depreciation can be written as

$$E_t(\Delta s_{t+\tau}) = E_t(\Delta c_{t+\tau}) + E_t(\Delta x_{t+\tau}), \tag{8}$$

where $E_t(\Delta c_{t+\tau})$ is the expected change in the central rate and $E_t(\Delta x_{t+\tau})$ is the expected change in the exchange rate within the band.

From equations (6) and (8), we have

$$E_t(\Delta c_{t+\tau}) = \delta_t - E_t(\Delta x_{t+\tau}). \tag{9}$$

Here, the expected change in the central rate can be represented as,

$$E_t(\Delta c_{t+\tau}) = p_{t+\tau} E_t(\Delta c_{t+\tau} | realignment) + (1 - p_{t+\tau}) E_t(\Delta c_{t+\tau} | no \quad realignment)$$

$$= p_{t+\tau} E_t(\Delta c_{t+\tau} | realignment), \tag{10}$$

where $p_{t+\tau}$ is the subjective probability that a realignment occurs in period $t+\tau$ and $E_t(\Delta c_{t+\tau}|realignment)$ is the expected change in the central rate, given that a realignment occurs. It can also be shown that,

$$E_t(\Delta x_{t+\tau}) = (1-p_{t+\tau})E_t(\Delta x_{t+\tau}|no \quad realignment) + p_{t+\tau}E_t(\Delta x_{t+\tau}|realignment)$$
$$= E_t(\Delta x_{t+\tau}|no \quad realignment) - p_{t+\tau}\{E_t(\Delta x_{t+\tau}|no \quad realignment) - E_t(\Delta x_{t+\tau}|realignment)\}, \tag{11}$$

where $E_t(\Delta x_{t+\tau}|(no)realignment)$ is the expected change in the exchange rate within the band, given that a realignment does occur (or does not occur). Substituting equations (10) and (11) into equation (9), we obtain,

$$p_{t+\tau}E_t(\Delta c_{t+\tau}|realignment) + p_{t+\tau}\{E_t(\Delta x_{t+\tau}|realignment) - E_t(\Delta x_{t+\tau}|no \quad realignment)\}$$
$$= \delta_t - E_t(\Delta x_{t+\tau}|no \quad realignment). \tag{12}$$

Here, let us define the left hand side of equation (12) as the expected rate of devaluation, with a larger expected rate of devaluation (in absolute value) being associated with lower credibility, and a smaller expected rate of devaluation (in absolute value) with higher credibility.

To estimate the expected rate of devaluation, we must identify the expected change in the exchange rate within the band,[12] $E_t(\Delta x_{t+\tau})$. This can be identified by

[12] Bertola and Svensson (1993) show that the exchange rate within the band displays mean reversion, and that the relationship between the expected change in the exchange rate within the band and the current exchange rate is probably non-linear. Lindberg, Söderlind and Svensson (1993), however, find that a linear approximation is satisfactory, by indicating that a simple linear

estimating the following equation,[13]

$$(x_{t+n} - x_t) = \sum_j \alpha_j d_j + \beta\, x_t + u_{t+n}\,, \tag{13}$$

where α_j and β are coefficients to be estimated; x_{t+n} is the exchange rate at time t+n (n days from time t); u_{t+n} is an error (forecast error) term; d_j is a dummy variable for regime j, *i.e.,* any period during which no realignment occurs; and the term $\sum_j \alpha_j d_j$ is allowed to vary across regimes.

In estimating equation (13), we have two problems. First, error terms are serially correlated because of overlapping observations, with the sampling interval shorter than the forecasting horizon.[14] Second, error terms are likely to be heteroskedastic because their conditional distribution has a non-normal shape. These problems of autocorrelation and heteroskedasticity will be fully addressed in the estimation of equation (13) below, by using the generalized method of moments (GMM) estimation procedure.

2.4.2. The Expected Future Exchange Rate and Target Zone Credibility

Before proceeding to apply this methodology, it may be useful to visually inspect the credibility of the presumed target zones by examining whether the expected future spot exchange rate is inside or outside the exchange rate band (Svensson (1991)). Figures 2.4 and 2.5 show the spot exchange rate and the estimated expected future

regression of the realized change in the exchange rate within the band on the current exchange rate consistently generates sensible results.

[13] Because subjective probability on the left hand side of equation (12) cannot be estimated directly, we estimate the expected rate of devaluation by using the right hand side variables in equation (12), *i.e.,* the interest rate differential δ_t and the expected change in the exchange rate

spot exchange rate over the 1, 3, and 12 month horizons for the Japanese yen and the deutsche mark, respectively. From these figures, we observe that the yen's expected 1-month future spot rate was almost entirely inside the presumed band, but that the expected 12-month future spot rate was frequently outside the band. The yen's expected 12-month future spot rate was largely outside the band from the latter half of March to May, when the yen was appreciating. On the other hand, we see that the deutsche mark's expected 1-month and 3-month future spot rates were both inside the band during the period, although the expected 12-month future spot rate was outside the band during a few days in May.

It can thus be concluded that the target zones following the Louvre Accord were not perfectly credible. Moreover, the target zone for the deutsche mark was more credible than that for the yen, to the extent that the yen's expected 12-month future spot rate was frequently outside the band, but the deutsche mark's expected 12-month future spot rate was hardly outside the band.

2.4.3. The Expected Rate of Devaluation for the Yen and the Deutsche Mark

Let us examine whether the credibility of the post-Louvre target zones, in which neither the target levels nor the bands were announced, by estimating the expected rate of devaluation as explained in Section 2.4.1. As it is said that the target level for the yen was rebased to 146 yen per dollar in the G5 meetings held in Washington D.C. on April 7, we set regime I as the period from February 23 to April 6, 1987 and regime II as the period from April 8 to October 18,1987.

Here, we attempt to estimate the expected 1-month change in the exchange rate within the band. Because one month consists of approximately 22 trading days, we set

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within the band $E_t(\Delta x_{t+\tau})$. The latter variable is identified by estimating equation (13).

n equal to 22. Table 2.3 shows the result from estimating equation (13), by the GMM method.[15] We find that the coefficient of x_t (β) is -1,08755 for the yen, and -1,07483 for the deutsche mark, both being significantly less than zero, indicating the presence of mean reversion in the exchange rate within the band. Next, we test the hypothesis that the coefficient of x_t (the deviation from the central rate) equals zero, by the augmented Dickey-Fuller unit root test.[16] It turns out that the test statistic is -4.802955 for the yen (with the p-value of 0.0005) and -4.163183 for the deutsche mark (with p-value of 0.0051), leading us also to reject the hypothesis at the 5 percent significance level. This indicates the presence of mean reversion in the exchange rate within the band.

Figure 2.6 shows the expected rates of devaluation[17] (*i.e.,* the difference between the interest rate differential and the expected exchange rate change within the band), for the yen and the deutsche mark, which can be characterized as follows. First, the expected rates of devaluation fluctuated throughout the sample period. Second, they were considerably more volatile than the interest rate differentials.[18] Third, the expected rate of devaluation for the yen was highly negative (*i.e.,* a revaluation was expected) before the realignment of April 7, as well as from the latter half of April to May.

Fourth, the expected rate of devaluation for the yen became highly negative when

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[14] See, for example, Hansen and Hodrick (1980).

[15] In the ordinary least squares (OLS) estimation of equation (13), the Durbin-Watson statistic for serial correlation was 0.16 for the yen and 0.12 for the deutsche mark. The White statistic for heteroskedasticity was 48.6608 for the yen (with the p-value of 0.000) and 10.7755 for the deutsche mark (with the p-value of 0.005). Clearly, the error terms were serially correlated and heteroskedastic.

[16] Because the error terms were serially correlated in equation (13), we adopted the augmented Dickey-Fuller test.

[17] Although the expected rate of devaluation is represented as an annual percent rate in Bertola and Svensson (1993) and Svensson (1993), it is represented as a monthly percent rate in this paper.

[18] The variance of the expected rate of devaluation was 8.363 for the yen and 2.390 for the mark, while the variance of the interest rate differential in the Euro market was 0.277 and 0.179,

the yen was close to the lower edge of the band, but became highly positive when the yen was close to the upper edge. Fifth, we can reject the hypothesis that the expected rate of devaluation for the yen was zero during most of the sample period. Sixth and finally, the expected rate of devaluation for the deutsche mark was less volatile than that for the yen, and we cannot reject the hypothesis that it was zero during some months of 1987. We may, therefore, reasonably conclude that target zone for the yen-dollar exchange rate was not credible, although the target zone for the deutsche mark-dollar exchange rate was comparatively more so.

The lack of credibility can be further confirmed by observing the distribution of the exchange rate (Figure 2.7). We observe that for both the yen and the deutsche mark exchange rates the distribution was not U-shaped but hump-shaped. Moreover, Table 2.4 shows the result of estimating a simple linear regression of the interest rate differential on the exchange rate within the band. From this table, we can see that, for both exchange rates, the correlation between the exchange rate and the interest rate differential was not negative but positive. Hence, it can be concluded that either target zone was not perfectly credible.

2.5. Summary and Concluding Remarks

In this chapter, we have presented an empirical analysis of central bank intervention during the 10-month period following the Louvre Accord of February 1987 in order to determine whether or not a target zone might have existed, as is commonly suggested in the popular literature. We have first examined whether the Bank of Japan and the Federal Reserve adopted a target zone regime during the period, by

respectively.

using daily foreign exchange intervention data. It was shown that, at least operationally, the central banks did adopt a target zone during the period, to the extent that they intervened to stabilize the exchange rate at the presumed Louvre target levels and frequently coordinated their intervention operations at the lower edge of the band for the US dollar.

We have then tested the credibility of the target zone by estimating the expected future exchange rate and the expected rate of devaluation. It was shown that the target zone for the yen-dollar exchange rate was less credible than that for the deutsche mark-dollar exchange rate, because (1) the yen's expected future exchange rate was frequently outside the presumed exchange rate band and (2) the expected rate of devaluation was more volatile for the yen-dollar rate than that for the deutsche mark-dollar rate.

The conclusion that the post-Louvre target zone for the yen-dollar rate was not credible may be explained by the fact that neither the levels nor the bands were announced for the exchange rate. On the other hand, the relative credibility of the target zone for the deutsche mark-dollar rate may be explained by the possibility that the market participants may have expected Germany, a participant in the ERM of EMS, to be firmer in defending the target zone.[19] This seems to suggest that, if exchange rate stability can only be earned through credibility, a greater degree of transparency must be attached to a target zone than was accorded to the unannounced arrangement of mutual intervention that apparently existed between the Japanese and US monetary authorities following the Louvre Accord.

[19] In this context, it should be noted that the ERM was stable during the sample period.

Appendix

Daily data on the exchange rates of the US dollar against the Japanese yen and the deutsche mark, as well as the 1, 3, and 12-month Euro-yen, Euro-dollar, and Euro-mark rates, for the period of February 23 (the day after the Louvre Accord) through October 18, 1987 (the day before the stock market crash) were obtained from the *Nihon Keizai Shinbun.* The exchange rates were the closing prices of the yen-dollar rate in New York (at 4:30 a.m. EST, or at 6:30 a.m. Japan Time) and the yen-dollar rate at 7:00 a.m. EST (or 9:00 p.m. Japan Time) in Section 2.3. In Section 2.4, the reported exchange rates were the closing prices in London. Information on daily foreign exchange intervention was obtained from the newspaper accounts of the *Nihon Keizai Shinbun.* From this information, a dummy variable was constructed, with no intervention, dollar-selling intervention, and dollar-purchasing intervention taking the values of 0, -1 and 1, respectively.

Svensson (1993) argues that the ERM target zone was almost perfectly credible from 1987 to 1990.

Table 2.1.

Estimation of the Multinomial Logit Model of Central Bank Intervention [a, b]

	Bank of Japan [c]	Federal Reserve	Coordinated [d]	Combined [e]
(1) sample period I (February 23- October18,1987)				
Coefficients				
α_0	20.5937***	46.1122***	10.7241	44.0785**
	(2.93521)	(2.53579)	(1.21155)	(2.38725)
α_1	-0.150349***	-0.289975**	-0.088412#	-0.277223**
	(-3.09581)	(-2.40252)	(-1.44538)	(-2.26091)
β_0		61.0151***		62.0643***
		(3.05552)		(3.20408)
β_1		-0.405859***		-0.408557***
		(-3.0384)		(-3.16282)
Number of observations	168	169	168	168
Logit slope derivatives [f] (x 100, to convert into percentages)				
$dP(I_t = 1) / dS_{BO}$	-2.4657	-1.3095	-0.83476	-2.4355
$dP(I_t = -1) / dS_{BO}$		1.1126		1.1130
(2) sample period II (April 7- October 18, 1987)				
α_0	33.5555***	72.5194***	30.3325**	67.0835**
	(3.12536)	(2.63135)	(1.98428)	(2.52505)
α_1	-0.242847***	-0.469715***	-0.227634**	-0.433965**
	(-3.23455)	(-2.55912)	(-2.12367)	(-2.45102)
β_0		115.252***		99.1351***
		(3.57532)		(3.47934)
β_1		-0.782966***		-0.665486***
		(-3.59263)		(-3.48181)
Number of observations	138	139	138	138
Logit slope derivatives [f] (x 100, to convert into percentages)				
$dP(I_t = 1) / dS_{BO}$	-3.5603	-2.6449	-1.7332	-3.6729
$dP(I_t = -1) / dS_{BO}$		1.7179		1.6509

(Notes)

[a] Estimation of equations (2-a) and (2-b) in the text; ***, **, * and # indicate that the statistic is significant at the 1, 5, 10, and 15 percent levels. The numbers in parentheses are t-statistics, which are calculated by using the standard errors computed from analytic second derivatives (Newton's method).

[b] We estimate the multinomial logit model by using the maximum likelihood method.

[c] Equation (2-b) cannot be estimated because no dollar-selling intervention is observed for the Bank of Japan. Instead, the following equation is estimated,

$$\ln\left(P(I_t = 1) / P(I_t = 0)\right) = \alpha_0 + \alpha_1 S_{BO},$$

where $P(I_t = 1) \equiv \dfrac{\exp(\alpha_0 + \alpha_1 S_{BO})}{1 + \exp(\alpha_0 + \alpha_1 S_{BO})}$, and $P(I_t = 0) \equiv \dfrac{1}{1 + \exp(\alpha_0 + \alpha_1 S_{BO})}$.

[d] Coordinated intervention is defined as an operation in which both of the central banks intervene in the same direction on the same day.

[e] Combined intervention is defined as an operation in which either the Bank of Japan or the Federal Reserve intervenes.

[f] We report the effects of one-unit change in the regressors on the probability of intervention (also expressed in percentage points), evaluated at the mean of the data.

Table 2.2.
The Estimation of the Yen / Dollar Target Levels [a]

	Federal Reserve	Combined (Bank of Japan or Federal Reserve)
(1) sample period I (February 23- October18,1987)		
Target level (1)	150.34	151.91
Target level (2)	148.07	150.08
Lewis (1995) [b]	147.0	150.4
The target level reported by Funabashi(1989) [c]	149.75	149.75
The average exchange rate	146.00	146.00
(1) sample period II (April 7- October 18, 1987)		
Target level 1	147.20	148.97
Target level 2	146.68	148.02
The target level reported by Funabashi(1989)	146	146
The average exchange rate	144.80	144.80

(Notes)

[a] The target level for the Bank of Japan cannot be estimated because no dollar-selling intervention is observed for the Bank of Japan.

[b] Lewis (1995) uses the second definition of the target level for the entire period of February 22 to October 18, 1987 only, with data obtained from the *New York Times*, the *Wall Street Journal*, and the *Financial Times*.

[c] According to Funabashi (1989), the central rate following the Louvre Accord was initially 153.50 yen per dollar but was rebased to 146 yen per dollar at the G5 meeting in Washington D.C. on April 7. 149.75 reported here is the average of the two.

Table 2.3.
Estimation of the Expected Change in the Exchange Rate within the Band for a 1 Month Horizon, February 23 - October 15, 1987 [a, b]

	estimated coefficients	standard errors [c]	t- statistics
(1) The yen- dollar rate			
α_1	-0.038352***	0.00373526	-10.2676
α_2	-0.00938790*	0.00571644	-1.64226
β	-1.08755***	0.131833	-8.24942

J(q-k) test [d]
test statistic: 2.00952, degree of freedom: 2, p-value: 0.3661.

Augmented Dickey-Fuller test [e]
test statistic: -4.802955, p-value: 0.0005.

(2) The deutsche mark-dollar rate			
α_1	-0.00238820	0.00294583	-0.810707
β	-1.07483 ***	0.161458	-6.65700

J(q-k) test [d]
test statistic: 2 .96804, degree of freedom: 2, p-value: 0.2267.

Augmented Dickey-Fuller test [e]
test statistic: -4.163183, p-value: 0.0051.

(Notes)

[a] Estimation of equation (13) in the text; ***, **, and * indicate that the statistic is significant at the 1, 5, 10 percent levels.

[b] We use the generalized method of moments (GMM) estimators. The instruments applied are the dummy variable for regime j, time trend, interest rate differential, and the exchange rate within the band. Because one month has approximately 22 observations, we set n to be 22.

[c] The standard errors are calculated by adjusting for autocorrelation and heteroscedasticity in error terms. Standard errors are computed from a heteroscedstic-consistent matrix, so that they are robust with respect to autocorrelation.

[d] J (q-k) test is a test of overidentifying restrictions, where q is the number of instruments applied and k is the number of coefficients. The test statistic is distributed asymptotically as a chi-square with q-k degrees of freedom.

[e] A unit root test for the hypothesis that the coefficient of x_t equals zero. On the basis of Schwarz's Bayesian information criterion (SBIC), one lag length was chosen for the yen and the deutsche mark.

Table 2.4.

Regression of the Interest Rate Differential on the Exchange Rate [a, b]

	One- month	Three- month
(1) The yen- dollar rate sample period April 8-October 18, 1987		
Coefficients		
constant	-0.033205***	-0.032989***
	(-12.6875)	(-21.5120)
the exchange rate	0.115407***	0.094090***
	(4.46299)	(3.6143)
AR(1) ρ [c]	0.963703	0.943652
Durbin-Watson statistic	2.02838	2.51355
(2) The deutsche mark-dollar rate sample period February 23- October18,1987		
constant	-0.036303***	-0.035859***
	(-10.0512)	(-13.0675)
the exchange rate	0.093365***	0.063469**
	(4.32081)	(2.43134)
AR(1) ρ [c]	0.981834	0.973443
Durbin-Watson statistic	2.22312	2.57114

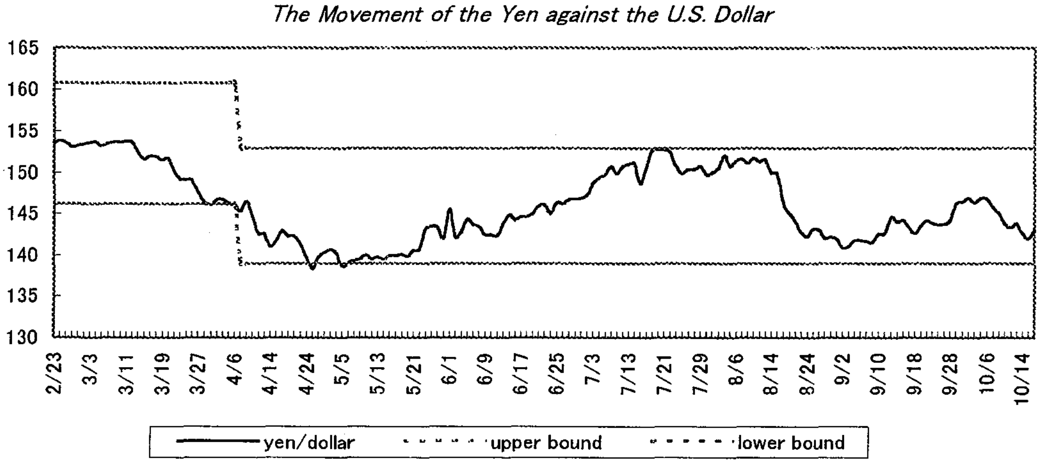
(Notes)

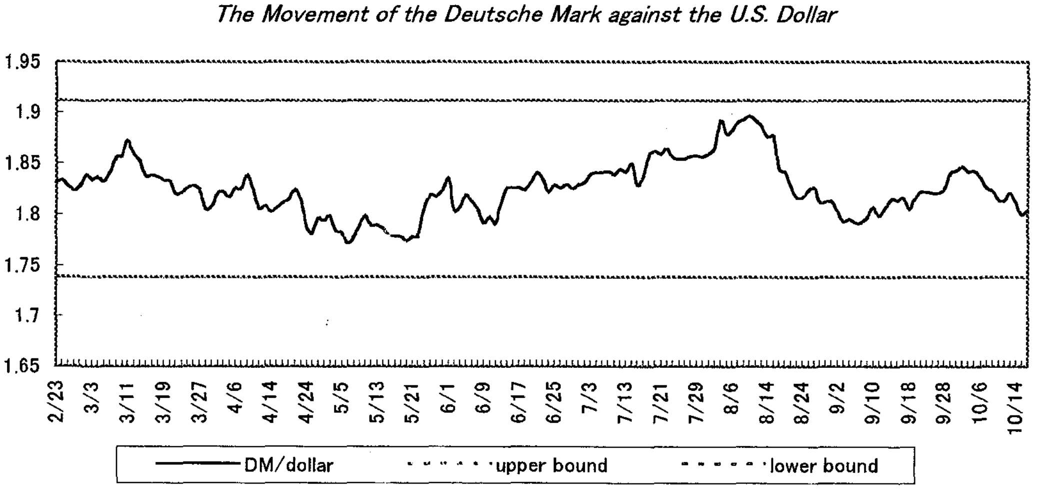
[a] The exchange rate is expressed as a logarithm of its deviation from the central rate. ***, **, and * indicate that the statistic is significant at the 1, 5, 10 percent levels. The numbers in parentheses are t-statistics.

[b] We use the instrumental variable method with a serial correlation correction. The instruments applied are a constant term, the time trend, the interest rate differential, and the exchange rate within the band.

[c] AR(1) is the estimated coefficient of first-order serial correlation in error terms.

Figure 2.1.
The Movement of the Yen and the Deutsche Mark against the US Dollar, February 23 to October 18, 1987

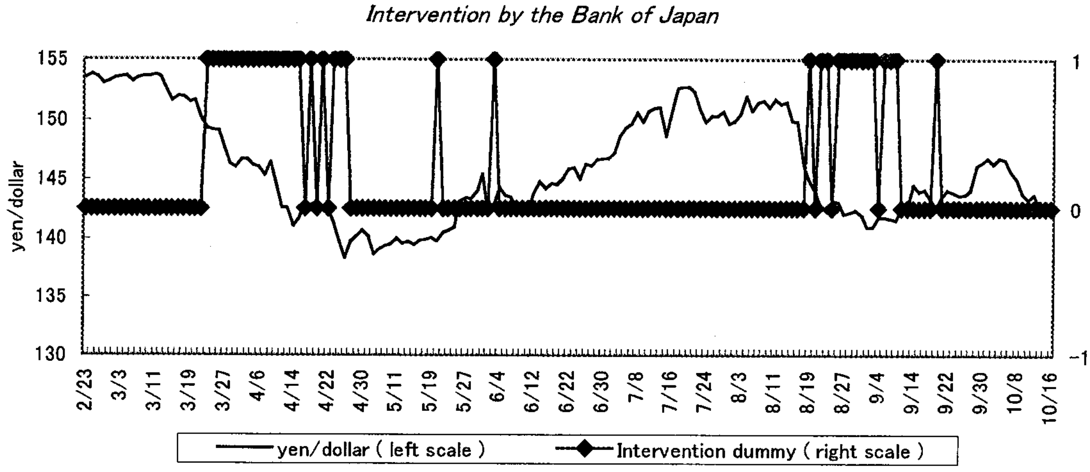




Note: The bands are based on Funabashi (1989).

Sources: The *Nihon Keizai Shinbun* and Funabashi (1989).

Figure 2.2.
Daily Foreign Exchange Intervention by the Bank of Japan and the US Federal Reserve, February 23 to October 18, 1987



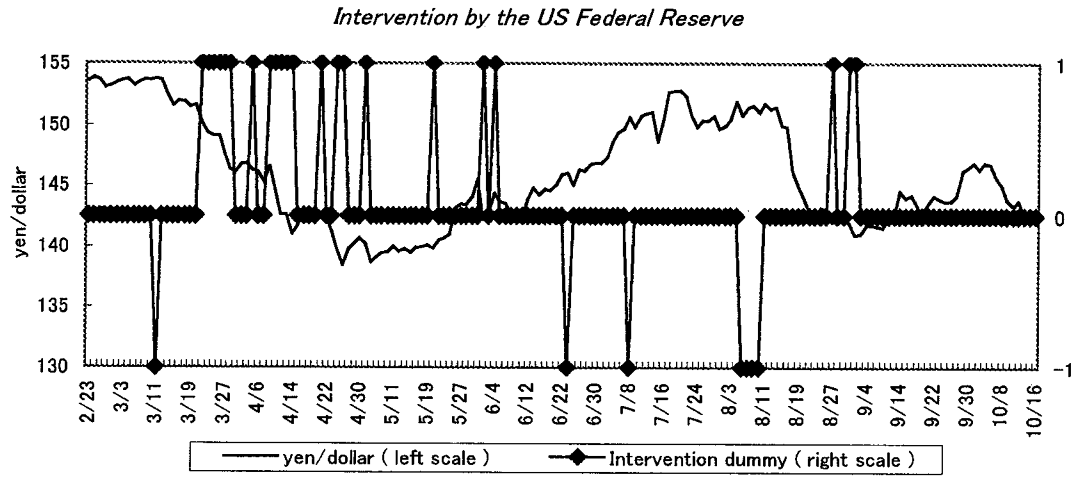
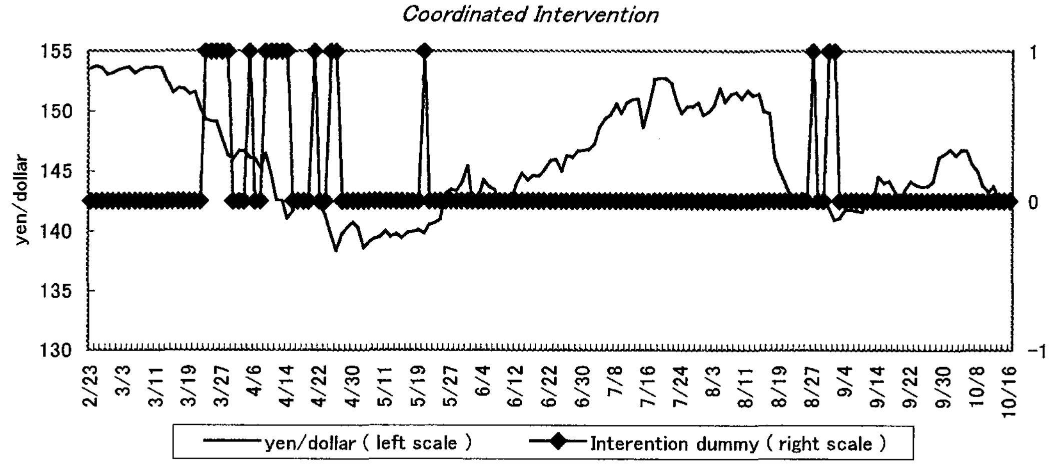
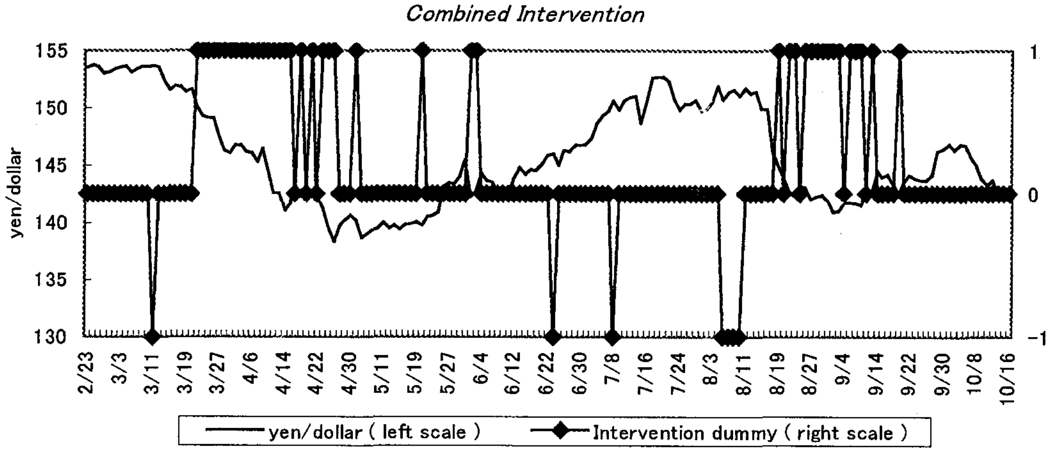


Figure 2.2 (Continued).





Note: Daily foreign exchange intervention data are compiled from the newspaper accounts of the *Nihon Keizai Shinbun*. No intervention, dollar-selling intervention, and dollar-purchasing intervention are defined 0, -1, and 1, respectively. Coordinated intervention is defined as an operation in which both of the central banks intervene in the same direction, while combined intervention is defined as an operation in which either the Bank of Japan or the Federal Reserve intervenes.

Figure 2.3.

The Estimated Probability of Intervention from the Multinomial Logit Model, February 23 to October 18, 1987

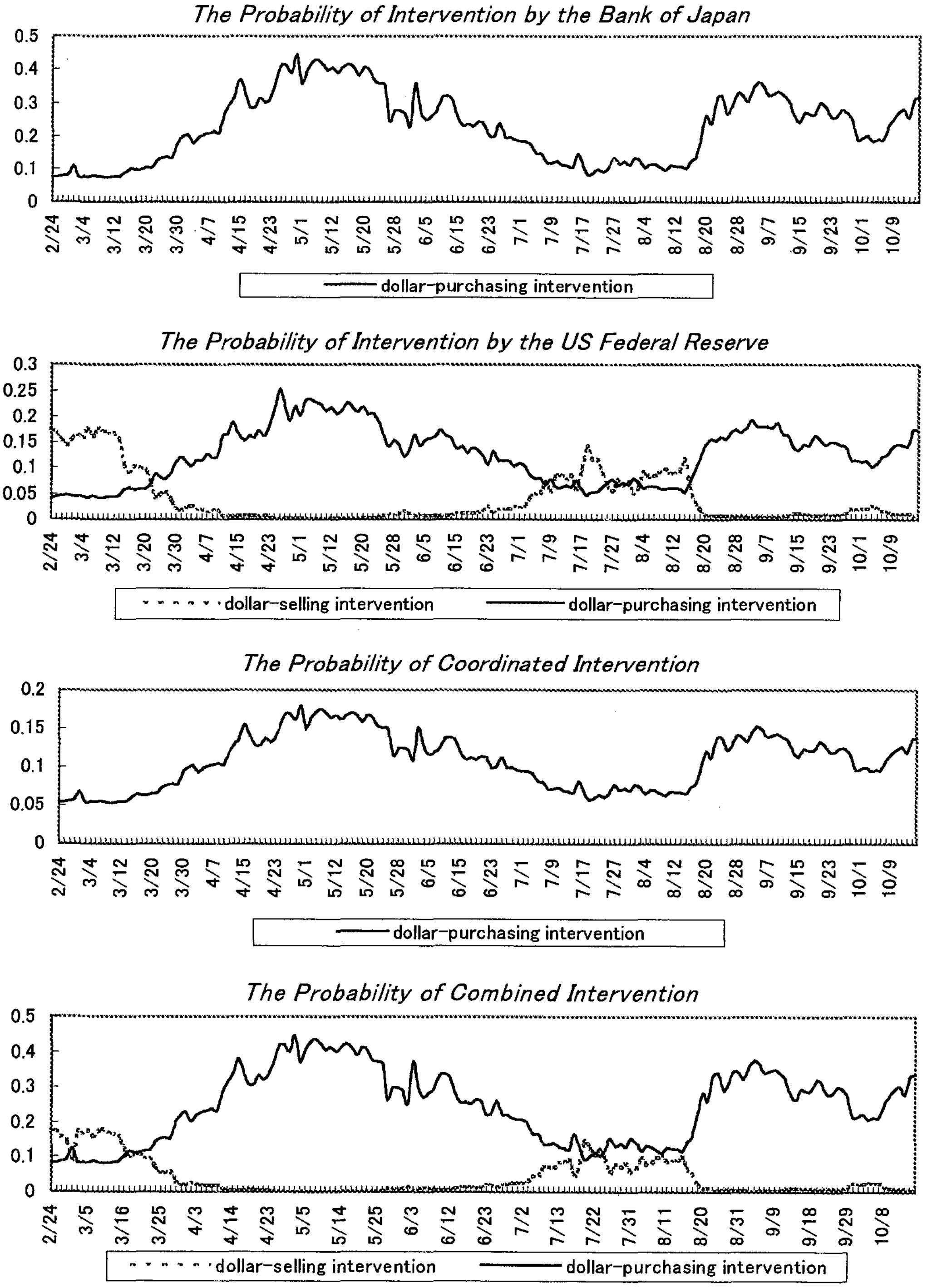
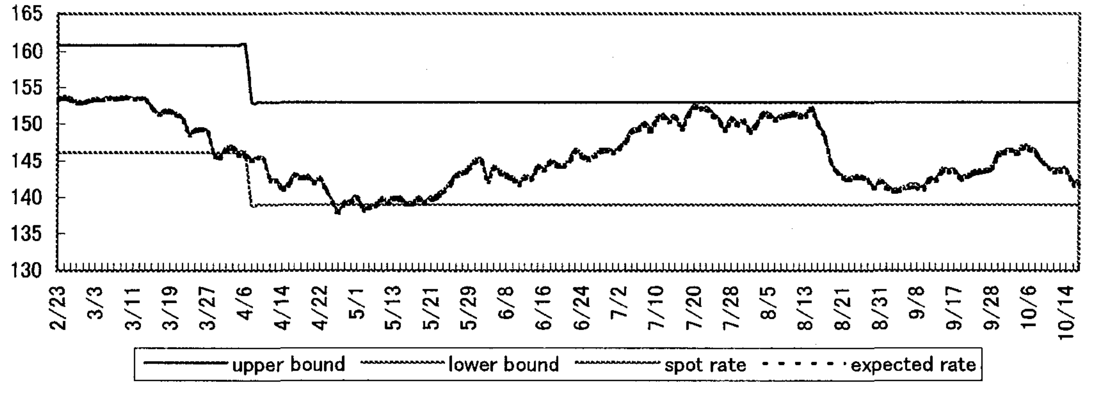
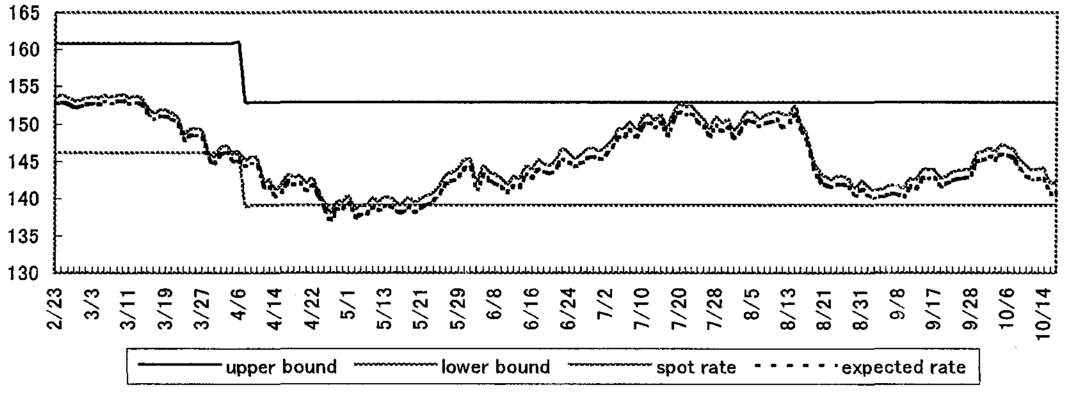


Figure 2.4.
The Yen's Expected 1-month, 3-month and 12-month Future Spot Rate, February 23 to October 18, 1987

The Yen's Expected 1-month Future Spot Rate



The Yen's Expected 3-month Future Spot Rate



The Yen's Expected 12-month Future Spot Rate

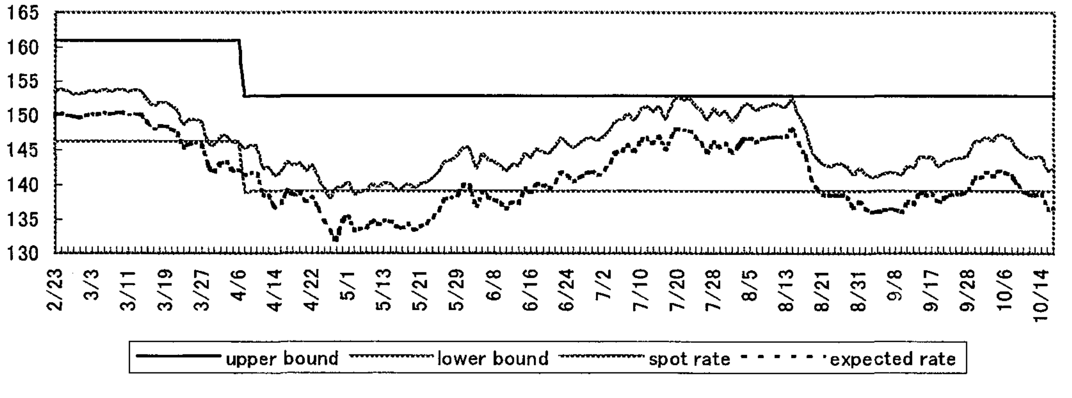
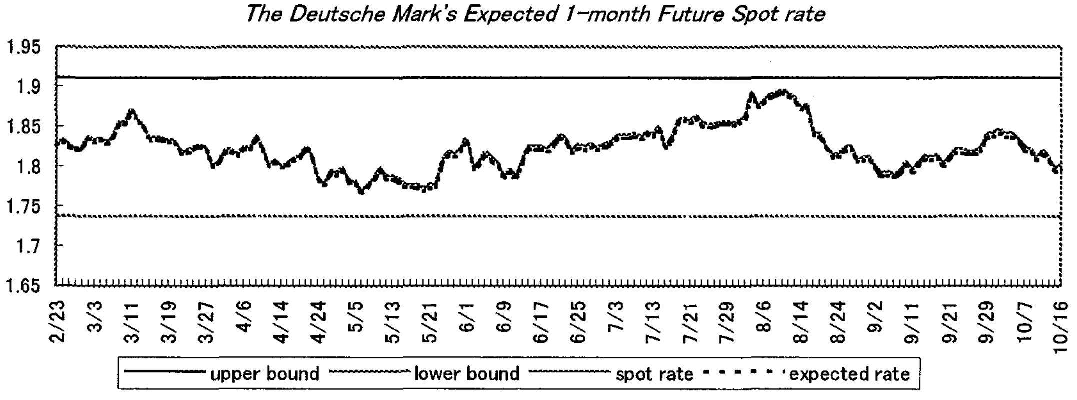
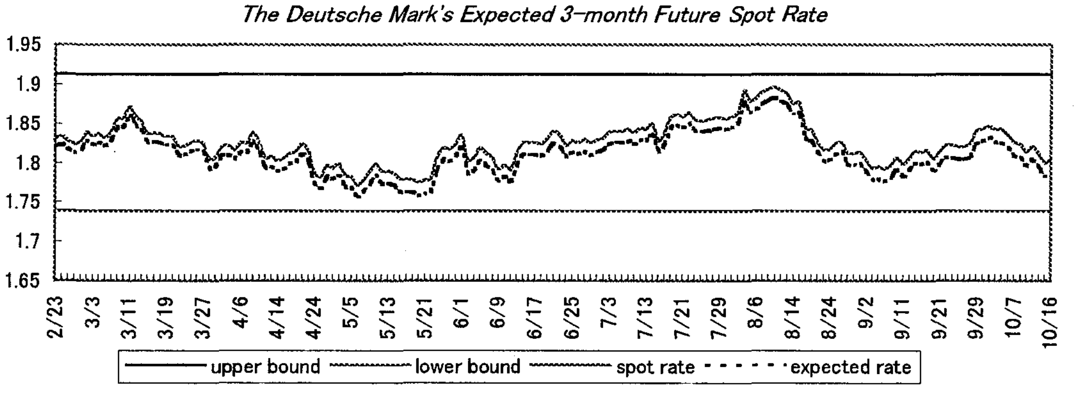


Figure 2.5.
The Deutsche Mark's Expected Future 1-month, 3-month and 12-month Future Spot Rate, February 23 to October 18, 1987





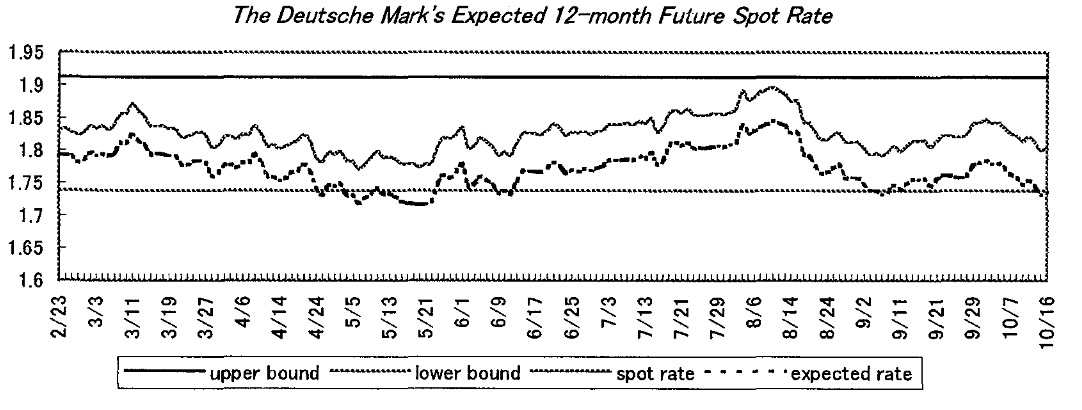
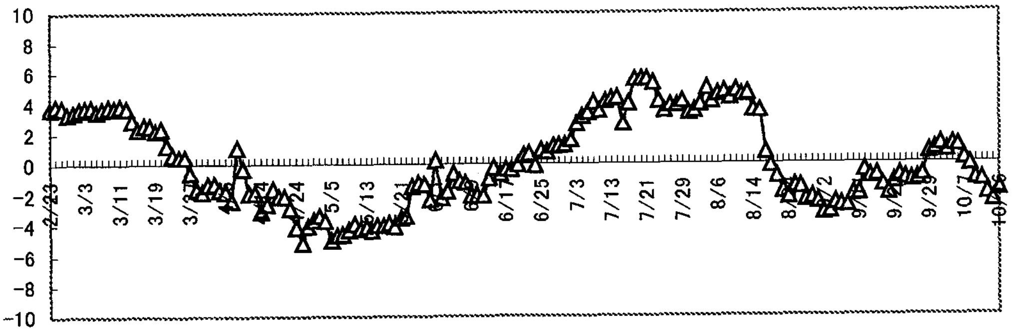


Figure 2.6.

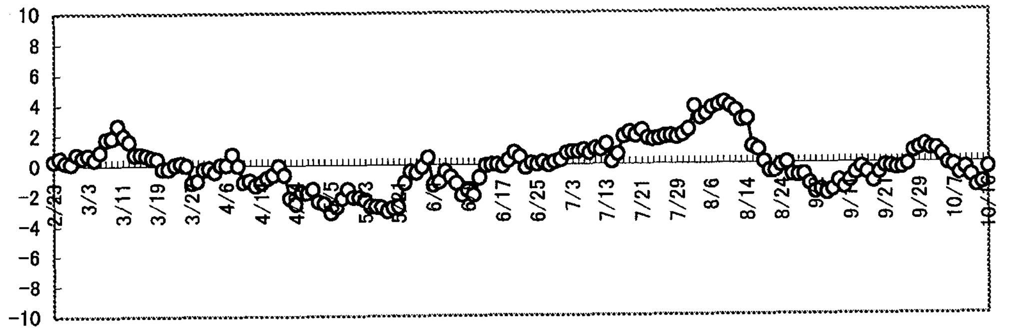
The Expected Rate of Devaluation over the 1 Month Horizon,

February 23 to October 18, 1987

The Expected Rate of Devaluation for the Yen



The Expected Rate of Devaluation for the Deutsche Mark



The Expected Rate of Devaluation for the Yen and the Deutsche Mark

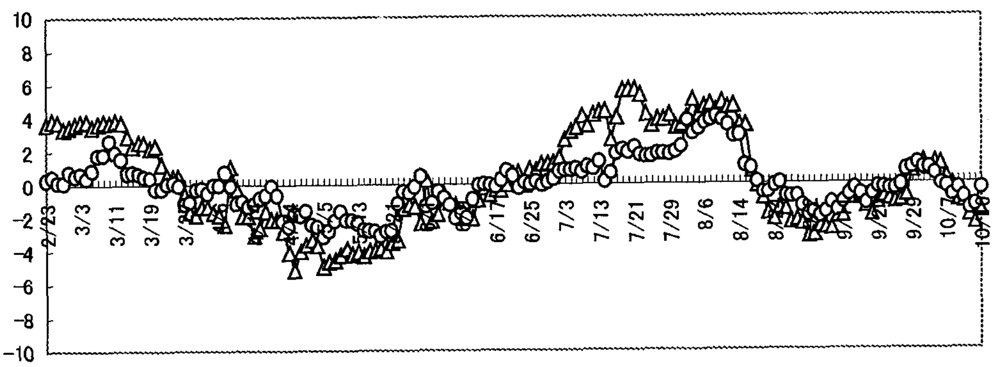
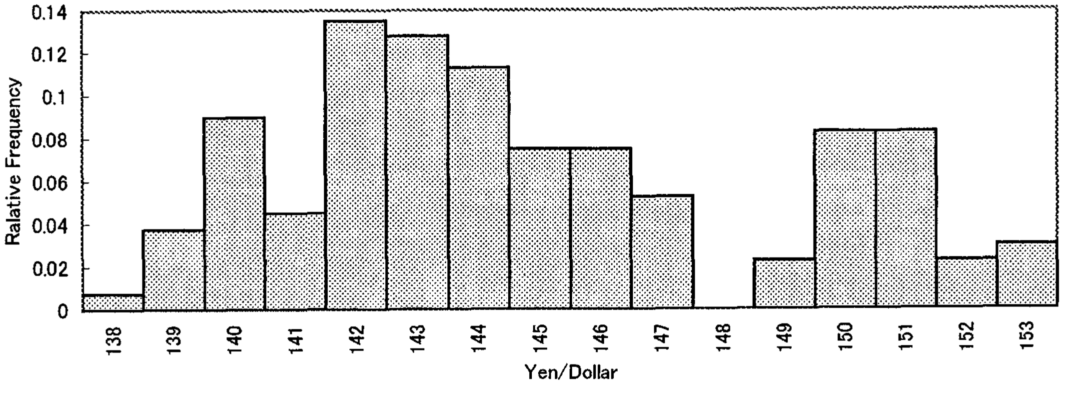


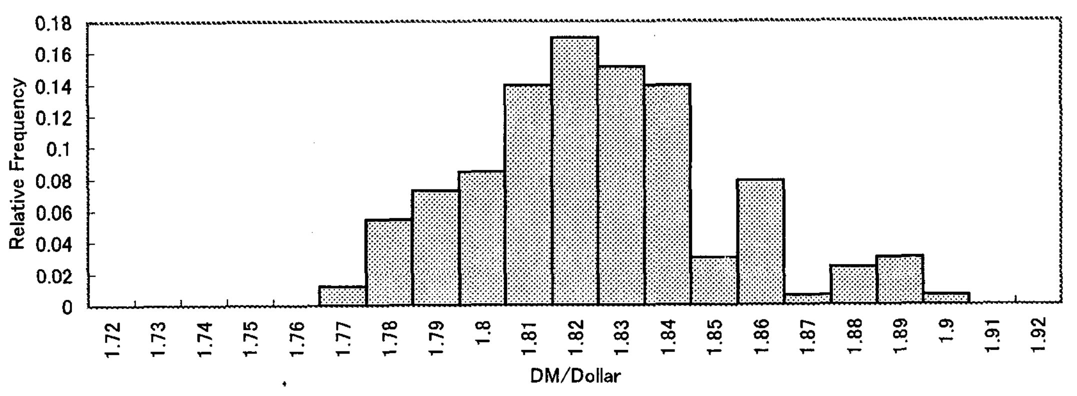
Figure 2.7.

The Frequency Distribution of the Exchange Rate

(1) The Yen-Dollar Rate, April 8-October 18,1987



(2) The Deutsche Mark-Dollar Rate, February 23-October 18,1987



3. Was It Really a Dollar Peg?: The Exchange Rate Policies of East Asian Countries, 1980-97

3.1. Introduction

Subsequent to the dramatic fall of the Thai baht on July 2, 1997, other East Asian currencies were also subjected to speculative pressure and the authorities were forced to allow them to depreciate sharply. In the context of this Asian currency crisis of 1997, it is often claimed that one of the major ingredients of the environment leading to the crisis was the exchange rate policies of East Asian countries, whereby their currencies were effectively pegged to the US dollar (see Ito, Ogawa and Sasaki (1998)). It is said that the presumed dollar peg contributed to excessive capital inflows by minimizing the exchange rate risk for international investors (see Radelet and Sachs (1998a, 1998b)). If the currencies had been freely floating, the argument goes, the currency crisis might well have been prevented.

This common argument linking the supposed dollar peg with the currency crisis of 1997, however, is curious for at least two reasons. First, at least officially, all of the East Asian countries or regions, except Hong Kong, had claimed to have a relatively flexible exchange rate policy during the period of at least 10 years leading up to the currency crisis. For example, according to the classification system of the International Monetary Fund (IMF), Thailand had a basket peg, Korea, Indonesia, Malaysia and Singapore had a managed float, and the Philippines even had an independent float. By casually looking at the behavior of many of these currencies, particularly the Indonesian rupiah and the Philippine peso, we find that the US dollar exchange rates did fluctuate fairly substantially over this period.

Second, the pioneering work of Frankel (1992) and Frankel and Wei (1994) indicates that the implicit weight of the dollar was large, but not unity, in the determination of the nominal values of these currencies. Moreover, the literature also suggests that the weight of the dollar was by no means fixed and that the weight of the yen did seem to increase over time (Kwan (1995)). An important insight provided by Takagi (1996) is that the weight of the Japanese yen tends to become underestimated if it is calculated during "tranquil" times, and that some East Asian monetary authorities systematically moved their currencies in response to sharp movements of the yen-dollar exchange rate. It is thus possible that the exchange rate policies of East Asian countries were too involved to be characterized as a simple dollar peg or possibly even as a de facto dollar peg.

This chapter begins where these earlier studies have left off, by quantifying the changing weights of the US dollar and the Japanese yen in the exchange rate policies of Indonesia, Korea, Malaysia, the Philippines, Singapore and Thailand during 1980-97 on the basis of a time-varying parameter model. First, Section 3.2 will present an overview of the movements of the East Asian currencies against the US dollar and the Japanese yen. Section 3.3 will explain the methodology of estimating the implicit weights of foreign currencies in the nominal values of East Asian currencies. Section 3.4 will estimate the conventional constant weights of foreign currencies in the determination of the nominal values of the East Asian currencies and test for the stability of the estimated coefficients. On the basis of this test, Section 3.5 will present the time-varying weights of the foreign currencies which are obtained by the Kalman filter technique. Section 3.6 will present a summary and concluding remarks. Finally, the Appendix will present a test of the statistical significance of the variation of the estimated time-varying weights.

3.2. The Movements of East Asian Currencies: An Overview

At the outset, it may be useful to review the movements of the East Asian currencies against the US dollar and the Japanese yen during 1980-97, against the general background of the substantial appreciation of the yen against the dollar. As Figure 3.1 indicates, the yen appreciated markedly against the dollar over the period, moving from around 250 yen per dollar to the range of 120 yen. Particularly, following the Plaza Agreement of the Group of Five (G5) countries in September 1985, the yen appreciated sharply from 240 yen per dollar to 120 yen in November 1988. Again, the yen appreciated sharply from 125 yen per dollar in January 1993 to 83 yen in May 1995. The yen, however, depreciated from 83 yen in May 1995 to 125 yen in April 1997.

Against this general background, Figure 3.2 shows the movements of the East Asian currencies against the US dollar and the Japanese yen from January 1980 to June 1997. It is clear from this figure that, in each of the countries considered, the US dollar exchange rate was less volatile than the yen exchange rate, evidently indicating that the East Asian countries were stabilizing their currencies more in term of the US dollar.

Looking at individual currencies separately, the Thai baht was pegged to the US dollar at 23 baht per dollar from July 1981 to November 1984, when Thailand officially began to peg its currency to an undisclosed basket of Thailand's major trading partners. The baht was stable at around 26 baht per dollar from November 1984 to June 1997, while depreciating against the Japanese yen over the entire period, except during the last two years (2a).

The Korean won was linked to a basket of major currencies from February 1980 to March 1990, when the authorities adopted a "market average rate (MAR) system", in which market factors were allowed to play a role in the determination of the won-dollar exchange rate within a 2.25 percent daily limit (Black (1996)). The won-dollar exchange rate fluctuated between 580 and 890 won per dollar, while the won depreciated considerably against the Japanese yen over the entire period (2b).

The Singapore dollar was managed by the Monetary Authority of Singapore (MAS) according to a trade-weighted basket of foreign currencies, which might be changed from time to time. The Singapore dollar appreciated against the US dollar from February 1985 to June 1997, while remaining fairly stable against the Japanese yen (2c).

From September 1975, the Malaysian ringgit was pegged to a basket of major currencies by Bank Negara Malaysia. The ringgit fluctuated between 2.1 and 2.8 ringgit per dollar, while depreciating against the Japanese yen over the entire period, except during the last two years (2d).

The Indonesian rupiah was linked to the US dollar until November 1978, when the currency began to be pegged a basket of major currencies subject to periodic devaluations in a crawling peg fashion (2e). Over the period under consideration, there were substantial devaluations twice, namely, in April 1983 and September 1986. Not surprisingly, the rupiah depreciated against both the dollar and the yen throughout the period.

Finally, the Philippine peso was under a managed float until October 1984, when an independent float was officially introduced. The peso depreciated against the US dollar over this period, suggesting that, in practice, the regime was a crawling peg to the US dollar. As evidenced by the relative short-run stability against the dollar, it is said that the central bank heavily intervened to control the peso's fluctuations against that currency in the short run(2f).

3.3. The Methodology of Estimating the Implicit Weights of Foreign Currencies

Following the convention of the literature, we will characterize the exchange rate policies of East Asian countries by estimating the implicit weights of major currencies in the determination of the nominal values of their currencies. Unlike the existing literature, however, we will use a time-varying parameter model by using the Kalman filter technique. In what follows, we will first estimate the implicit weights by using the usual constant parameter model and then, after testing for the stability of the estimated coefficients, we will present the time-varying weights. Before proceeding further, we will explain the two alternative models and describe the data set below.

3.3.1. The Constant Parameter Model

The constant weights can be estimated by the ordinary least squares (OLS) method. With the Swiss franc chosen as the numeraire (as in Frankel and Wei (1994)), we estimate the following regression equation for each of the six East Asian currencies,

$$\Delta \ln DEX(t) = \beta_0 + \beta_1 \Delta \ln JYEN(t) + \beta_2 \Delta \ln USD(t) + \beta_3 \Delta \ln DM(t) + \varepsilon(t), \tag{1}$$

where Δ is a first difference operator (*i.e.,* $\Delta \ln USD(t) = \ln USD(t) - \ln USD(t-1)$);

$DEX(t)$ is the exchange rate against the Swiss franc at time t; $JYEN(t)$, $USD(t)$ and $DM(t)$ are, respectively, the exchange rates of the Japanese yen, the US dollar and the deutsche mark against the Swiss franc at time t; β_0, β_1, β_2 and β_3 are the coefficients to be estimated, with β_1, β_2 and β_3 being the implicit weights of the Japanese yen, US dollar and deutsche mark; and $\varepsilon(t)$ is an error term.

3.3.2. The Time-Varying Parameter Model

How the authorities of East Asian countries might have altered the weights of foreign currencies over time can be examined by using a time-varying parameter model. We estimate the following regression equation by using the Kalman filter technique,

$$\Delta \ln DEX(t) = \beta_0(t) + \beta_1(t)\Delta \ln JYEN(t) + \beta_2(t)\Delta \ln USD(t) + \beta_3(t)\Delta \ln DM(t) + \varepsilon(t), \qquad (2)$$

$$\beta_j(t) = \beta_j(t-1) + \eta_j(t), \quad j = 0,\ldots,3, \qquad (3)$$

where $\varepsilon(t)$ is an independently and normally distributed error term with zero mean and a constant variance σ^2, and $\eta_j(t)$ are random error terms with zero mean and variances υ_j^2, assumed independent across all j and t, and of $\varepsilon(t)$. This model assumes that $\beta_j(t)$ follows a random walk process, such that changes in the parameters are randomly driven by disturbances $\eta_j(t)$. The reasoning underlying this formulation is that changes in $\beta_j(t)$ reflect the random arrival of new information. The profiles of $\beta_j(t)$, therefore, evolve over time according to equation (3).

In *state-space form*, equation (1) can be rewritten as a *measurement equation* (2) and the parameter profiles given by equation (3) can be expressed as a *transition equation*. This model is a systematically time-varying parameter model[1] that can be estimated by the maximum likelihood method through the Kalman filter technique (see, for example, Harvey (1981), Harvey (1989) and Hamilton (1994)). The Kalman filter recursively updates the estimate of $\beta_j(t)$, by using the new information in $\Delta \ln DEX(t)$, $\Delta \ln JYEN(t)$, $\Delta \ln USD(t)$ and $\Delta \ln DM(t)$ for each observation. In this sense, it can be viewed as a Bayesian method.[2]

3.3.3. The Data Set

We use monthly (end-of-month) data from January 1980 up to June 1997 (the month preceding the Asian currency crisis). The exchange rate series involve the six East Asian currencies (*i.e.,* the Thai baht, the Korean won, the Singapore dollar, the Malaysian ringgit, the Indonesian rupiah and the Philippines peso) as well as three major currencies (*i.e.,* the Japanese yen, the US dollar and the deutsche mark). The exchange rates are expressed in terms of the Swiss franc, by dividing the US dollar exchange rates (as reported in the International Monetary Fund, *International Financial Statistics* (IFS), various issues) by the Swiss francs-US dollar rate.

3.4. The Stability of Estimated Fixed Coefficients

In this section, we will estimate the implicit weights of foreign currencies by OLS, assuming that the weights are constant. We will then test for the stability of the

[1] Because the parameters in this setup are allowed to follow a random walk, this specification is much less restrictive than other types of time-varying parameter models.

[2] The Kalman filter may be viewed as mimicking a sequential optimal learning process. The predictions are rational in the sense that the agent optimally utilizes current and past information when learning about his stochastic environment (see Harvey (1981) and Hamilton (1994)).

estimated coefficients.

3.4.1. Estimating the Fixed Coefficients

Table 3.1 reports the estimated fixed implicit weights of major foreign currencies, for the entire sample period (January 1980 to June 1997) as well as for two subperiods (January 1980 to December 1989, and January 1990 to June 1997). First, in the determination of the nominal value of the Thai baht, the weights of the Japanese yen (9.2 percent) and the US dollar (82.2 percent) were statistically significant for the entire sample period. Second, for the Korean won, the weights of the US dollar (92.8 percent) and the Japanese yen (5.0 percent) were likewise statistically significant.

Third, in the determination of the nominal value of the Singapore dollar, the weights of the Japanese yen (12.0 percent), the US dollar (72.6 percent) and the deutsche mark (8.2 percent) were all statistically significant over the entire sample period. Fourth, for the Malaysian ringgit, the weights of the Japanese yen (5.9 percent), US dollar (80.7 percent) and deutsche mark (9.3 percent) were all likewise significant. Finally, for the Indonesia rupiah and the Philippines peso, while the weight of the US dollar was large and significant (97.9 percent for the rupiah and 106 percent for the peso), the weights of the other currencies were not significant, indicating that they were closely linked to the US dollar.

3.4.2. Testing for the Stability of Estimated Coefficients

Here, we apply the CUSUMSQ test of Brown, Durbin and Evans (1975) to see whether or not the estimated coefficients (β_j) were stable (Table 3.2). For the entire sample period, as the CUSUMSQ plot of the recursive residuals obtained from equation (1) violates the 5 percent line,[3] we are able to reject the null hypothesis of parameter stability at the 5 percent level for all six currencies. However, we are not able to reject the null hypothesis (at the 10 percent level) in the cases of the Thai baht and the Singapore dollar during the subperiod of 1990-97.

As an additional test for the Thai baht and the Singapore dollar during that subperiod, we apply the Chow test for these currencies only (Figure 3.3). Figure 3.3 shows the p-values (*i.e.,* the probability that the F statistic of the chow test does not exceed the critical value) for the Thai baht and the Singapore dollar, respectively. It should be noted that the null hypothesis of parameter stability is rejected if the p-value is less than a chosen significance level. From these figures, on the basis of the Chow test, we are also able to reject the null hypothesis at the 5 percent level for both currencies during the subperiod of 1990-97.

In estimating equation (1) by OLS, it is implicitly assumed that the estimated weight (β_j) is constant. However, the CUSUMSQ test and the Chow test indicate that the weights were not constant but that they varied over time. This means that the OLS weights are biased and provide an inaccurate picture of the exchange rate policies of East Asian countries. A time-varying parameter model is called for.

3.5. Estimating the Time-Varying Weights

In this section, in the light of the evidence that the implicit weights of the foreign currencies were apparently time-varying, we estimate equations (2) and (3) by the Kalman filter technique. Given the insight provided by Takagi (1996), we are

[3] Although the CUSUMSQ test may fail to detect coefficient variation because of lack of power, it is almost certain that the estimated weights of foreign currencies changed during the sample period.

particularly interested in quantifying in a more rigorous manner how the East Asian countries shifted the weight of the Japanese yen in response to the sharp movement of the yen-dollar exchange rate. As the starting values of β required in initializing the Kalman filter estimation, we follow the convention of using the estimates from OLS.[4]

Figures 3.4-3.9 show the estimated time-varying weights of the three major currencies (*i.e.,* β_1, β_2 and β_3) in the exchange rate policies of Thailand, Korea, Singapore, Malaysia, Indonesia and the Philippines. These figures indicate that the estimated weights did vary over time, in a manner consistent with the result of the earlier stability tests. On the whole, the weight of the US dollar was large, ranging between 60 and 130 percent, while the weight of the Japanese yen was relatively small, ranging between zero and 25 percent (ignoring the negative values). However, it should be noted that the weight of the Japanese yen varied substantially over the entire sample period,[5] with the weight generally higher in the 1990s than in the 1980s.

The Thai Baht

Looking at individual currencies separately, from Figure 3.4, we observe that the implicit weight of the Japanese yen in the determination of the Thai baht was stable at around 6 percent during the first half of the 1980s. Between January and August of 1985, the weight briefly fell to 0.5 percent. Following the Plaza Agreement of September 1985, however, it began to increase gradually and continued to do so through the rest of the 1980s, reaching 11 percent in June 1991.[6] From June 1991 to

[4] We also estimated the time-varying parameter model by using alternative starting values, and found that the coefficients consistently converged to the same maximum values.

[5] Following McNelis and Neftçi (1982), we tested if the individual coefficients varied over time, and found that the weights of the Japanese yen and the US dollar changed substantially during the sample period for all six currencies (see the Appendix and Appendix Figures A3.1-A3.6).

[6] From September 1985 to June 1991, the variation in the estimated weight of yen was

June 1997, the weight of the yen was stable at around 11 percent. On the other hand, the weight of the US dollar ranged between 77 and 86 percent for the entire period, while was stable around 82 percent from 1987 to 1997. Given the stability of the yen and dollar weights, it appears that the baht was pegged strictly to a basket of currencies, including the yen and the dollar.

The Korean Won

For the Korean won, we observe from Figure 3.5 that the implicit weight of the Japanese yen fluctuated around 2 percent during the period of 1980-94. However, from May 1995 to April 1997, the weight rose sharply from 2 percent to 14 percent, as the yen depreciated from 83 yen per dollar to 125 yen,[7] suggesting that the Korean authorities significantly raised the weight of the yen, possibly to maintain Korea's export competitiveness (Takagi (1996)). On the other hand, the weight of the US dollar varied substantially, ranging between 90 and 98 percent for the entire period. Although the Korean won was officially linked to a basket of currencies from February 1980 to February 1990, it was thus effectively pegged to the US dollar.

The Singapore Dollar

For the Singapore dollar, we observe from Figure 3.6 that the implicit weight of the Japanese yen jumped from 8 to 17 percent during the two months following the Plaza Agreement of September 1985.[8] The weight of yen was stable at around 18 percent from 1991 to 1992, when the yen-dollar exchange rate was stable. When the

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significant, because the test statistic of the yen's weight rose from -2.5 to 0.5 (Figure A3.1).

[7] From May 1995 to April 1997, the variation in the estimated weight of yen was significant, because the test statistic of the yen's weight rose from 0.5 to 4 (Figure A3.2).

[8] During the two months following the Plaza Agreement, the variation in the estimated weight of yen was significant, because the test statistic of the yen's weight jumped from -1.5 to 2 (Figure A3.3).

Japanese yen depreciated against the US dollar from 83 yen in May 1995 to 125 yen in April 1997, the weight of the yen did not rise but it was stable at around 15 percent. These movements confirm the official view (supported by Takagi (1996)) that the Singaporean authorities give priority to price stability. Thus, they raised the weight of the yen, when there was a secular appreciation of the yen against the US dollar.

The Malaysian Ringgit

For the Malaysian ringgit, we observe from Figure 3.7 that the weight of the Japanese yen fluctuated between 2 and 6 percent during the 1980s. When the yen was around 110 yen per dollar (in February 1994), its weight was 12 percent but fell to 5 percent when the yen appreciated sharply against the dollar (during June 1994- May 1995). On the other hand, when the yen depreciated from 83 yen per dollar in May 1995 to 125 yen in April 1997, the weight of the yen rose sharply from 5 to 13 percent.[9] These movements indicate that the Malaysian authorities stressed export promotion as an objective of exchange rate policy (Takagi (1996)), hence systematically shifted the weight of the yen whenever the yen moved sharply against the US.

The Indonesia Rupiah

Figure 3.8 shows that, for the Indonesian rupiah, the implicit weight of the Japanese yen was algebraically small throughout the sample period, taking negative values during much of the later years. On the other hand, the weight of the US dollar was between 95 and 114 percent, suggesting that the rupiah, was closely linked to the US dollar. It is interesting, however, that the weight of the yen increased in algebraic value consistently from the devaluation of September 1986.

[9] From May 1995 to April 1997, the variation in the estimated weight of yen was significant,

The Philippine Peso

Finally, for the Philippine peso, we observe from Figure 3.9 that the implicit weight of the US dollar varied around 100 percent, while the weight of the yen was small. There is even no evidence that the yen's weight increased during the later years. This suggests that, although the Philippines officially had a free float, the peso was effectively closely linked to the US dollar.

3.6. Summary and Concluding Remarks

In this chapter, we have examined the exchange rate policies of six East Asian countries during the 17-year period preceding the currency crisis of 1997. We have first estimated the implicit weights of three major foreign currencies in the determination of the nominal values of their currencies by OLS, assuming that the weights were constant. We have then tested for the stability of the estimated weights and found that they were not constant but varied over time. Accordingly, in the final section, we have estimated a time-varying parameter model by the Kalman filter.

The time-varying weights suggest that the authorities of Korea and Malaysia significantly raised the weight of the Japanese, when the yen depreciated sharply against the US dollar (May 1995 - April 1997). On the other hand, the authorities of Singapore significantly raised the weight of the Japanese yen, when the yen appreciated sharply against the US dollar. These observations rigorously support the assessment of Takagi (1996) that the important objectives of exchange rate policies were export promotion in Korea and Malaysia and price stability in Singapore.

because the test statistic of the yen's weight rose from -0.2 to 2.3 (Figure A3.4).

It is true that the weight of the US dollar was large in the exchange rate policies of all countries, confirming the earlier result of Frankel (1992) and Frankel and Wei (1994). In the case of Indonesia and the Philippines, the weight was virtually unity, so that it may rightly be said that the rupiah and the peso were effectively pegged to the US dollar. For the other currencies, however, the weight of the yen did increase in the 1990s. The weight was by no means insignificant in Thailand (almost 12 percent), Korea (almost 14 percent), Singapore (about 14 percent) and Malaysia (around 10 percent) when the currency crisis struck. In the crisis countries of Thailand, Korea and Malaysia, moreover, the weight was steadily rising during the preceding two years along with the depreciation of the yen against the US dollar.[10] Thus, it is incorrect to presume, as has often been done in the past, that the authorities were sitting idly, watching their currencies appreciate against the Japanese yen.

Appendix. Significance Tests of the Variation in Estimated Weights

In the appendix, we test for the significance of the variation in the estimated values of $\beta_j(t)$ obtained by the Kalman filter. Following McNelis and Neftçi (1982), we calculate the test statistics as follows,

$$T_j(t) = \frac{\hat{\beta}_j(t) - \hat{\beta}_j(0)}{\hat{\sigma}(\beta_j(0))}$$

[10] So far, in this chapter, we did not examine the optimal basket weights in East Asia. Ito, Ogawa and Sasaki (1998) show the optimal basket weights to minimize the fluctuation of growth rate of trade balance, and the optimal weights of the yen range between 40 and 80 percent in East Asia. From the standpoint of their research, it may be said that the authorities of Thailand, Korea and Malaysia did not increase the weight of the yen sufficiently before the Asian crisis.

where $\hat{\beta}_j(t)$ is the Kalman filter estimate; $\hat{\beta}_j(0)$ is the starting value for Kalman filter estimation, which in this case is the constant OLS estimate; and $\hat{\sigma}(\beta_j(0))$ is the standard error of the OLS estimate.

The test statistic can be interpreted as a time-varying version of the standard t-test statistic. If the test statistics exceed the critical level, the variation of the estimated weights obtained by the Kalman filter is considered significant (Appendix Figures A3.1- A3.6). Moreover, if the change of the test statistic exceed the critical level ± 1.96, the variation of the estimated weight is significant at the 5 percent level during a period. For example, in the case of the Korean won, the test statistic of the yen's weight rose from -0.5 to 4 (i.e., the change of the test statistic is 4.5) during the period of May 1995 to April 1997 (Appendix Figure A3.2). From these figures, we find that the variations in the estimated weights of the Japanese yen and the US dollar were all significant during the entire sample period.

Table 3.1.
The Fixed Implicit Weights of Three Major Currencies in the Nominal Values of Six East Asian Currencies [a, b, c]

Currency	Period	Constant	Yen	USD	DM	Adj-R^2	DW
Thai baht	80:1-97:6	0.001	0.092***	0.822***	0.021	0.857	2.013
		(1.609)	(2.644)	(27.49)	(0.319)		
	80:1-89:12	0.002	0.074	0.821***	-0.007	0.781	2.015
		(1.417)	(1.155)	(16.66)	(-0.057)		
	90:1-97:6	0.003E-01**	0.107***	0.825***	0.057***	0.998	2.160
		(2.157)	(21.05)	(161.6)	(5.489)		
Korean won	80:1-97:6	0.002***	0.050**	0.928***	0.044	0.946	0.986
		(3.736)	(2.644)	(48.11)	(1.018)		
	80:1-89:12	0.001#	-0.010	0.933***	0.115*	0.944	0.634
		(1.561)	(-0.304)	(36.79)	(1.900)		
	90:1-97:6	0.003***	0.115***	0.923***	-0.063	0.956	1.646
		(4.273)	(4.079)	(32.94)	(-1.097)		
Singapore dollar	80:1-97:6	-0.001**	0.120***	0.726***	0.082*	0.907	2.098
		(-2.267)	(4.801)	(33.63)	(1.689)		
	80:1-89:12	-0.006E-01	0.095**	0.737***	-0.012	0.883	2.044
		(-0.586)	(2.381)	(23.87)	(-0.157)		
	90:1-97:6	-0.003***	0.142***	0.708***	0.206***	0.950	2.187
		(-4.046)	(5.504)	(27.53)	(3.925)		
Malaysian ringgit	80:1-97:6	0.001	0.059**	0.807***	0.093#	0.883	1.743
		(1.145)	(1.934)	(30.80)	(1.587)		
	80:1-89:12	0.002**	0.020	0.776***	-0.008	0.894	1.699
		(2.008)	(0.526)	(26.20)	(-0.108)		
	90:1-97:6	-0.007E-01	0.076*	0.866***	0.210**	0.888	1.796
		(-0.581)	(1.671)	(19.06)	(2.262)		
Indonesian rupiah	80:1-97:6	0.006***	-0.003	0.979***	0.135	0.546	2.012
		(2.747)	(-0.032)	(12.76)	(0.786)		
	80:1-89:12	0.008**	0.007E-01	0.970***	0.229	0.425	2.017
		(2.123)	(0.004)	(7.656)	(0.760)		
	90:1-97:6	0.003***	0.007	0.987***	0.020	0.995	1.761
		(13.51)	(0.716)	(106.5)	(1.080)		

Table 3.1 (Continued).

Currency	Period	Constant	Yen	USD	DM	Adj-R^2	DW
Philippine peso	80:1-97:6	0.006***	0.020	1.059***	0.034	0.609	2.105
		(2.741)	(0.239)	(14.59)	(0.207)		
	80:1-89:12	0.009***	0.066	1.050***	-0.126	0.538	2.125
		(2.737)	(0.481)	(9.833)	(-0.496)		
	90:1-97:6	0.001	-0.034	1.070***	0.263#	0.762	1.804
		(0.631)	(-0.412)	(12.92)	(1.552)		

(Notes)
[a] OLS, with the Swiss francs as the numeraire.
[b] ***, **, * and # indicate that the statistic is significant at the 1, 5, 10 and 15 percent levels, respectively. The numbers in parentheses are t-statistics.
[c] The estimated weights do not necessarily add up to unity because no such constraint is imposed.

Table 3.2.

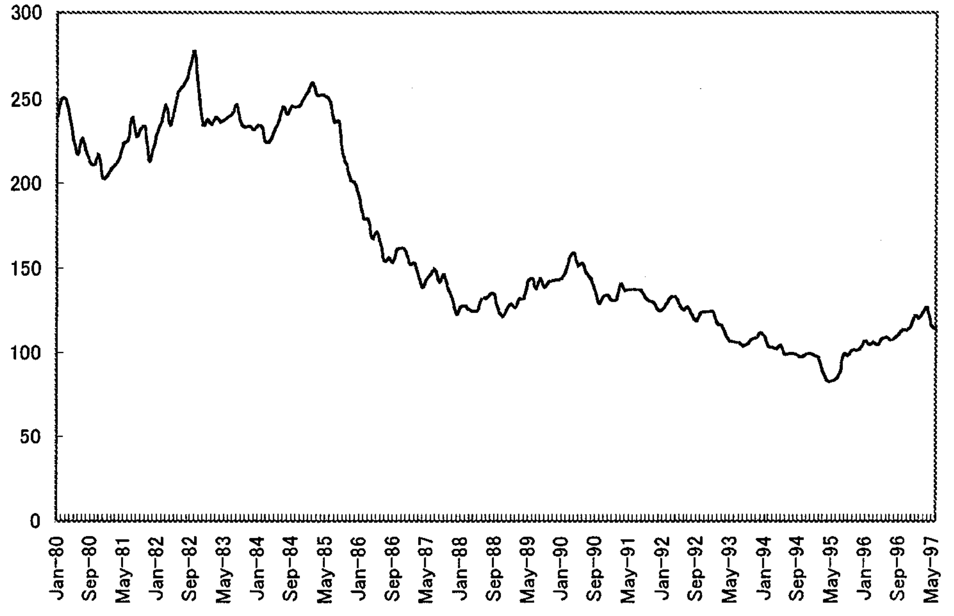
Test Statistics on the Stability of Coefficients

	CUSUMSQ test statistics		
	80:1-97:6	80:1-89:12	90:1-97:6
Thai baht	0.665*** [0.000]	0.473*** [0.000]	0.133 [0.298]
Korean won	0.146** [0.016]	0.308*** [0.000]	0.376*** [0.000]
Singapore dollar	0.229*** [0.000]	0.156* [0.076]	0.106 [0.557]
Malaysian ringgit	0.191*** [0.001]	0.175** [0.035]	0.219** [0.017]
Indonesian rupiah	0.608*** [0.000]	0.328*** [0.000]	0.429*** [0.000]
Philippine peso	0.414*** [0.000]	0.353*** [0.000]	0.292*** [0.000]

(Note)

***, *, and * indicate that the statistic is significant at the 1, 5, and 10 percent levels, respectively. The numbers in brackets are p-values.

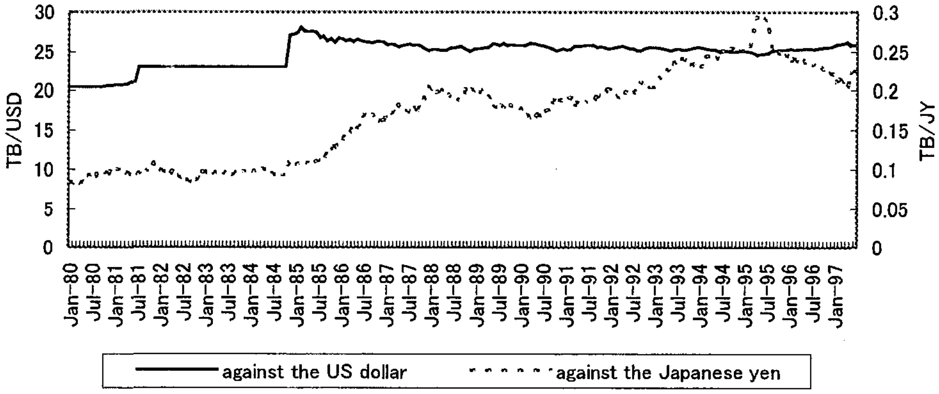
Figure 3.1.
The Movement of the Japanese Yen against the US Dollar,
January 1980 - June 1997



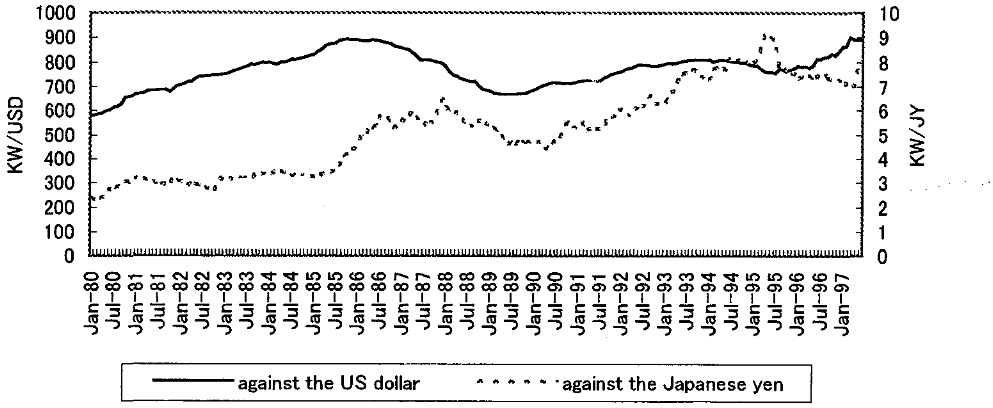
Source: IMF, *International Financial Statistics.*

Figure 3.2.
The Movements of East Asian Currencies against the US Dollar and the Japanese Yen, January 1980 - June 1997

2a. The Thai Baht



2b. The Korean Won



2c. The Singapore Dollar

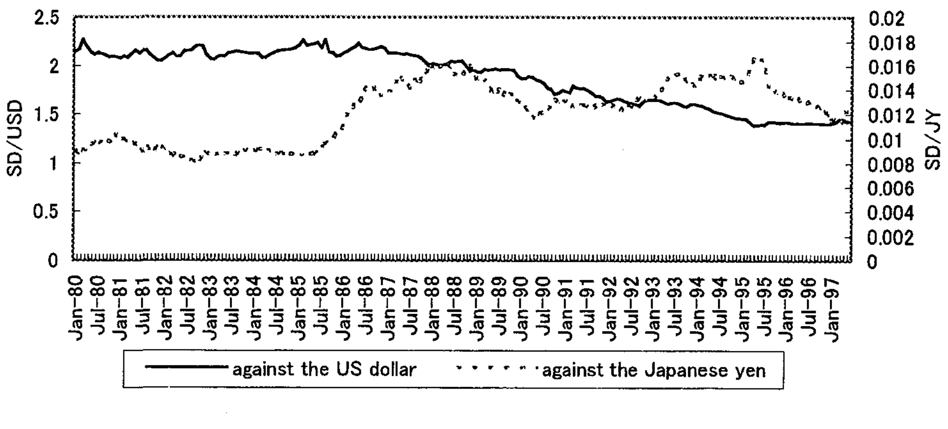
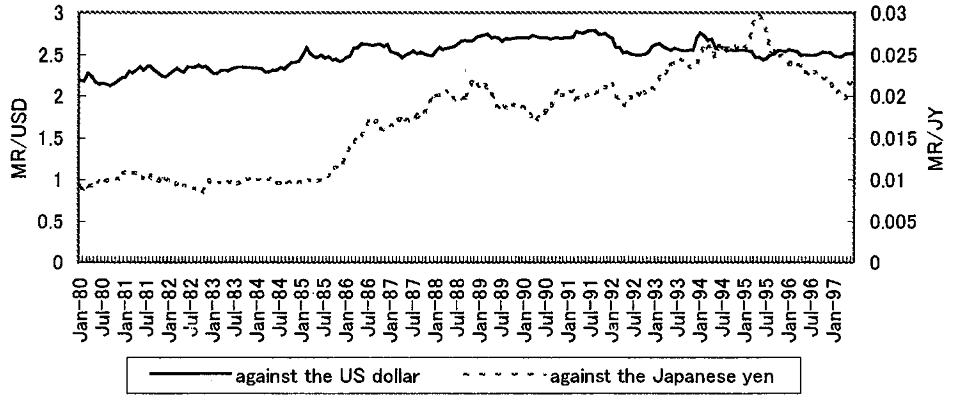
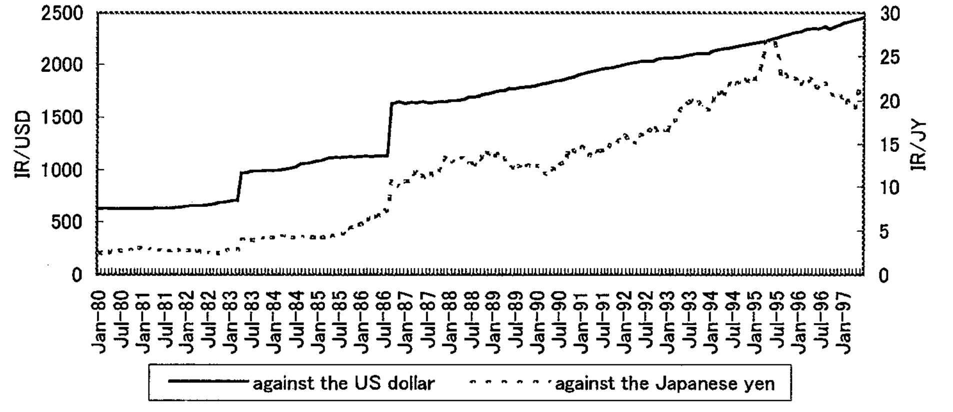


Figure 3.2 (Continued).

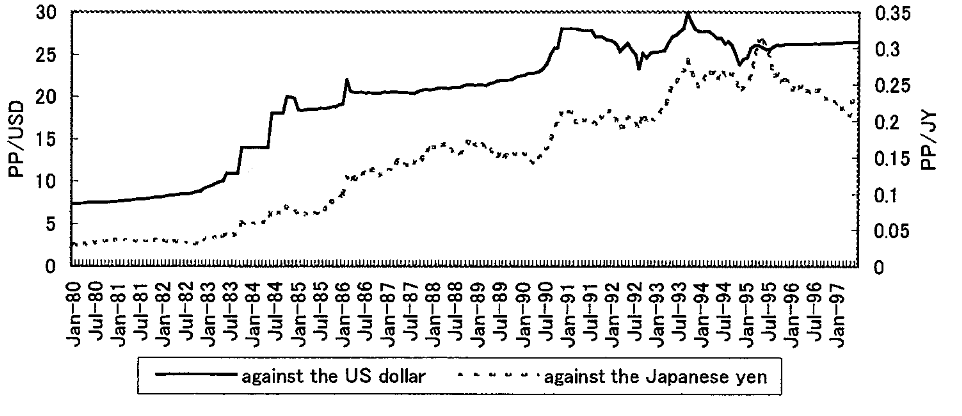
2d. The Malaysian Ringgit



2e. The Indonesian Rupiah



2f. The Philippine Peso



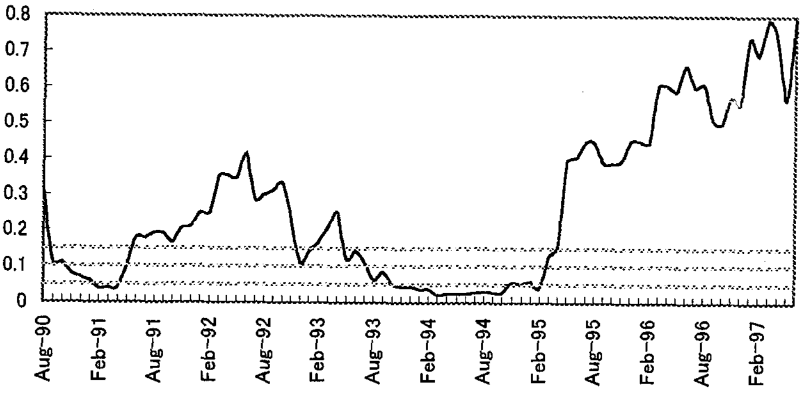
Note: End-of-month data.

Source: IMF, *International Financial Statistics.*

Figure 3.3.

P-Values of the Chow Test for the Thai Baht and the Singapore Dollar, 1990-1997

The Thai Baht



The Singapore Dollar

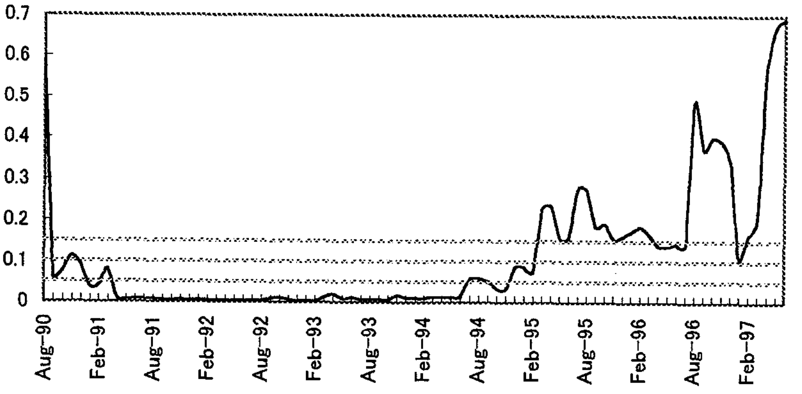
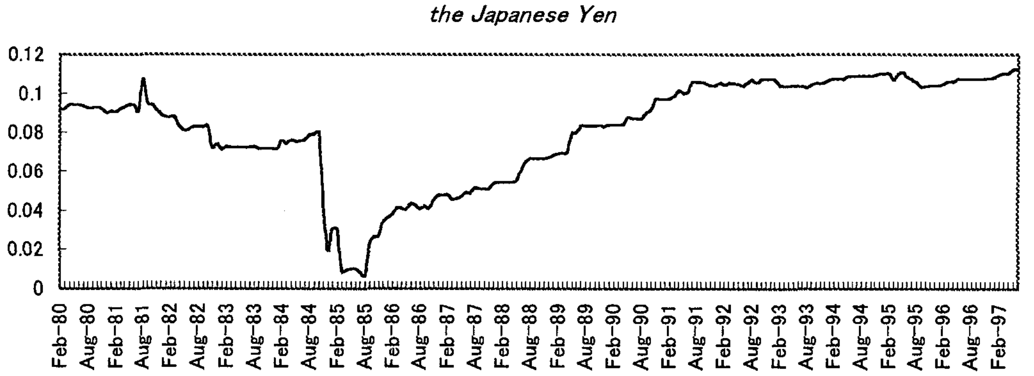
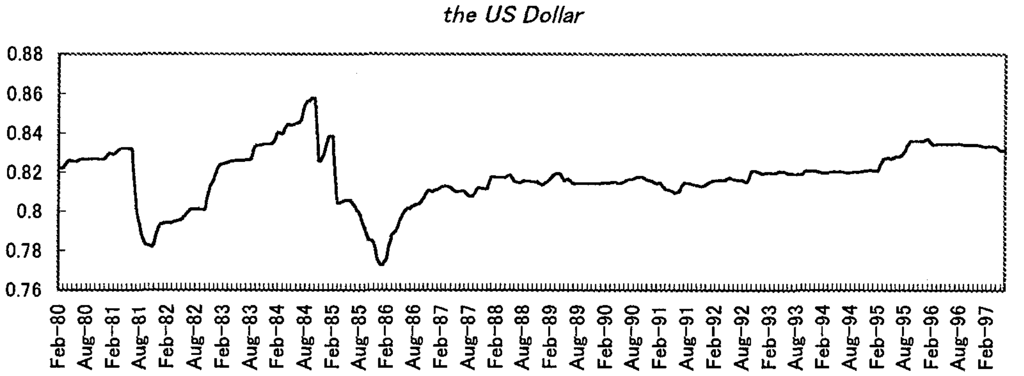


Figure 3.4.

The Thai Baht: The Time-Varying Weights of Major Currencies





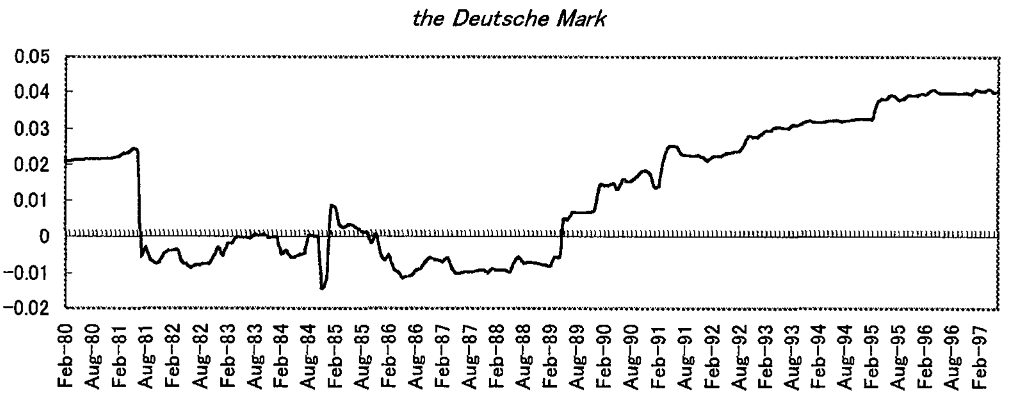
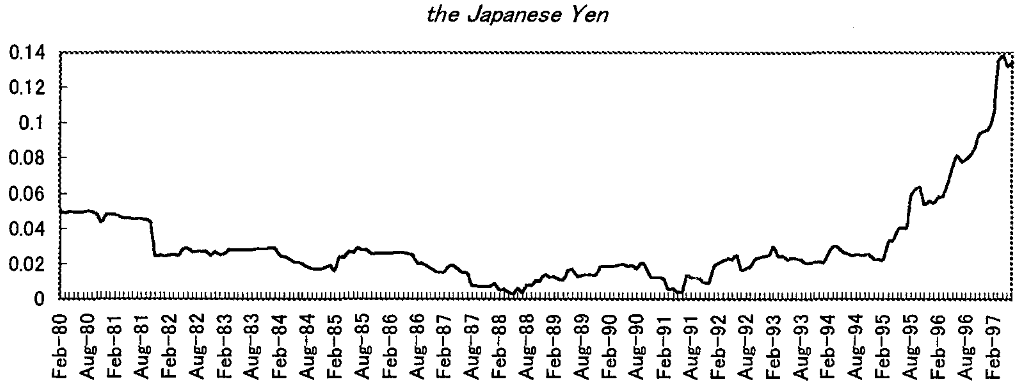
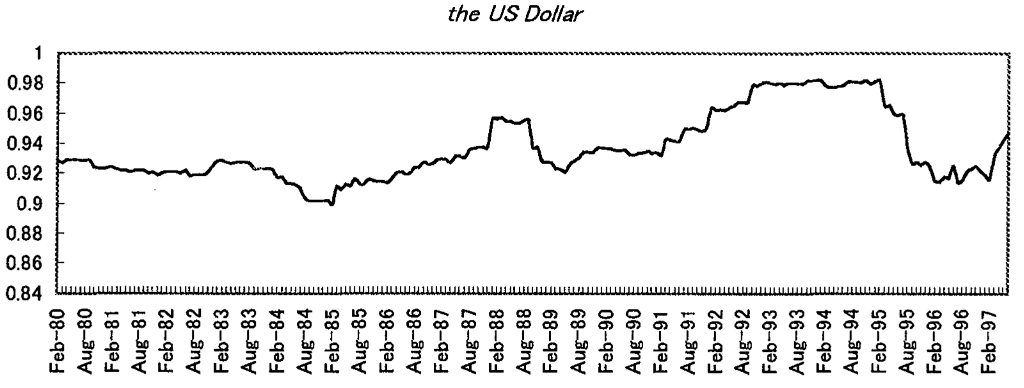


Figure 3.5.

The Korean Won: The Time-Varying Weights of Major Currencies





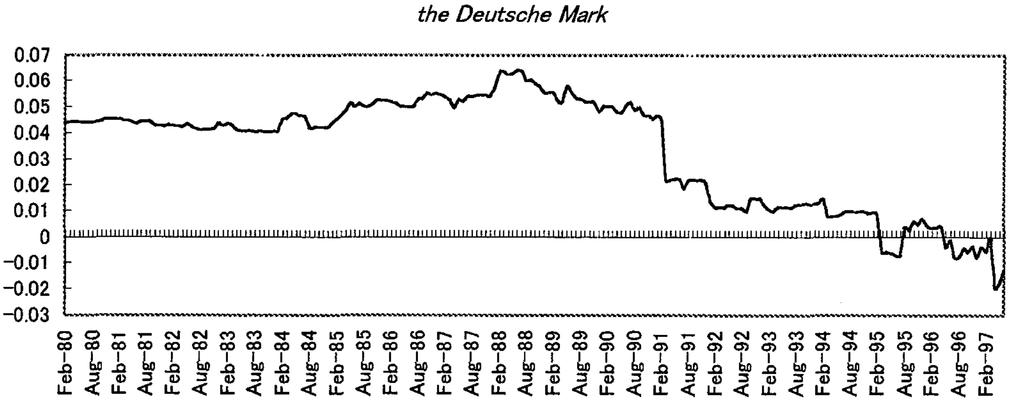


Figure 3.6.

The Singapore Dollar: The Time-Varying Weights of Major Currencies

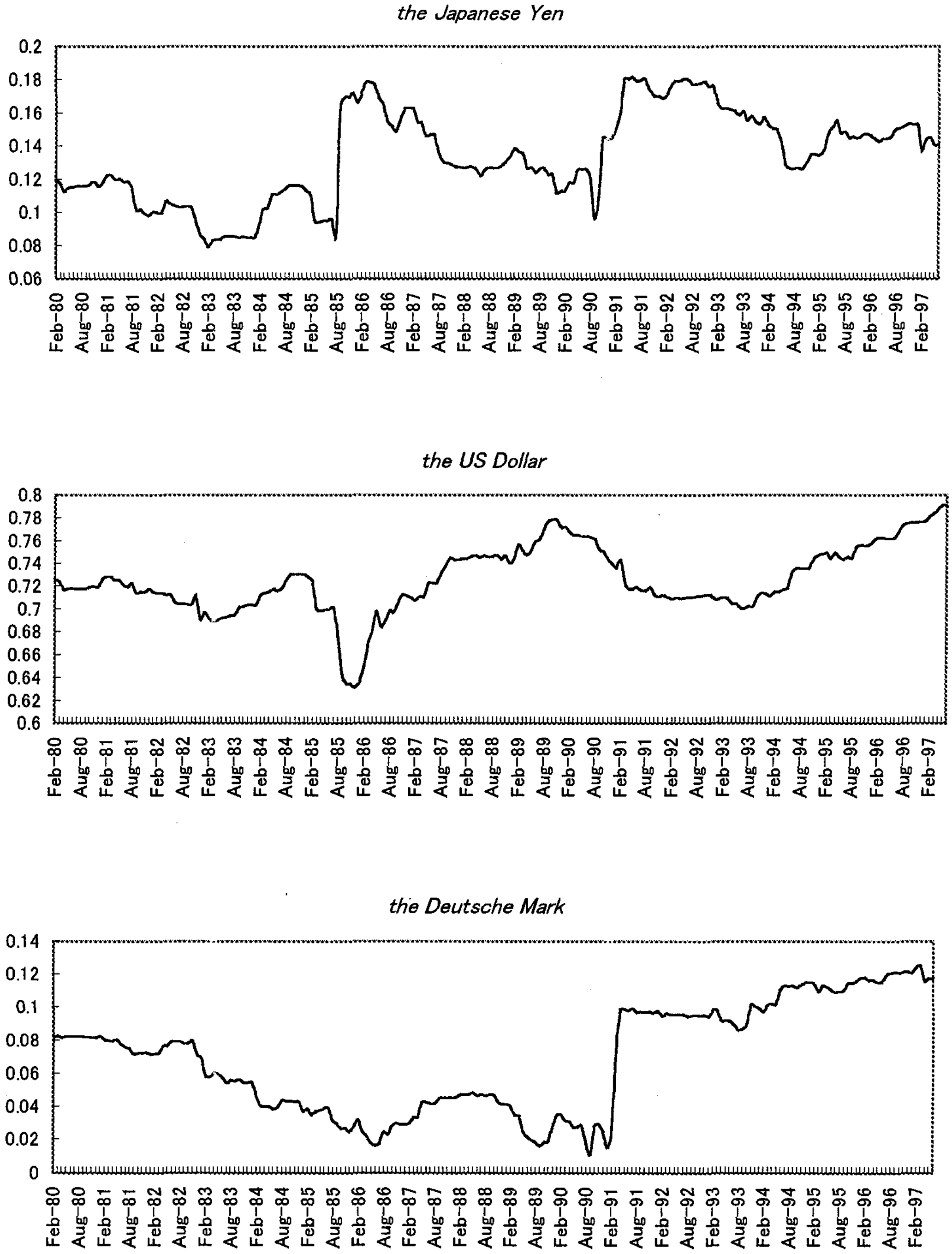
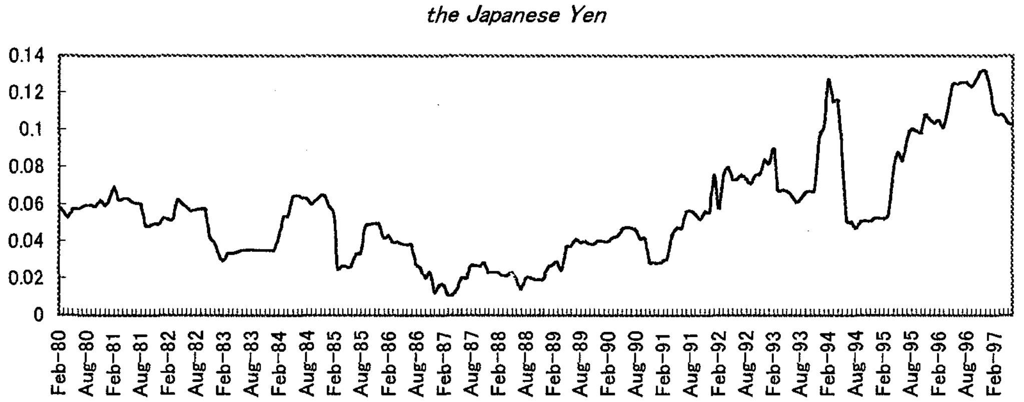
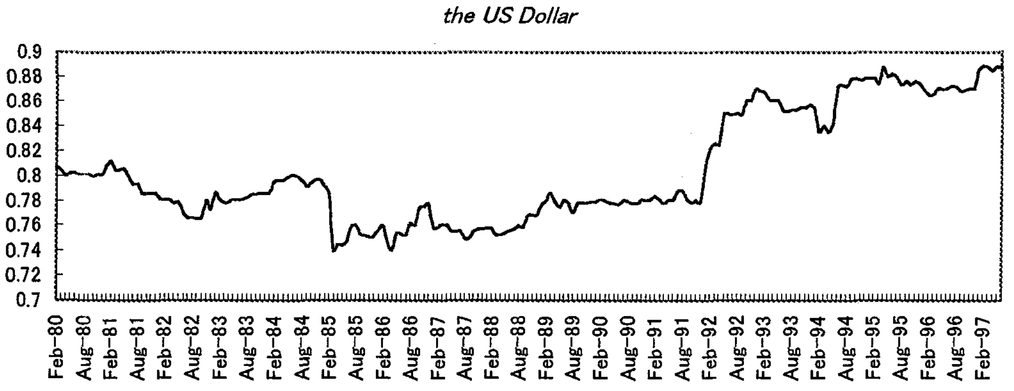


Figure 3.7.

The Malaysian Ringgit: The Time-Varying Weights of Major Currencies





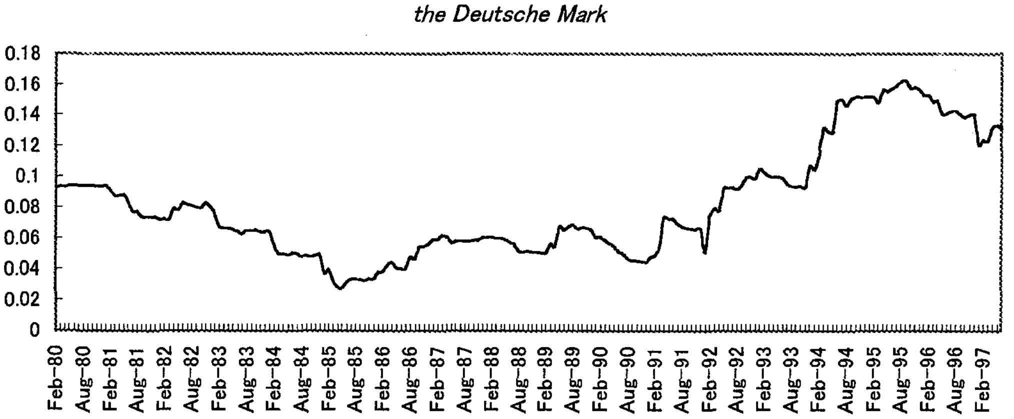


Figure 3.8.

The Indonesian Rupiah: The Time-Varying Weights of Major Currencies

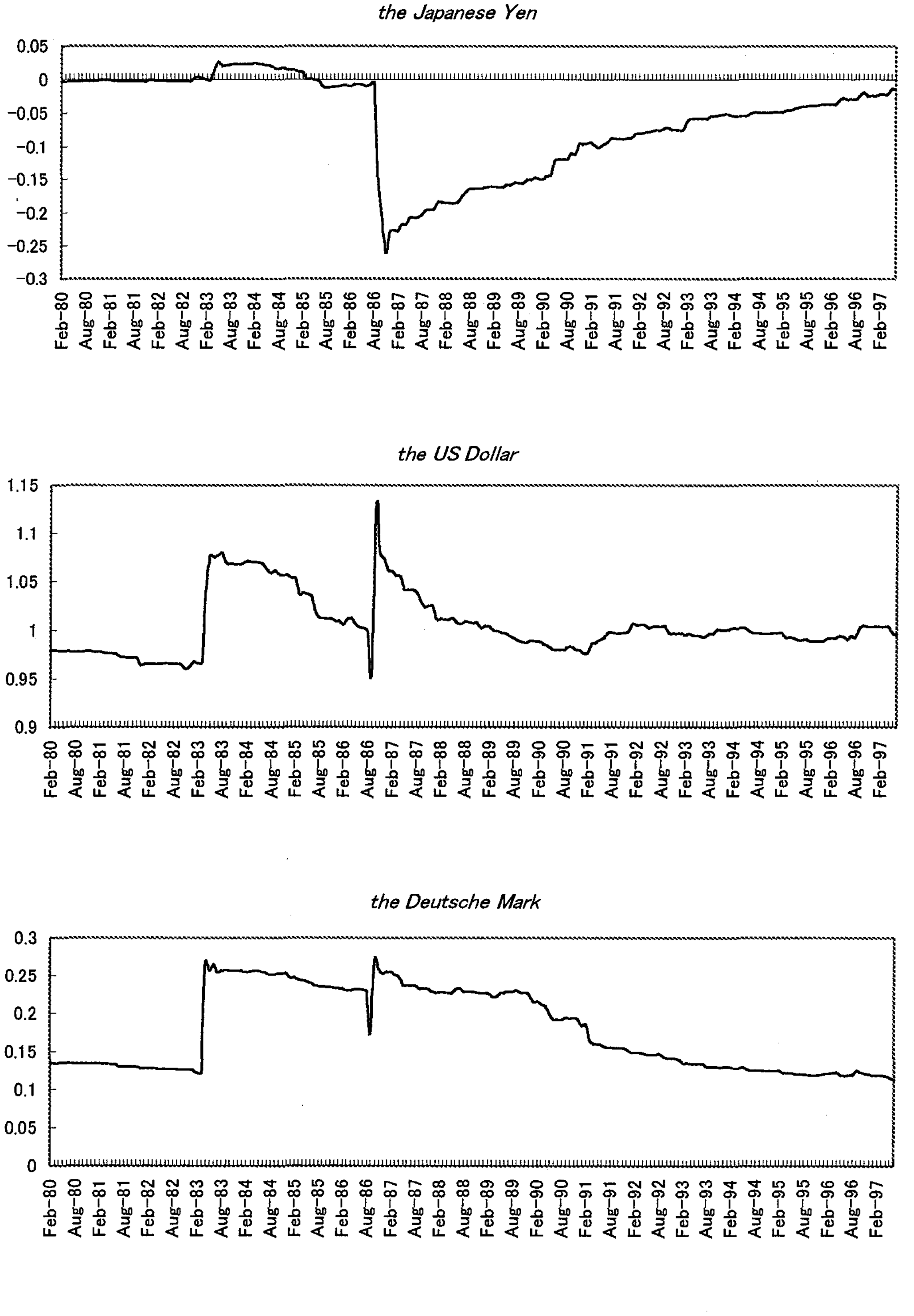
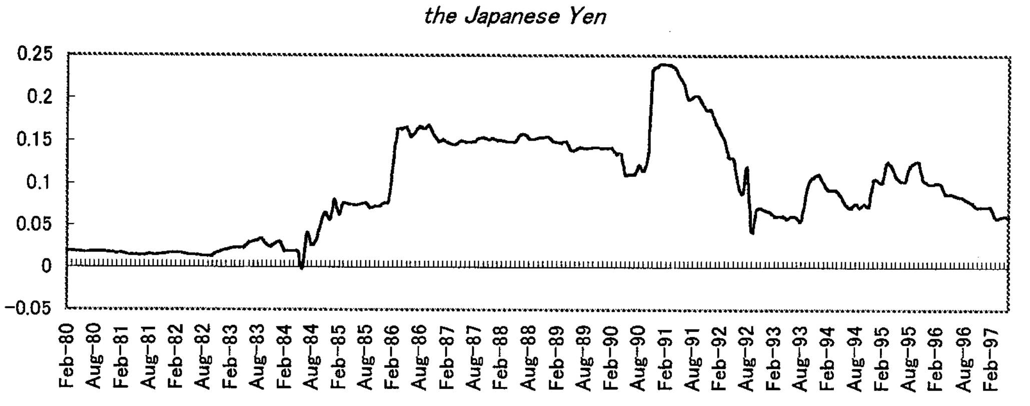
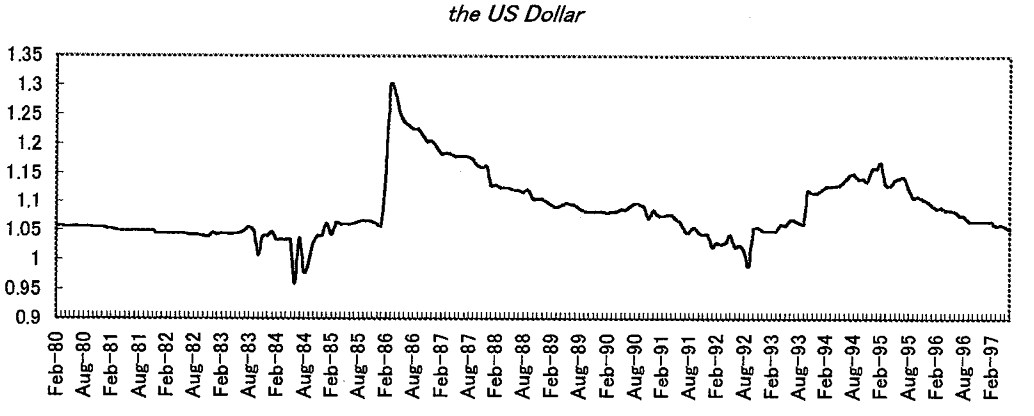
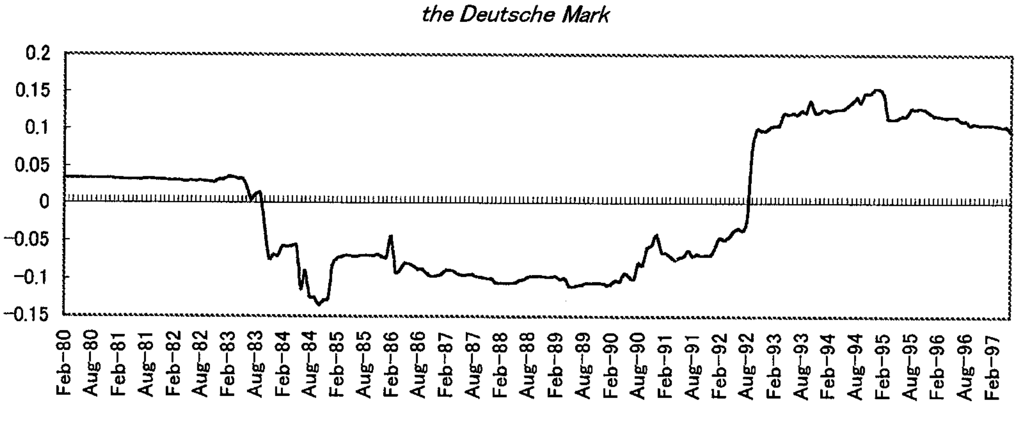


Figure 3.9.

The Philippine Peso: The Time-Varying Weights of Major Currencies



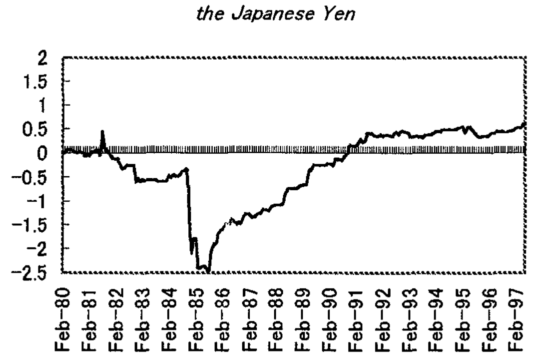


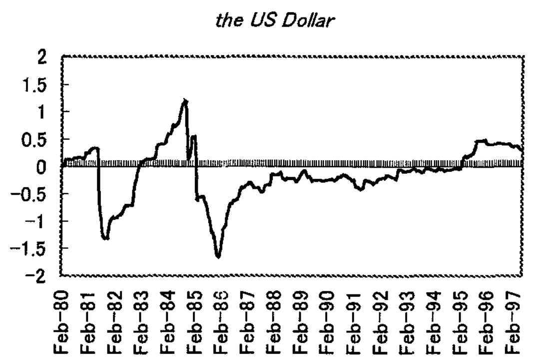


Appendix Figures.

Significance Tests of the Variation in Estimated Weights Obtained by the Kalman Filter

Figure A3.1. The Thai Baht: Test Statistics for the Variation of the Estimated Weights of Major Currencies





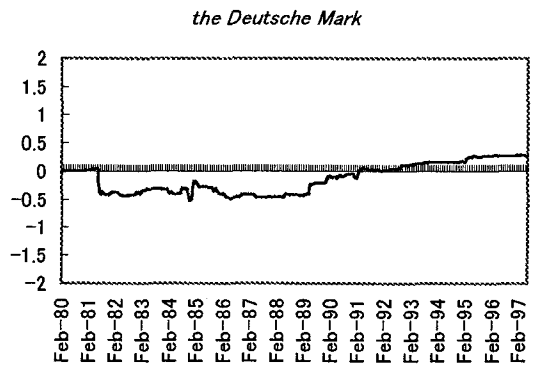
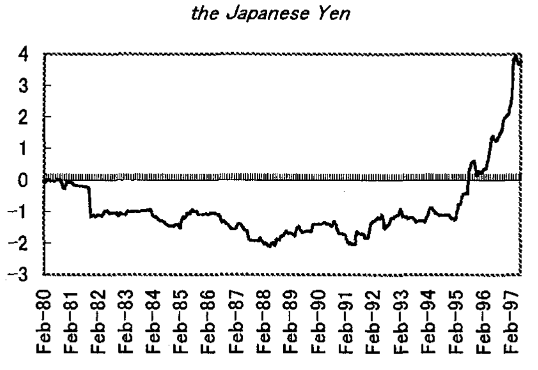
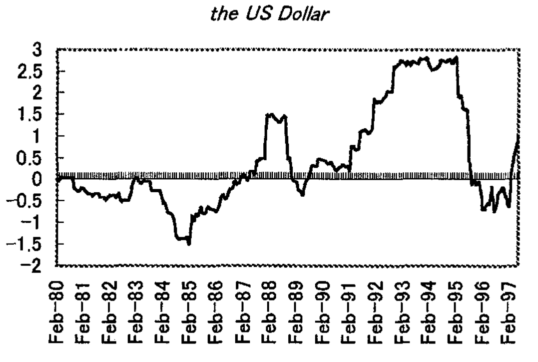


Figure A3.2. The Korean Won: Test Statistics for the Variation of the Estimated Weights of Major Currencies





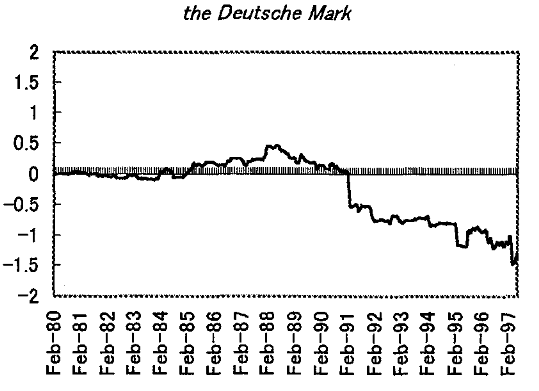
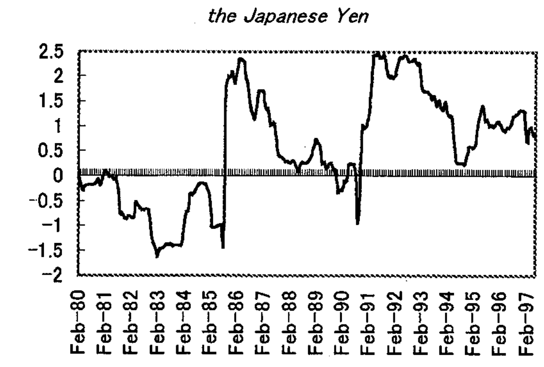
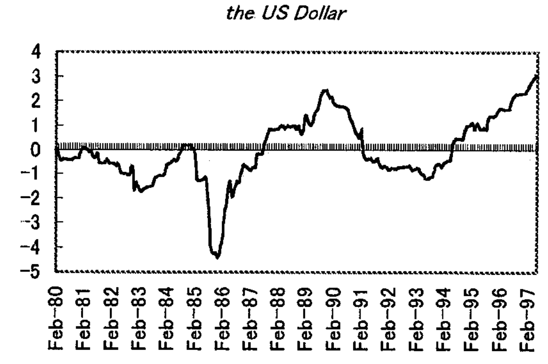


Figure A3.3. The Singapore Dollar: Test Statistics for the Variation of the Estimated Weights of Major Currencies





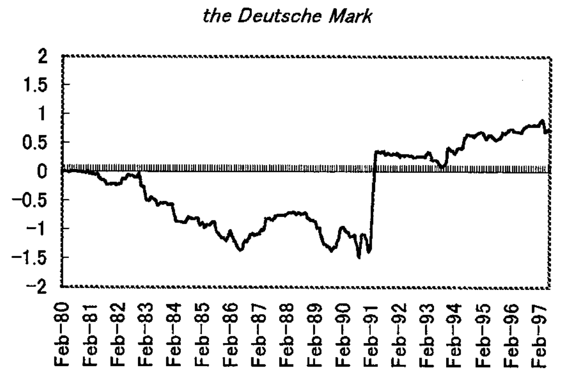
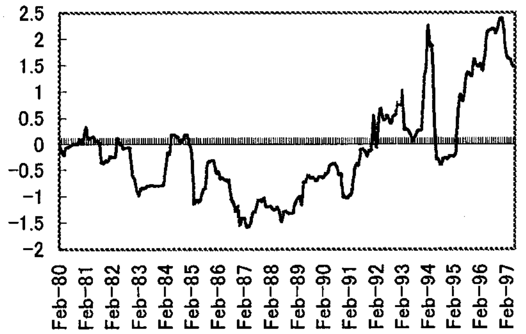
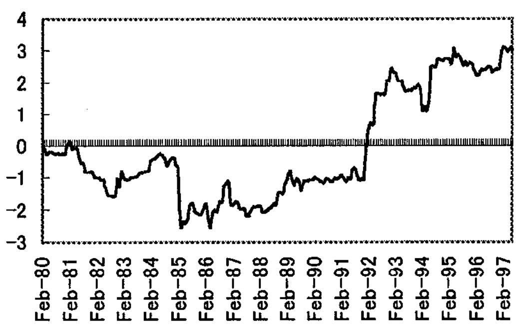


Figure A3.4. The Malaysian Ringgit: Test Statistics for the Variation of the Estimated Weights of Major Currencies

the Japanese Yen



the US Dollar



the Deutsche Mark

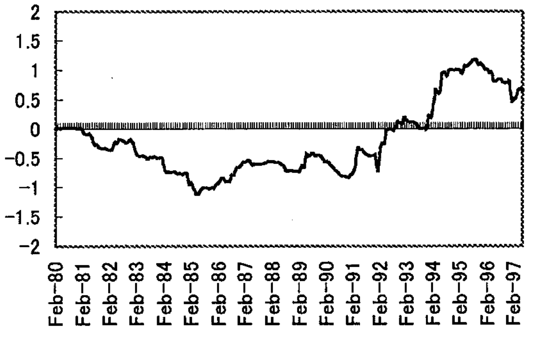
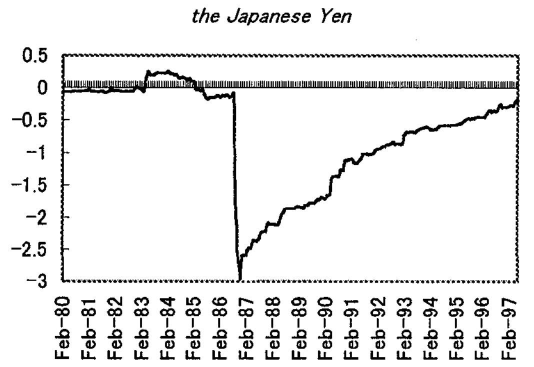
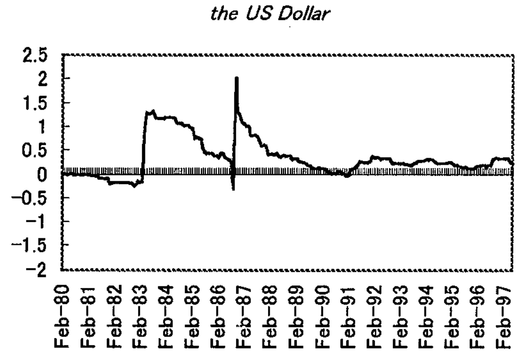


Figure A3.5. The Indonesian Rupiah: Test Statistics for the Variation of the Estimated Weights of Major Currencies





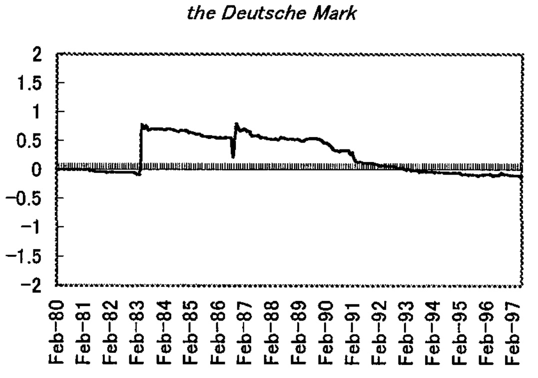
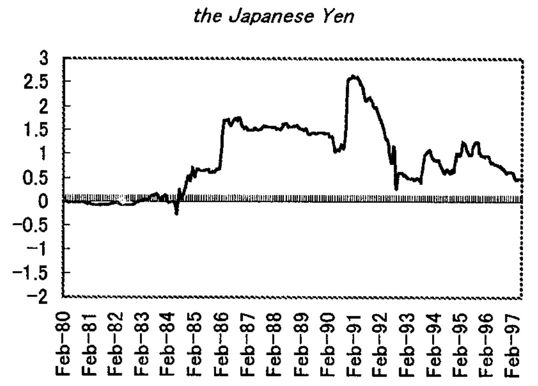
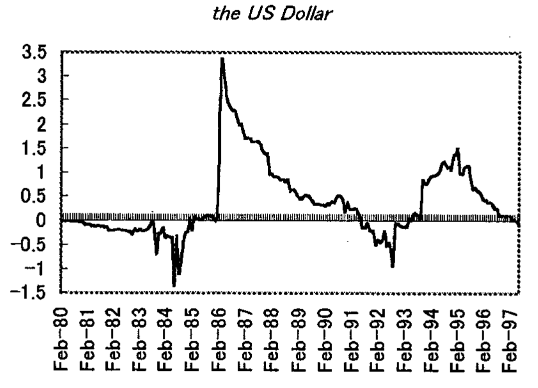
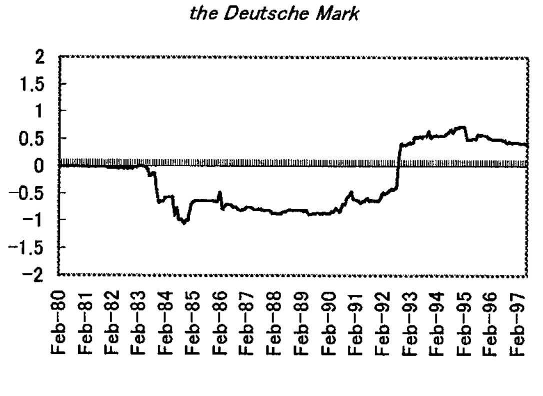


Figure A3.6. The Philippine Peso: Test Statistics for the Variation of the Estimated Weights of Major Currencies







4. Sterilization and the Capital Inflow Problem in East Asia, 1987-97

4.1. Introduction

At the end of the 1980s, a large volume of capital began to flow into the emerging market economies of East Asia, owing to both external (or "push") and internal (or "pull") factors.[1] Among other things, the factors that were external to the recipient countries included the lower interest rates, recessions, and regulatory changes favoring international portfolio diversification, all taking place in the industrialized world. The factors that were internal to the recipients included their sound economic policies (supported, for instance, by trade and capital market liberalization), exchange rate stability and deposit guarantees, and strong economic fundamentals. Roughly, the beginning of the surge in capital inflows can be identified as 1988 for Thailand, 1989 for Malaysia and the Philippines, 1990 for Indonesia, and 1990/91 for Korea (Calvo, Leiderman and Reinhart (1996), Bartolini and Drazen (1997), Chuhan, Claessens and Mamingi (1998), Montiel (1998) and Villanueva and Seng (1999)).

East Asia led the developing world in attracting private capital flows in the late 1980s, and became the most important destination for private capital flows in the early 1990s, with its share in total global capital flows to developing countries rising from around 10 percent in the early 1980s to over 40 percent in the 1990s. While the largest portion (about a half) of capital inflows was initially foreign direct investment (FDI), an increasing amount of inflows took the form of short-term borrowing in later years (Chen and Khan (1997) and Alba, Bhattacharya, Claessens, Ghosh and

Hernandez (1998)). In fact, for the period as a whole, the bulk of the capital inflows were in the form of offshore borrowing by banks and private corporations, except for Malaysia where FDI inflows remained larger than bank and private sector borrowing (Radelet and Sachs (1998)).

On an individual level, the capital inflows were massive indeed. In terms of GDP, the volume of cumulative capital inflows from 1988 to 1995 amounted to 51.5 percent in Thailand, 45.8 percent in Malaysia, 23.1 percent in the Philippines, 9.3 percent in Korea, and 8.3 percent in Indonesia. Of the two largest recipients, Malaysia received surges of massive capital inflows in 1992 and 1993, amounting to 15.3 and 23.2 percent of GDP, respectively, while Thailand received consistent flows averaging about 10 percent of GDP annually (Villanueva and Seng (1999)). At the end of 1996, the balance of claims held by foreign banks against these countries stood at $261.2 billion; of this total, $100 billion was accounted for by Korea, $69.4 billion by Thailand, $58.7 billion by Indonesia, $28.8 billion by Malaysia, and $14.1 billion by the Philippines. Except in Korea, more than a half of these claims were the obligations of the nonbank private sector (Radelet and Sachs (1998)).

Undoubtedly, capital inflows have both benefits and costs. As benefits, they promote investment and economic growth in the recipient countries, allow intertemporal smoothing in consumption, and thus raise welfare across countries. At the same time, as costs, they may lead to a rapid monetary expansion, an excessive rise in domestic demand and inflationary pressures, an appreciation of the real exchange rate, and widening current account deficits. They may even increase the vulnerability of recipients to a sudden reversal in capital flows. For these reasons, and perhaps in the light of the earlier international debt crisis, the surge in capital

[1] Latin America (particularly Argentina, Brazil, Mexico and Venezuela) was another region

inflows was, almost from the inception, perceived by the recipient countries as posing a challenge for domestic macroeconomic management, and soon began to be referred to as the "capital inflow problem" in the literature on open economy macroeconomics (Isard (1995) and Montiel (1998)).

This chapter will examine the extent to which a part of this capital inflow problem was policy-induced in the East Asian countries of Indonesia, Korea, Malaysia, the Philippines and Thailand during the decade preceding the outbreak of the currency crisis in July 1997. The motivation for this investigation comes from the large accumulation of official foreign exchange reserves in the recipient countries that was associated with the capital inflows. This indicated that the volume of capital inflows was in excess of the current account deficits; during this period, the reserve accumulation in each country amounted to between 25 and 35 percent of the total capital flows (see Section 4.2 for details). The accumulation of reserves might have been an offsetting response to the tight stance of monetary policy, which was supported by various measures to limit the expansionary impact of reserve inflows in the first place. The chapter will indirectly test whether or not such tight monetary policy measures — called in the paper broadly as sterilization — promoted additional capital inflows by keeping the level of interest rates high, by examining the effectiveness of sterilization in limiting the impact of reserve inflows on the growth of monetary aggregates.[2]

The rest of the chapter is organized as follows. Section 4.2 presents an overview of the capital inflow episode in the context of Indonesia, Korea, Malaysia, the Philippines and Thailand, by emphasizing the relationship between the capital inflows

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that attracted a large volume of capital from the late 1980s into the 1990s.

[2] The exclusive emphasis of this chapter is on the domestic monetary system of the recipient country, as our primary interest lies in the effectiveness of sterilization as a monetary policy measure. In the other hand, Montiel and Reinhart (1999) directly test the effect of sterilization on

and the growth of monetary aggregates. Section 4.3 summarizes the policy responses, collectively called sterilization, taken by the East Asian monetary authorities to limit the expansionary impact of reserve inflows on the growth of monetary aggregates. Section 4.4 tests for the effectiveness of sterilization in limiting the growth of monetary aggregates, by using both time-series and structural approaches. Finally, Section 4.5 presents concluding remarks.

4.2. An Overview of the Capital Inflow Episode in East Asia

During the capital inflow episode, the volume of capital inflows (as measured by the surplus in the capital and financial account) exceeded the deficit in the current account in all of the countries concerned, hence resulting in increases in the foreign asset source component of the monetary base. In Indonesia, for example, there was a capital inflow of $4,495 million against the current account deficit of $2,988 million in 1990 (the year in which the surge in inflows began), with an increase in the foreign exchange reserve of $2,088 million (or about 46 percent of the net capital inflows).[3] For the period 1989-96, about 26 percent of the net capital inflows were accumulated as foreign exchange reserves in Indonesia.

A similar story can be told for the other countries. In Korea, the proportion of the net capital inflows which were accumulated as foreign exchange reserves was about 32 percent for the period 1992-96. It was particularly high in 1992 (when there was a net capital inflow of $6,994 million against the current account deficit of $3,944 million) and in 1993 (when there was a net capital inflow of $3,217 million against the

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the volume and composition of capital inflows.

[3] These balance of payments figures do not necessarily add up to zero because of errors and omissions. The figures are all from the IMF, *International Financial Statistics*.

current account surplus of $990 million). In Malaysia, almost 80 percent of the net capital inflows were accumulated as foreign exchange reserves from 1989 (when the surge in inflows began) to 1993. However, it lost reserves in 1994 and 1995 before moderately gaining them again in 1996. About a third of the net capital inflows were accumulated as foreign exchange reserves in both the Philippines and Thailand during the inflow period.

Reflecting the accumulation of foreign exchange reserves, the foreign assets (FA) source component of the monetary base rapidly expanded in these countries.[4] At the same time, all the countries saw a rapid growth in both narrow and broad money (M1 and M2). In Indonesia, for example, FA rose about 5 times from 1989 to 1996, with M1 rising 2.5 times and M2 4.7 times during the same period; over the entire sample period, however, there seems to be a closer correspondence between FA and M1 (Figure 4.1). In Korea, FA, M1 and M2 all increased by roughly the same percentage (*i.e.,* 2.6 times, 1.8 times and 2.1 times, respectively, from 1991 to 1996); one can observe volatile changes in the growth of M1 (Figure 4.2).

In Malaysia, FA rose 3.2 times from 1989 to 1996, with M1 and M2 both rising 3.6 times. Corresponding to the surge of capital inflows, there was a rapid growth in FA from 1992 to early 1994; M1 then contracted through the first part of 1995 (Figure 4.3). In the Philippines, FA rose 5.8 times from 1989 to 1996, with M1 rising 2.9 times and M2 about 4 times. There were volatile fluctuations in the growth of FA; similar but more subdued fluctuations were observed for the growth of M1, sometimes displaying negative correlations between the two (Figure 4.4). Finally, in Thailand, FA rose 5.5 times from 1988 to 1996, with M1 rising 2.9 times and M2 3.9 times (Figure 4.5).

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[4] There is not necessarily a perfect correspondence between changes in the value of foreign assets held by the monetary authorities and the official settlement accounts in the balance of payments, owing to valuation and other accounting differences.

In each country, there was a sustained growth in FA, which was associated with the sustained growth in M1 and M2, hence giving rise to the common view that the surge in FA associated with the capital inflows somehow caused the rapid growth of monetary aggregates during the capital inflow episode. The validity of this view will be the subject of our investigation in the sections to follow.

4.3. Policy Responses to the Capital Inflows

As stated earlier, it was feared from the very beginning that the capital inflows might lead to a rapid monetary expansion, an excessive rise in aggregate demand and inflationary pressures, an appreciation of the real exchange rate, and widening current account deficits. For this reason, the monetary authorities of East Asian countries resorted to various policy measures to mitigate that possibility, including capital controls, trade liberalization, greater exchange rate flexibility, fiscal contraction and a variety of monetary measures (Montiel (1998), Reinhart and Reinhart (1998) and Villanueva and Seng (1999)). The monetary measures, the focus of the present chapter, included the conventional form of sterilized intervention (designed to offset the effect of reserve inflows on the monetary base by open market sales of domestic securities), increases in reserve requirements (designed to limit the impact of reserve inflows on the growth of monetary aggregates by reducing the money multiplier), shifting of government deposits from commercial banks to the central bank, an increase in the discount rate or otherwise a greater limit on the discount window, moral suasion and credit controls. Of these and other monetary measures, sterilized intervention and the tightening of reserve requirements were the most common and were employed by all of the central banks concerned at one time or another.

By far, the most common and extensive was sterilized intervention, at least initially. Often lacking the depth of markets in government securities, the East Asian central banks supplemented operations in government securities by issuing their own debt instruments (Villanueva and Seng (1999)). For example, in 1987, the Bank of Thailand (BOT) began to issue short-term BOT bonds with maturities of 6 months to one year. Monetary Stabilization Bonds (MSBs) and Bank Indonesia Certificates (SBIs) were the principal tools of open market operations used by the Bank of Korea and Bank Indonesia, respectively.[5] The Central Bank of the Philippines had routinely used Central Bank Certificates of Indebtedness (CBCIs) at least until 1994, when open market operations in government securities gained prominence. Even in Malaysia where the market for government securities is fairly well developed by East Asian standards, Bank Negara issued series of Bank Negara Bills and Malaysian Savings Bonds during the peak inflow period of 1993.[6]

After the initial period, however, most of the central banks began to rely much less on conventional sterilized intervention, in part owing to the quasi-fiscal costs of such operations. The quasi-fiscal cost arises because, in sterilized intervention, the central bank typically exchanges high-yielding domestic assets for low-yielding foreign assets (Calvo (1991) and Kletzer and Spiegel (1998)). In the consolidated government and central bank portfolio, the public sector ends up paying more on its liabilities than it receives on its assets, as more of government debt is held outside the central bank. Villanueva and Seng (1999) identify the period of active sterilized intervention as 1988-95 for Thailand, 1989 and 1992-93 for Korea, 1990-93 and 1996 for Indonesia, 1990-93 for the Philippines, and 1992-93 for Malaysia. Thus, it was only in Thailand

[5] In Korea, the first auction in MSBs was conducted in April 1993, although they had been issued earlier. In Indonesia, SBIs were first issued in 1984.

[6] In Malaysia, central bank securities were first issued in 1987.

that sterilized intervention was used consistently throughout much of the capital inflow episode.

In addition to sterilized intervention, other measures were also used to control either the monetary base or the growth of monetary aggregates. Measures to control base money included central bank borrowing from commercial banks, and the shifting of government deposits from commercial banks to the central bank. The latter tool was frequently used in Malaysia, Thailand and Indonesia. In Malaysia, the most important funds to be so shifted were deposits of the Employee Provident Fund (EPF). It is said that more than US$2.6 billion of EPF funds were shifted from commercial banks to Bank Negara in 1992 (Villanueva and Seng (1999)). In the Philippines, the government borrowed from the private sector to make deposits at the central bank. Access to the discount window was reduced in Korea during 1986-88, in Thailand during 1989-90, and in Malaysia during 1995-96. In Indonesia, moral suasion and various reporting requirements were imposed on commercial banks during 1994-96. Some control measures acted almost like cross-border capital controls, such as the ceiling on the external liabilities of domestic banks and the prohibition of sales of short-term financial instruments to foreigners, both imposed by Malaysia for several months during 1994.

The most common tool of containing the growth of monetary aggregates (while accepting the increase in base money itself) was to effect a rise in reserve requirements. Malaysia frequently raised reserve requirements and expanded the coverage of institutions and deposits subject to the requirements. Indonesia and Thailand, although initially reluctant to raise reserve requirements, became more active users of this tool in later years. Villanueva and Seng (1999) identify the period during which the reserve requirements were raised as 1989-92, 1994 and 1996 for Malaysia, 1990 for

Korea and the Philippines, 1995-96 for Thailand, and 1996 for Indonesia.

In this chapter, as elsewhere in the recent literature on this subject, what we call sterilization includes not only the conventional form of sterilized intervention (in which domestic and foreign securities are exchanged in an open market transaction), which may be terms "sterilization in the narrow sense", but also any form of transaction which is designed to limit the impact of reserve inflows on the growth of monetary aggregates, which may be termed "sterilization in the broader sense". Whether it is defined narrowly or broadly, sterilization tends to raise the level of domestic interest rates, provided that foreign and domestic assets are imperfect substitutes and hence sterilization is effective.[7]

In the case of narrowly defined sterilization, domestic interest rates rise so as to induce the market participants to hold the greater amount of domestic assets willingly. In the case of broadly defined sterilization, domestic interest rates rise so as to clear the money market, given the restricted money supply. In either case, a rise in foreign assets would be prevented from increasing the volume of monetary aggregates at least one to one, and the resulting rise in interest rate differentials favoring the domestic assets would promote additional capital inflows, given flexible but stable nominal exchange rates (Takagi (1999)). Of course, no additional capital inflows would result if the market participants correctly perceived that the higher interest rates only reflected the higher risk premium of domestic assets and the non-zero probability of currency depreciation. However, it is said that many market participants tried to exploit the interest rate differentials that existed between US-dollar denominated and East Asian currency-denominated assets by taking unhedged short-term positions for supposed financial gains, believing that the markets were imperfect (Furman and

[7] It should be noted that, in practice, sterilization was generally supported by tight fiscal

Stiglitz (1998), particularly footnote 34).

4.4. Estimating the Effectiveness of Sterilization

The foregoing discussion makes it clear that, in testing for the effectiveness of sterilization, the conventional method of estimating the offset coefficient of the capital flow equation along with the monetary policy reaction function would be inappropriate in the context of the East Asian experience (for an example of the conventional method applied to developing countries, see Takagi (1986)). In East Asia, various monetary measures were used at various times in various intensities in order to sterilize the effect of capital inflows on the growth of monetary aggregates. For this reason, in what follows, we will test for the effectiveness of sterilization by estimating the extent to which foreign assets (FA) in the monetary base explains or predicts monetary aggregates, setting aside the question of how sterilization is actually effected.

We will use quarterly data for the 10-year period from the first quarter of 1987 through the second quarter of 1997, immediately preceding the outbreak of the Thai crisis in July 1997. Both narrow money (M1) and broad money (M2) are used as measures of monetary aggregates, and consumer price indices are used as the price level (P). For Korea and the Philippines, real GDP is used for output (Y), whereas industrial production is used for the other three countries. For the interest rate (i), the money market rate is used (see the appendix for the sources and descriptions of the data). Table 4.1 summarizes the time-series properties of the variables, where all but the interest rate are expressed in natural logarithm. The table overwhelmingly suggests that the variables are integrated of order one, i.e., I(1). The only exceptions

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policy, which reinforced the upward pressure on the level of interest rates.

are nominal and real FA, and Y in Thailand. Although not formally reported in the table, all the variables are found to become stationary when they are differenced once.

4.4.1. Cointegration Tests

Before proceeding further, we test for the presence of cointegration between money and foreign assets by using Johansen's trace tests, with lag length chosen by Schwarz's Bayesian information criterion (SBIC). In a bivariate system (expressed in natural logarithm), we find that neither M1 nor M2 is found to be cointegrated with FA, except for M2 in Indonesia (Table 4.2). In a multivariate system (consisting of real M1 or M2, real FA, i and Y, where all but i are expressed in natural logarithm), a cointegrating relationship is found only for M1 in the Philippines. In what follows, given the overwhelming evidence that all variables are I(1) and the general absence of cointegration, we will estimate regression equations in first difference form without an error correction term.

4.4.2. Granger Causality Tests

First, we will test for Granger causality between money and foreign assets. A stationary time-series x (*e.g.,* FA) is said to Granger cause a stationary time-series z (*e.g.,* M1 or M2), if the hypothesis that the coefficients c_j are collectively zero can be rejected at a given level of significance.

$$z_t = a + \sum b_j z_{t-j} + \sum c_j x_{t-j} + sw + e_t, \qquad (1)$$

where t is a time subscript, a is a constant, $\sum$ is a summation from 1 to k (where lag length (k) is chosen by SBIC), b_j' s are the coefficients of the lagged dependent variables, w is a vector of other variables, including seasonal dummies and, in a multivariate system, the lagged values of other variables, such as output and the interest rate, s is a vector of coefficients associated with w, and e is a random error term. Both causality from FA to M1 or M2 and causality from M1 or M2 to FA are tested, although only the first type of causality, which is the focus of this chapter, is discussed in the text below.[8]

In a bivariate system with FA and M1 or M2 (in logarithmic differences), FA is found to Granger cause M only in Malaysia when M1 is used and in the Philippines when M2 is used, both at the 10 percent level of significance (Table 4.3). At the 5 percent level of significance, however, no Granger causality is found from FA to either M1 or M2.[9] In a multivariate system with real FA, real M1 or M2, Y and i (in logarithmic differences, except for i which is expressed in simple first difference), no Granger causality is found at the 10 percent level of significance or lower (Table 4.4). To the extent that the multivariate system can generally be considered more appropriate,[10] we conclude that no Granger causality was found from foreign assets to monetary aggregates during 1987-97 in any of the countries.[11]

Another important channel of influence concerns how a change in FA might have

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[8] As we are considering the impact of a change in FA on monetary aggregates (which must be effected through the banking sector and presumably takes some time), we believe that the use of quarterly data is appropriate. If the adjustment of monetary aggregates in response to a change in FA is completed quickly within a quarter, however, Granger causality is not revealed in quarterly data.

[9] We have also followed the procedure of Toda and Yamamoto (1995) to apply Granger causality tests in the levels of integrated or cointegrated variables. In a bivariate system, the only evidence of causality from FA to money is found in the case of Malaysia (at the 5 percent level of significance) when M1 is used.

[10] If the true model includes more variables, the bivariate system of foreign assets and money may show a spurious relationship.

[11] As an additional test, we have also applied Granger causality tests in Johansen's error correction model (ECM) framework, given the possible presence of cointegration between FA and M2 in Indonesia and between real M1, real FA, Y and i in the Philippines (see Table3.2). On the basis of the procedure of Toda and Philips (1993), the only evidence of causality (from FA to M1) was found for the Philippines at the 10 percent level of significance. Hence, our conclusion based on Table 3.3 and Table 4 does not change.

affected the level of interest rates. Our earlier discussion suggested that effective sterilization would limit the growth of monetary aggregates and raise the level of interest rates at the same time. So far, the causality tests (along the lines of equation (1)) have suggested the possibility that sterilization was effective in limiting the growth of monetary aggregates. How was then the level of interest rates affected by sterilization, given a change in FA? Table 4.5 reports the results of multivariate causality tests in logarithmic differences (except for i which is expressed in simple differences). The tests suggest, rather surprisingly, that no Granger causality was found from FA to the money market rate during 1987-97 for any of the countries, except for the Philippines, where causality was found at the 1 percent significance level regardless of whether M1 or M2 was used. This may mean that sterilization was effective, not necessarily in raising the level of interest rates, but in keeping it from falling toward the world interest rates. More will be said on this point in the concluding section.

4.4.3. Tests of structural equations

Second, as an additional test of the effect of foreign assets on the growth of monetary aggregates, we will next estimate the following structural equation.

$$\Delta \ln(M_t / P_t) = d + h\Delta \ln(FA_{t-1} / P_{t-1}) + qv + u_t, \tag{2}$$

where Δ is a first difference operator, M is either M1 or M2, d is a constant, h is the coefficient of lagged foreign assets, q is a vector of coefficients, v is a vector of other explanatory variables, including seasonal dummies, $\Delta \ln Y$ and $\Delta\ i$, and u is a random error term.

Equation (2) includes lagged FA, and not current FA, because a change in FA is believed to affect M1 or M2 over time through the banking sector. Use of lagged FA also has an additional advantage in that it alleviates the potential difficulty with M1 or M2 affecting FA contemporaneously. Moreover, in the light of the earlier causality test that, except for the Philippines, there was no causality between FA and i in either direction, there is no need to worry about correlation between lagged FA and i, either (except for the Philippines, of course). However, equation (2) is estimated both with and without i in order to check robustness. We are particularly interested in the estimated value of h.

Table 4.6 through Table 4.10 (first two columns under each heading M1 or M2) report the results of estimating equation (2) by OLS for Indonesia, Korea, Malaysia, the Philippines and Thailand. The F-statistics are generally significant (except for Indonesia and Malaysia when M2 is used); considering that the regression equation is estimated in first difference form, the R-squared is remarkably high, especially when M1 is used. The coefficient of output is positive when it is significant, while the coefficient of the interest rate is negative when it is significant. Many of the coefficients of the seasonal dummies (not formally reported in the tables) are significant.

From these tables, we find that regardless of whether M1 or M2 is used or whether i is included or not, the coefficient of lagged FA (h) is not significantly different from zero. The only exception is found for the Philippines when M2 is used and i is included. Because of the potential simultaneity problem, not too much confidence can be placed in the present result at the present time. So far as this result is concerned, however, the coefficient (h) is negative, suggesting that a rise in foreign assets reduces M2 in the next period. All in all, the overall weight of the evidence seems to suggest that sterilization was effective in limiting the growth of monetary aggregates during 1987-97 in all countries, affirming the results of the Granger causality tests.

Finally, the tables (last two columns under each heading) also report the results of estimating equation (2) by including a slope dummy for the coefficient of $\Delta \ln(FA_{t-1} / P_{t-1})$, with the dummy indicating the intensity of sterilization.

$$\Delta \ln(M_t / P_t) = d + h_1 \Delta \ln(FA_{t-1} / P_{t-1}) + DUM_t h_2 \Delta \ln(FA_{t-1} / P_{t-1}) + qv + u_t, \qquad (3)$$

where DUM is the dummy variable which takes the value of unity when sterilization is considered to be particularly intense, and h_1 and h_2 (replacing h) are the coefficients of lagged real foreign assets under normal conditions and under intense sterilization, respectively. The annual series of dummy variables were constructed on the basis of information provided by Villanueva and Seng (1999) and a similar construction of the sterilization index presented by Reinhart and Reinhart (1998) and Montiel and Reinhart (1999). The quarterly series are created by simply assuming that, during a given calendar year, they take the same value as the annual series. Here, sterilization was considered to be intense if open market operations were large in scale and accompanied by increased reserve requirements or transfers of government deposits from commercial banks to the central bank (see the annual series in Table 4.11).

We consider equation (3) in order to see whether or not the relationship between FA and monetary aggregates was invariant through time. If the policy of intense sterilization was particularly effective in limiting the impact of an increase in FA on the growth of M1 or M2, we should expect the value of h_2 to be negative, so that the coefficient of (FA_{t-1} / P_{t-1}) under intense sterilization (i.e., $h_1 + h_2$) is algebraically smaller than that under normal conditions (h_1). Because no sterilization was considered intense in Korea, the results are reported for the other four countries only. The last two columns under each heading show that the coefficient h_2 is not statistically significant in any of the countries regardless of whether M1 or M2 is chosen (confirming the earlier results obtained without the slope dummies), although it is indeed negative in Indonesia, the Philippines and Thailand. We can thus reaffirm our earlier conclusion that sterilization was effective in limiting the growth of monetary aggregates during 1987-97, with the additional insight that the effectiveness of sterilization was indeed greater (albeit marginally) when it was intense.

4.5. Summary and Concluding Remarks

The East Asian countries of Indonesia, Korea, Malaysia, the Philippines and Thailand received large volumes of capital inflows from the end of the 1980s through early 1997. The cumulative inflows were massive indeed, amounting to 50 percent of GDP in Malaysia and Thailand, over 20 percent in the Philippines and about 10 percent in Indonesia and Korea. Although a large portion of the inflows initially took the form of FDI, they increasingly took the form of offshore borrowing by banks and non-bank private corporations in later years. Because of the potential risks they entail, these capital inflows were almost from the inception considered as posing a serious challenge for macroeconomic management, leading the profession to coin the term "the capital inflow problem".

An important aspect of the capital inflow episode was that the volume of inflows far exceeded the current account deficits, such that the increases in foreign exchange reserves amounted to between 25 and 35 percent of the net capital inflows. Needless to say, the accumulation of reserves was the result of foreign exchange market intervention to maintain the level of nominal exchange rates. Short of allowing the exchange rate to appreciate, the East Asian monetary authorities responded decisively to the massive reserve inflows, first by the conventional form of sterilization and then by taking a wide range of measures to limit the effect of the reserve inflows on the growth of monetary aggregates, the measures which are called "broadly defined sterilization" in the chpater.

We began the chapter by noting that, whether narrowly or broadly defined, effective sterilization should not only limit the growth of monetary aggregates in response to an increase in foreign assets, but also raise the level of domestic interest rates. The resulting tight monetary condition (often supported by tight fiscal policy) and higher domestic interest rates should then promote additional capital inflows. The Granger causality tests and OLS estimates of structural parameters, however, gave the somewhat perplexing results indicating that, while sterilization was apparently effective in fully limiting the growth of monetary aggregates arising from an increase in foreign assets, it was not causing the level of interest rates to rise.

At this point, a word of reservation might be expressed about the nature of the methodologies used. We noted at the outset that, given the variety of tools used to mitigate the impact of reserve inflows on the growth of monetary aggregates in these countries, the conventional method of estimating the offset coefficient of the capital flow equation along with the monetary policy reaction function would be inappropriate as a test of the effectiveness of sterilization. Instead, what we decided to do was to use a "black box" way of measuring the effectiveness of sterilization by essentially estimating the statistical significance of FA in the equation describing the growth of

M1 or M2, without explicitly considering how sterilization is actually effected. While we believe that this is an intuitively appealing procedure, given the ultimate objective of sterilization, we also recognize that it may be subject to potential problems. For instance, the lack of statistical significance may reflect, not the effectiveness of sterilization, but the much smaller magnitude of FA relative to that of either M1 or M2; the results may also be sensitive to the choice of lag length, particularly when the methodologies are applied in first difference form. In the future, it will be useful to check the robustness of our methodologies against alternative specifications or alternative sample countries.[12]

Subject to these and other limitations, our result (suggesting the effectiveness of sterilization, while indicating little evidence of an interest rate rise) are capable of yielding two possible interpretations. First and most likely, the lack of evidence linking a rise in foreign assets to a rise in interest rates may simply suggest that sterilization was effective, not necessarily in raising the level of interest rates, but in keeping it from falling toward the lower world interest rates. To support this claim, the moving average representations of the estimated VAR system (reported earlier) suggest that interest rates do rise in response to an innovation in foreign assets in all countries except Korea (Figure 4.6). It is also possible that a more systematic relationship between foreign assets and interest rates might have been evident for a more appropriate interest rate or interest rate *differential*. In Indonesia, for example, it is said that the interest rate on SBIs rose sharply from 11.6 percent in 1988 to 18.8 percent in 1990 and 21.5 percent in 1991; Furman and Stiglitz (1998) note that interest rate differentials did widen in East Asian countries over the period of sterilization.[13]

[12] In this context, as a robustness check, the referee has suggested the usefulness of applying our methodologies to countries under currency boards. Data limitations, however, have prevented us from pursuing this course.

[13] According to Furman and Stiglitz (1998), in Thailand, short-term money market rates rose

Second, as another possible interpretation of the seeming lack of evidence on the interest rate channel, it is possible that sterilization was not so effective in limiting the growth of overall monetary assets, although it was effective in limiting the growth of M1 or M2 which is under the supervised banking sector. Although broadly defined sterilization measures, such as changes in reserve requirements, credit controls and moral suasion may be effective against the supervised banking sector, they may result in disintermediation in an environment where there is a viable nonbank financial sector. In the case of Korea, for example, Spiegel (1995) documents that the share of assets controlled by the banking sector declined over the period 1986-93, although no such evidence was found for the Philippines and Malaysia, where the nonbank financial sector is not well-developed. It should be noted that this disintemediation interpretation is not necessarily incompatible with the story that sterilization kept the level of interest rates high.

In either case, the policy of sterilization pursued by the monetary authorities of East Asia during the capital inflow episode was effective in fully limiting the growth of M1 or M2, and possibly magnified the risk of capital inflows by keeping the level of interest rates high (hence promoting additional capital inflows), by channelling resources to the relatively unsupervised nonbank financial sector, or both. In this context, the work of Montiel and Reinhart (1999) suggest that the sterilization policy of the Asian monetary authorities not only magnified the volume of capital inflows but also skewed the composition of capital flows towards short-term maturities. Both through additional capital inflows with a short-term bias and through possible disintermediation, it is likely that the capital inflow problem of East Asia leading up to the crisis of 1997 was made more serious by the active and persistent policy of

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400 basis points above comparable US interest rates in 1996, and similar spreads were observed for

sterilization.

Appendix. Sources of Data

Except for industrial production in Indonesia and Thailand (which were obtained from the Bank of Japan's economic database), all data were obtained from the International Monetary Fund, *International Financial Statistics*, as follows. Foreign Assets (FA) were obtained from line 11. Narrow money (M1) and quasi-money were obtained from lines 34 and 35, respectively; M1 and quasi-money constitute broad money (M2). Interest rates were obtained from the money market rate (line 60b) for Indonesia, Korea, Malaysia and Thailand, and from the Treasury bill rate (line 60c) for the Philippines.

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other East Asian countries.

Table 4.1.
Augmented Dickey-Fuller Statistics, 1987-97

	variables	seasonal dummies	time trend
Indonesia:	ln M1	-0.320 (4) [0.922]	-2.904 (4) [0.160]
	ln M2	-0.920 (3) [0.781]	-2.009 (3) [0.596]
	ln FA	-0.112 (3) [0.948]	-2.666 (4) [0.250]
	ln (M1/P)	-0.741 (2) [0.835]	-2.642 (2) [0.260]
	ln (M2/P)	-1.009 (3) [0.749]	-2.090 (3) [0.552]
	ln (FA/P)	-0.154 (3) [0.943]	-2.554 (4) [0.301]
	ln Y	-0.312 (4) [0.923]	-2.544 (4) [0.306]
	i	-2.603 (2) [0.278]	-2.603 (2) [0.278]
Korea:	ln M1	-1.686 (3) [0.438]	0.171 (4) [0.995]
	ln M2	-2.240 (4) [0.191]	-1.269 (4) [0.895]
	ln FA	-0.289 (4) [0.926]	-3.087 (4) [0.110]
	ln (M1/P)	-1.367 (4) [0.597]	-0.516 (4) [0.982]
	ln (M2/P)	0.038 (4) [0.961]	-2.725 (4) [0.225]
	ln (FA/P)	-0.544 (4) [0.883]	-2.986 (4) [0.144]
	ln Y	-0.661 (4) [0.856]	-2.697 (3) [0.237]
	i	-2.079 (2) [0.557]	-2.079 (2) [0.557]
Malaysia:	ln M1	0.499 (2) [0.984]	-2.396 (2) [0.381]
	ln M2	1.155 (2) [0.995]	-2.992 (2) [0.134]
	ln FA	-0.957 (3) [0.768]	-1.583 (3) [0.798]
	ln (M1/P)	0.220 (4) [0.973]	-2.376 (2) [0.392]
	ln (M2/P)	1.014 (2) [0.994]	-2.778 (2) [0.204]
	ln (FA/P)	-1.043 (3) [0.736]	-1.597 (3) [0.793]
	ln Y	0.520 (4) [0.985]	-2.191 (4) [0.494]
	i	-1.918 (3) [0.644]	-1.918 (3) [0.644]
Philippines:	ln M1	0.070 (4) [0.964]	-1.617 (4) [0.785]
	ln M2	-1.643 (2) [0.460]	-2.071 (4) [0.562]
	ln FA	-0.217 (2) [0.936]	-2.280 (2) [0.444]
	ln (M1/P)	1.839 (4) [0.998]	0.169 (4) [0.995]
	ln (M2/P)	-0.842 (2) [0.806]	-1.491 (4) [0.831]
	ln (FA/P)	-0.261 (2) [0.930]	-2.679 (2) [0.244]
	ln Y	0.651 (4) [0.988]	-2.131 (4) [0.528]
	i	-2.008 (2) [0.596]	-2.008 (2) [0.596]
Thailand:	ln M1	-0.981 (2) [0.760]	-2.197 (4) [0.491]
	ln M2	-2.413 (3) [0.137]	-1.708 (4) [0.747]
	ln FA	-2.917 (3) [0.043]**	0.324 (3) [0.996]
	ln (M1/P)	-1.112 (2) [0.709]	-2.135 (4) [0.526]
	ln (M2/P)	-2.309 (2) [0.168]	-1.071 (4) [0.933]
	ln (FA/P)	-2.948 (3) [0.039]**	0.441 (3) [0.990]
	ln Y	-2.056 (4) [0.262]	-3.736 (4) [0.022]**
	i	-2.069 (2) [0.563]	-2.069 (2) [0.563]

Notes: The figures are augmented Dickey-Fuller statistics obtained from running a regression with a constant term and seasonal dummies (left column) or with a constant term and time trend (right column); for the interest rate only, neither seasonal dummy nor time trend is included (hence, the same statistics are reported in both columns). Lag length was chosen on the basis of Schwarz's Bayesian information criterion (SBIC). ** indicates that the statistic is significant (*i.e.,* the existence of a unit root can be rejected) at the 5 percent level; the figures in parentheses denote lag length; and those in brackets are p-values.

Table 4.2.
Tests of Cointegration Between Money and Foreign Assets, 1987-97

cointegration vectors (r)		null $r = 0$	$r \leq 1$	
Bivariate: M1 and FA (first row); M2 and FA (second row)				
Indonesia:	VAR (1)	11.90 [0.286]	0.337 [0.770]	r=0
	VAR (1)	22.99 [0.012]**	1.705 [0.603]	r=1
Korea:	VAR (3)	8.017 [0.622]	2.955 [0.424]	r=0
	VAR (3)	6.574 [0.735]	1.976 [0.565]	r=0
Malaysia:	VAR (1)	5.096 [0.826]	0.019 [0.801]	r=0
	VAR (1)	6.369 [0.749]	1.246 [0.665]	r=0
Philippines:	VAR (1)	6.036 [0.771]	0.005 [0.803]	r=0
	VAR (1)	6.958 [0.707]	0.006 [0.803]	r=0
Thailand:	VAR (1)	13.59 [0.180]	1.539 [0.626]	r=0
	VAR (1)	11.12 [0.347]	0.555 [0.747]	r=0
Multivariate: real M1, real FA, output and interest rate (first row); real M2, real FA, output and interest rate (second row)				
Indonesia:	VAR (1)	39.38 [0.284]	19.43 [0.544]	r=0
	VAR (1)	44.32 [0.119]	25.80 [0.184]	r=0
Korea:	VAR (1)	36.23 [0.432]	10.25 [0.930]	r=0
	VAR (1)	41.13 [0.217]	14.29 [0.821]	r=0
Malaysia:	VAR (1)	38.60 [0.318]	18.52 [0.602]	r=0
	VAR (1)	39.20 [0.292]	19.82 [0.518]	r=0
Philippines:	VAR (1)	55.73 [0.009]***	19.85 [0.516]	r=1
	VAR (1)	38.87 [0.400]	18.95 [0.575]	r=0
Thailand:	VAR (1)	39.87 [0.264]	21.27 [0.424]	r=0
	VAR (1)	45.10 [0.102]	21.65 [0.399]	r=0

Notes: Johansen's trace tests on a VAR system with a constant term and seasonal dummies; lag length (in parentheses) is chosen on the basis of Schwarz's Bayesian information criterion (SBIC); r denotes the number of cointegrating vectors; *** and ** indicate that the statistic is significant at the 1 and 5 percent levels; and the figures in brackets are p-values.

Table 4.3.
Granger Tests of Causality between Money and Foreign Assets, 1987-97
(Bivariate VAR)

		FA causes M		M causes FA	
M1 and FA (first row); M2 and FA (second row):					
Indonesia:					
	VAR (1)	F (1,34)	0.388 [0.537]	F (1,34)	0.337 [0.565]
	VAR (1)	F (1,34)	1.132 [0.295]	F (1,34)	0.022 [0.882]
Korea:					
	VAR (1)	F (1,34)	1.421 [0.241]	F (1,34)	0.474 [0.496]
	VAR (1)	F (1,34)	1.638 [0.209]	F (1,34)	0.231 [0.634]
Malaysia:					
	VAR (1)	F (1,34)	4.035 [0.053]*	F (1,34)	1.455 [0.236]
	VAR (1)	F (1,34)	0.000 [0.991]	F (1,34)	0.480 [0.493]
Philippines:					
	VAR (1)	F (1,34)	0.324 [0.573]	F (1,34)	4.622 [0.039]**
	VAR (1)	F (1,34)	3.146 [0.085]*	F (1,34)	1.349 [0.254]
Thailand:					
	VAR (1)	F (1,34)	0.039 [0.845]	F (1,34)	0.673 [0.418]
	VAR (1)	F (1,34)	1.077 [0.307]	F (1,34)	4.315 [0.045]**

Notes: F-statistics in a bivariate VAR of money (M1 or M2) and foreign assets (FA) with a constant term and seasonal dummies; lag length (in parentheses following VAR) was chosen on the basis of Schwarz's Bayesian information criterion (SBIC); p-statistics are in brackets; ** and * indicate that the statistic is significant at the 5 and 10 percent levels, respectively.

Table 4.4.
Granger Tests of Causality between Money and Foreign Assets, 1987-97
(Multivariate VAR)

		FA causes M	M causes FA
M1, Y, i and FA (first row); M2, Y, i and FA (second row):			
Indonesia:			
	VAR (1)	F (1,28) 0.000 [0.975]	F (1,28) 0.073 [0.788]
	VAR (1)	F (1,28) 0.002 [0.963]	F (1,28) 0.038 [0.845]
Korea:			
	VAR (1)	F (1,28) 1.235 [0.275]	F (1,28) 0.641 [0.429]
	VAR (1)	F (1,28) 1.191 [0.256]	F (1,28) 0.432 [0.515]
Malaysia:			
	VAR (1)	F (1,28) 1.520 [0.227]	F (1,28) 1.407 [0.245]
	VAR (1)	F (1,28) 0.093 [0.762]	F (1,28) 0.005 [0.942]
Philippines:			
	VAR (1)	F (1,28) 0.531 [0.472]	F (1,28) 6.674 [0.015]**
	VAR (1)	F (1,28) 2.048 [0.163]	F (1,28) 1.242 [0.274]
Thailand:			
	VAR (1)	F (1,28) 1.298 [0.264]	F (1,28) 2.351 [0.136]
	VAR (1)	F (1,28) 0.303 [0.586]	F (1,28) 0.918 [0.346]

Notes: F-statistics in a multivariate VAR of real money (M1 or M2), real foreign assets (FA), output and the interest rate, with a constant term and seasonal dummies; lag length (in parentheses following VAR) was chosen on the basis of Schwarz's Bayesian information criterion (SBIC); p-statistics are in brackets; ** indicates that the statistic is significant at the 5 percent level.

Table 4.5.
Granger Tests of Causality between Foreign Assets and the Interest Rate, 1987-97
(Multivariate VAR)

		FA causes i	i causes FA
M1, Y, i and FA (first row); M2, Y, i and FA (second row):			
Indonesia:			
	VAR (1)	F (1,28) 2.251 [0.144]	F (1,28) 2.009 [0.167]
	VAR (1)	F (1,28) 1.791 [0.191]	F (1,28) 1.933 [0.175]
Korea:			
	VAR (1)	F (1,28) 0.251 [0.619]	F (1,28) 0.011 [0.913]
	VAR (1)	F (1,28) 0.134 [0.716]	F (1,28) 0.011 [0.915]
Malaysia:			
	VAR (1)	F (1,28) 0.239 [0.628]	F (1,28) 0.428 [0.517]
	VAR (1)	F (1,28) 0.775 [0.386]	F (1,28) 0.234 [0.631]
Philippines:			
	VAR (1)	F (1,28) 12.27 [0.002]***	F (1,28) 2.217 [0.145]
	VAR (1)	F (1,28) 8.765 [0.006]***	F (1,28) 1.569 [0.220]
Thailand:			
	VAR (1)	F (1,28) 0.268 [0.608]	F (1,28) 2.692 [0.112]
	VAR (1)	F (1,28) 0.546 [0.465]	F (1,28) 0.327 [0.571]

Notes: F-statistics in a multivariate VAR of real money (M1 or M2), real foreign assets (FA), output and the interest rate, with a constant term and seasonal dummies; lag length (in parentheses following VAR) was chosen on the basis of Schwarz's Bayesian information criterion (SBIC); p-statistics are in brackets; *** indicates that the statistic is significant at the 1 percent level.

Table 4.6.
Indonesia: Money Supply Adjustment, 1987-97

	real narrow money (M1)				real broad money (M2)			
Constant	0.046*** (3.249)	0.045*** (3.180)	0.078* (1.930)	0.044*** (3.135)	0.054*** (3.661)	0.052*** (3.506)	0.060 (1.425)	0.051*** (3.498)
Output	0.284** (2.284)	0.285** (2.290)	0.270** (2.103)	0.297** (2.384)	0.028 (0.224)	0.029 (0.225)	0.040 (0.296)	0.047 (0.364)
Interest Rate	-0.003 (-0.936)		-0.003 (-0.884)		-0.005 (-1.349)		-0.001 (-0.222)	
Lagged Real Foreign Assets	-0.068 (-0.783)	-0.049 (-0.586)	0.021 (0.198)	0.021 (0.198)	-0.036 (-0.403)	-0.009 (-0.097)	0.093 (0.825)	0.093 (0.839)
DUM*Lagged Real Foreign Assets			-0.104 (-0.615)	-0.166 (-1.078)			-0.222 (-1.250)	-0.238 (-1.494)
F-statistic	1.912*	2.136*	1.800	1.983*	0.909	0.711	0.828	0.986
R-squared	0.259	0.239	0.282	0.265	0.142	0.094	0.154	0.152
Durbin-Watson	1.764	1.717	1.715	1.703	2.666	2.714	2.696	2.720

Notes: Coefficients are estimated in first difference form by ordinary least squares (OLS); ***, ** and * indicate that the statistic is significant at the 1, 5, and 10 percent levels, respectively; coefficients for the seasonal dummies are not reported; except for Korea, DUM is a slope dummy for intense sterilization; and t-values are in parentheses.

Table 4.7.
Korea: Money Supply Adjustment, 1987-97

	real narrow money (M1)		real broad money (M2)	
Constant	0.064 (1.203)	0.064 (1.226)	0.053*** (3.394)	0.054*** (3.459)
Output	0.280 (0.879)	0.278 (0.897)	0.064 (0.681)	0.062 (0.679)
Interest Rate	0.0003 (0.033)		0.0003 (0.134)	
Lagged Real Foreign Assets	0.120 (1.303)	0.120 (1.322)	0.040 (1.466)	0.040 (1.484)
F-statistic	10.97***	16.13***	16.13***	19.93***
R-squared	0.666	0.746	0.746	0.746
Durbin-Watson	2.680	2.738	2.738	2.743

Notes: See Table 4.6.

Table 4.8.
Malaysia: Money Supply Adjustment, 1987-97

	real narrow money (M1)				real broad money (M2)			
Constant	0.068***	0.066***	0.067***	0.064***	0.049***	0.048***	0.048***	0.047***
	(6.067)	(6.094)	(5.651)	(5.666)	(4.070)	(4.128)	(3.813)	(3.862)
Output	-0.100	-0.127	-0.054	-0.080	-0.053	-0.070	-0.033	-0.049
	(-0.421)	(-0.542)	(-0.216)	(-0.323)	(-0.208)	(-0.283)	(-0.123)	(-0.188)
Interest Rate	-0.009		-0.009		-0.006		-0.006	
	(-0.807)		(-0.788)		(-0.487)		(-0.475)	
Lagged Real Foreign Assets	0.025	0.046	0.008	0.028	0.022	0.035	0.014	0.027
	(0.424)	(0.844)	(0.115)	(0.447)	(0.338)	(0.603)	(0.195)	(0.408)
DUM*Lagged Real Foreign Assets			0.086	0.088			0.038	0.039
			(0.663)	(0.680)			(0.269)	(0.281)
F-statistic	2.921**	3.41**	2.524**	2.874**	1.314	1.564	1.105	1.282
R-squared	0.347	0.334	0.356	0.343	0.192	0.187	0.195	0.189
Durbin-Watson	2.514	2.434	2.433	2.356	1.882	1.862	1.892	1.872

Notes: See Table 4.6.

Table 4.9.
The Philippines: Money Supply Adjustment, 1987-97

	real narrow money (M1)				real broad money (M2)			
Constant	0.069 (1.109)	0.071 (1.147)	0.070 (1.129)	0.071 (1.161)	0.004 (0.063)	0.012 (0.175)	0.004 (0.063)	0.012 (0.171)
Output	0.620 (1.589)	0.600 (1.569)	0.626 (1.628)	0.615 (1.633)	0.643* (1.639)	0.536 (1.294)	0.645 (1.610)	0.543 (1.302)
Interest Rate	-0.001 (-0.412)		-0.001 (-0.251)		-0.008** (-2.175)		-0.008** (-2.079)	
Lagged Real Foreign Assets	0.028 (0.542)	0.037 (0.810)	0.071 (1.206)	0.077 (1.475)	-0.108** (-2.090)	-0.059 (-1.201)	-0.093 (-1.509)	-0.038 (-0.652)
DUM*Lagged Real Foreign Assets			-0.134 (-1.411)	-0.137 (-1.471)			-0.048 (-0.481)	-0.071 (-0.694)
F-statistic	34.60***	42.52**	30.83***	37.01***	7.836***	7.621***	6.593***	6.334***
R-squared	0.863	0.862	0.871	0.871	0.588	0.528	0.591	0.535
Durbin-Watson	2.534	2.563	2.442	2.466	2.139	2.180	2.109	2.168

Notes: See Table 4.6.

Table 4.10.
Thailand: Money Supply Adjustment, 1987-97

	real narrow money (M1)				real broad money (M2)			
Constant	0.094***	0.093***	0.090***	0.090***	0.048***	0.048***	0.047***	0.047***
	(4.917)	(4.949)	(4.681)	(4.713)	(7.805)	(7.585)	(7.522)	(7.307)
Output	0.078	0.071	0.076	0.070	0.072	0.080	0.072	0.080
	(0.378)	(0.352)	(0.372)	(0.345)	(1.108)	(1.176)	(1.098)	(1.164)
Interest Rate	-0.002		0.002		-0.003**		-0.003**	
	(-0.532)		(0.527)		(-2.002)		(-2.000)	
Lagged Real Foreign Assets	-0.086	-0.090	0.131	0.126	0.038	0.045	0.092	0.097
	(-0.728)	(-0.784)	(0.575)	(0.562)	(1.021)	(1.145)	(1.255)	(1.275)
DUM*Lagged Real Foreign Assets			-0.262	-0.263			-0.064	-0.064
			(-1.110)	(-1.126)			(-0.853)	(-0.804)
F-statistic	12.71***	15.52**	11.14***	13.25***	6.000***	5.879***	5.205***	4.956***
R-squared	0.698	0.695	0.709	0.707	0.522	0.464	0.532	0.474
Durbin-Watson	2.254	2.276	2.375	2.402	1.729	1.852	1.745	1.854

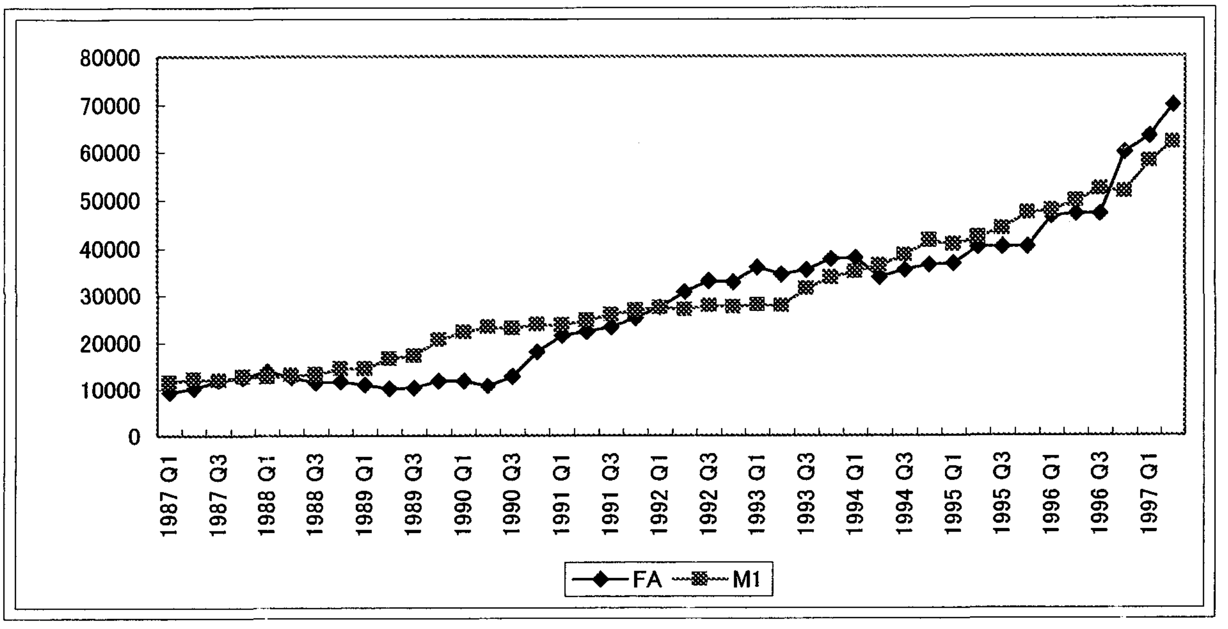
Notes: See Table 4.6.

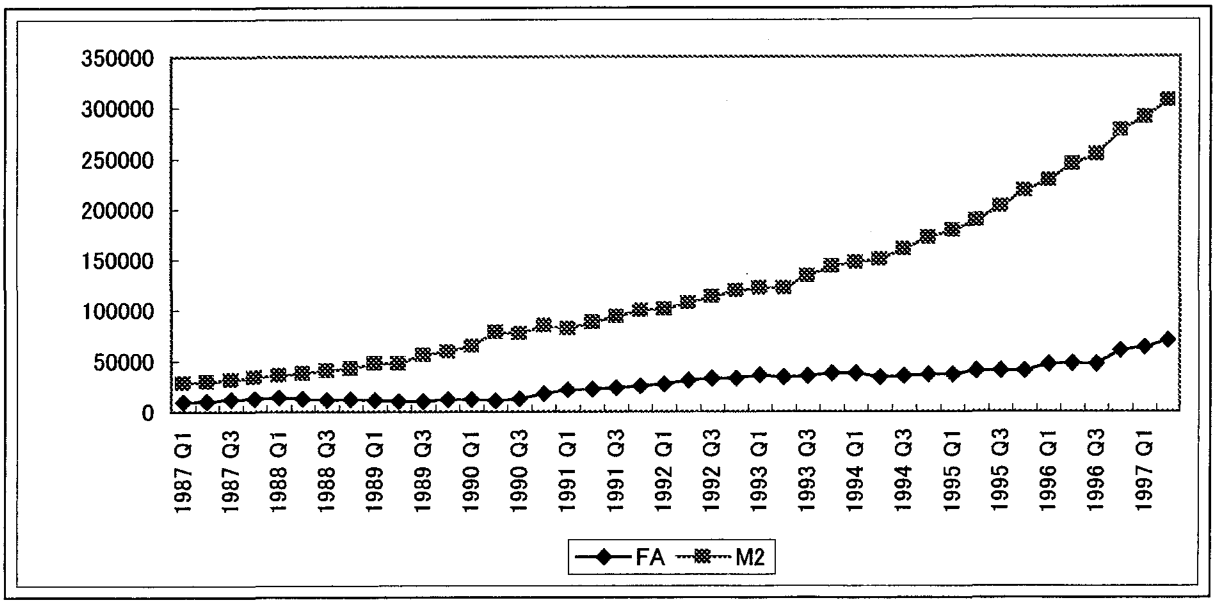
Table 4.11.
The "Intense Sterilization" Dummy

	Indonesia	Korea	Malaysia	Philippines	Thailand
1987	0	0	0	0	0
1988	0	0	0	0	0
1989	0	0	0	0	0
1990	0	0	0	1	1
1991	1	0	1	0	1
1992	1	0	1	0	0
1993	0	0	1	0	0
1994	0	0	0	0	0
1995	0	0	0	0	1
1996	1	0	1	0	0
1997	0	0	0	0	0

Note: The authors' judgement based on Reinhart and Reinhart (1998), Montiel and Reinhart (1999) and Villanueva and Seng (1999). Sterilization is considered intense (*i.e.,* a value of unity is assigned) if open market operations were large in scale and accompanied by increased reserve requirements or transfers of government deposits from commercial banks to the central bank. The quarterly series for a given year are assumed to have the same value as the annual series.

Figure 4.1.
Indonesia: Foreign Assets and Monetary Aggregates, 1987-97
(in billions of rupiah)



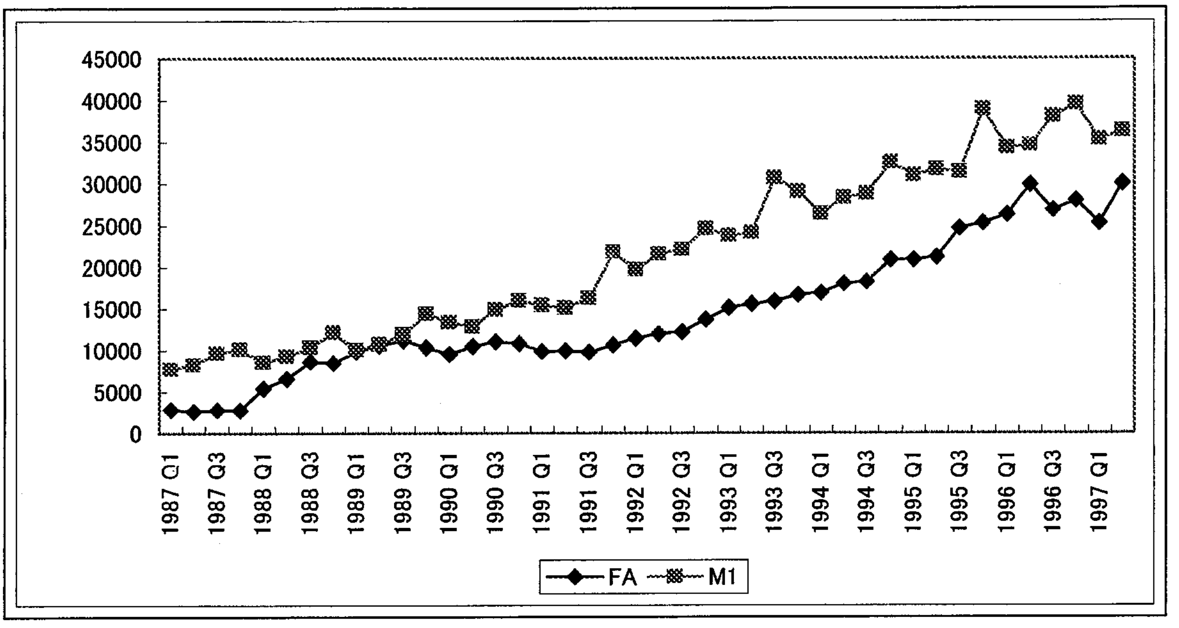


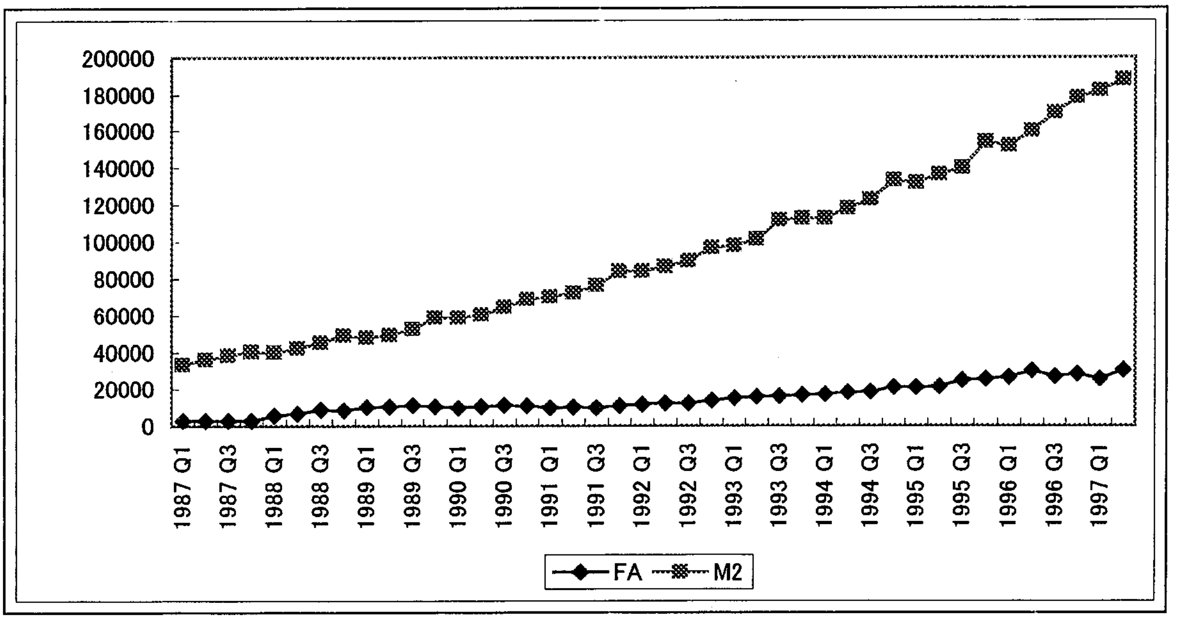
Source: IMF, *International Financial Statistics.*

Figure 4.2.

Korea: Foreign Assets and Monetary Aggregates, 1987-97

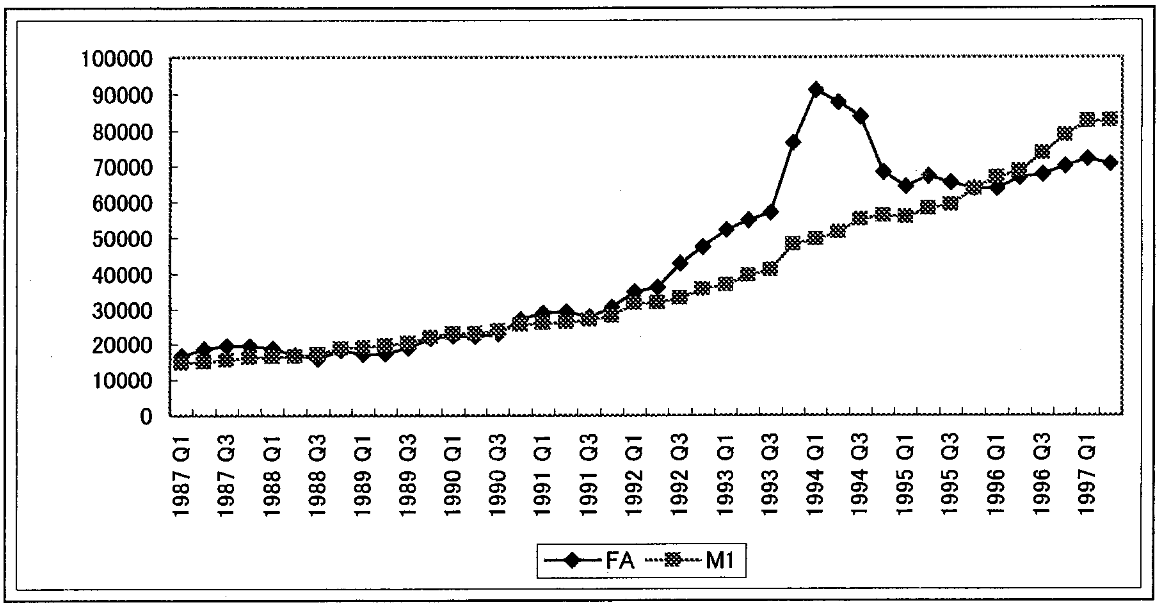
(in billions of won)

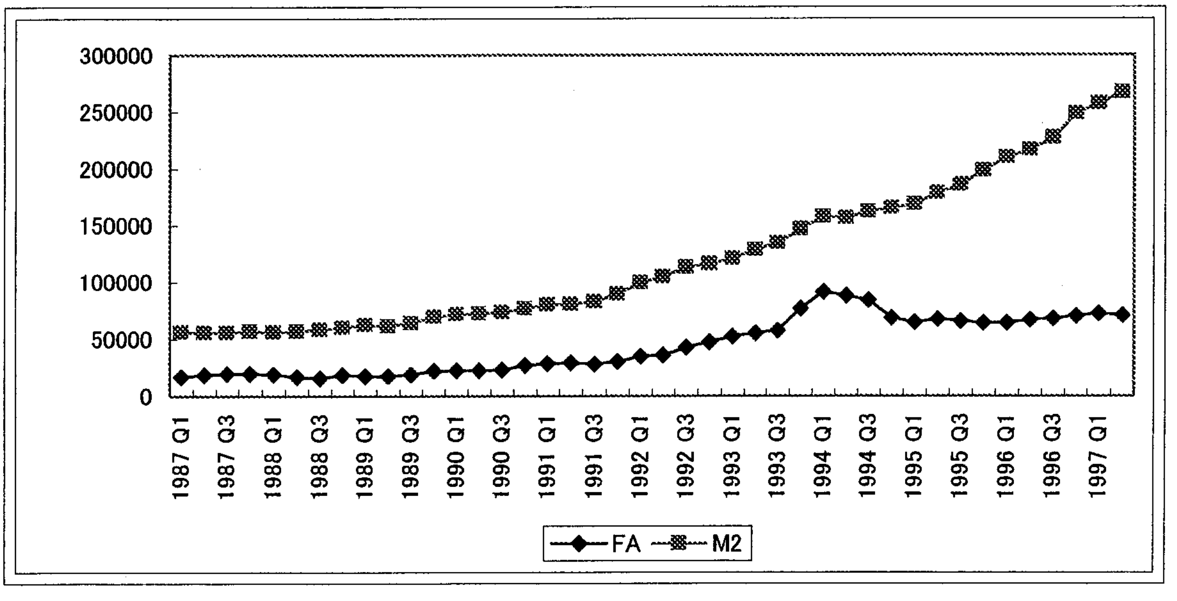




Source: IMF, *International Financial Statistics*.

Figure 4.3.
Malaysia: Foreign Assets and Monetary Aggregates, 1987-97
(in millions of ringgit)



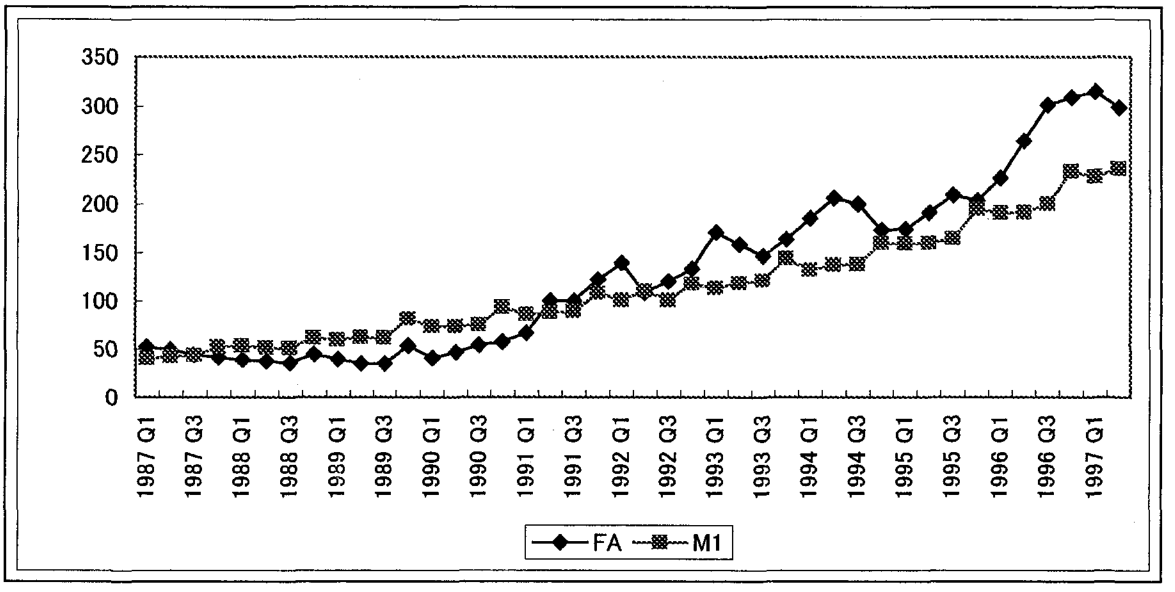


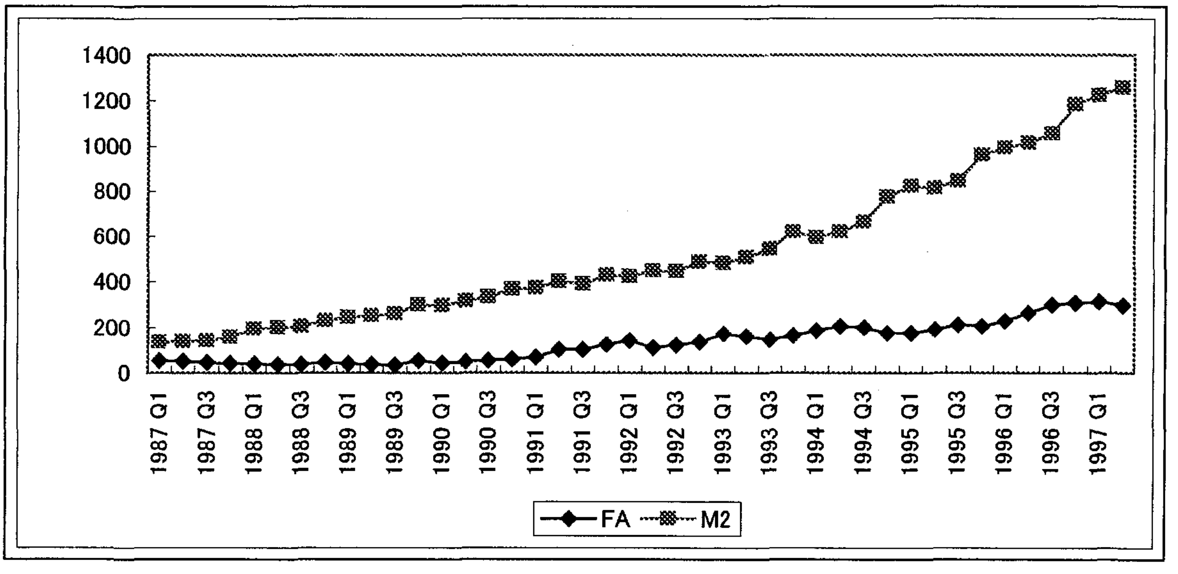
Source: IMF, *International Financial Statistics*.

Figure 4.4.

The Philippines: Foreign Assets and Monetary Aggregates, 1987-97

(in billions of pesos)



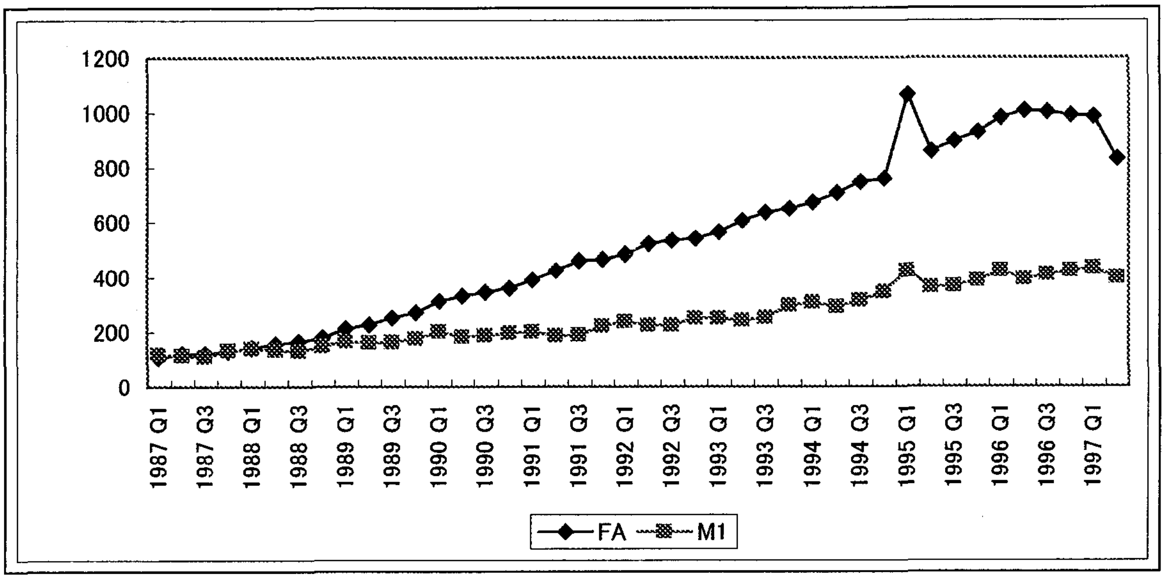


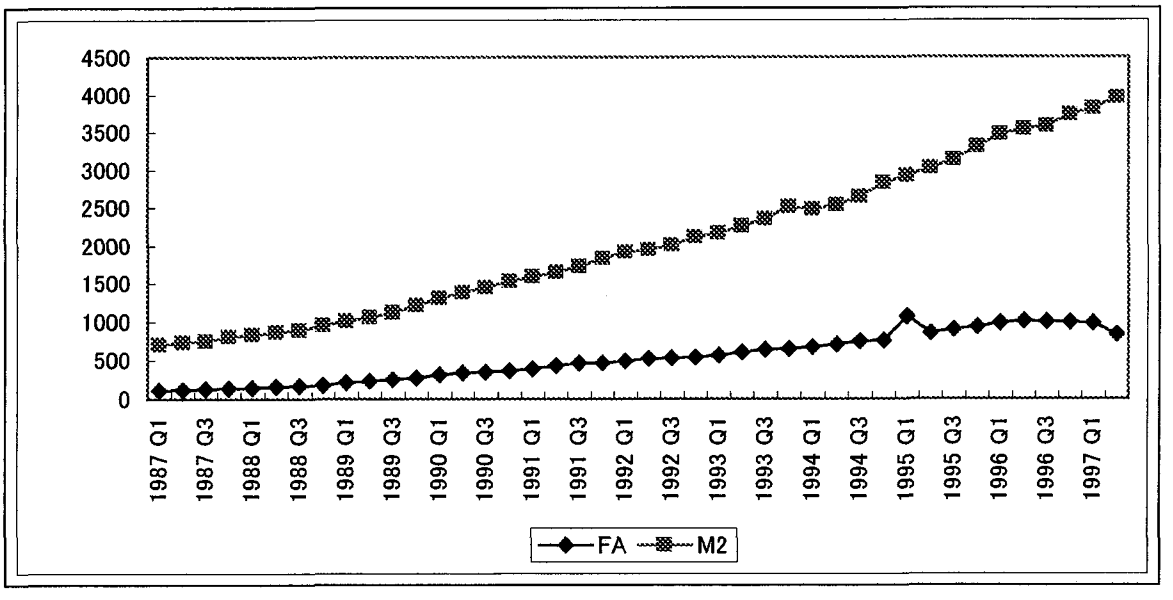
Source: IMF, *International Financial Statistics.*

Figure 4.5.

Thailand: Foreign Assets and Monetary Aggregates, 1987-97

(in billions of baht)

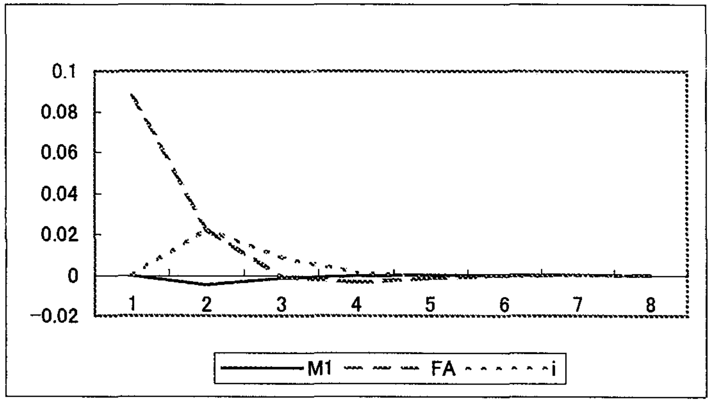


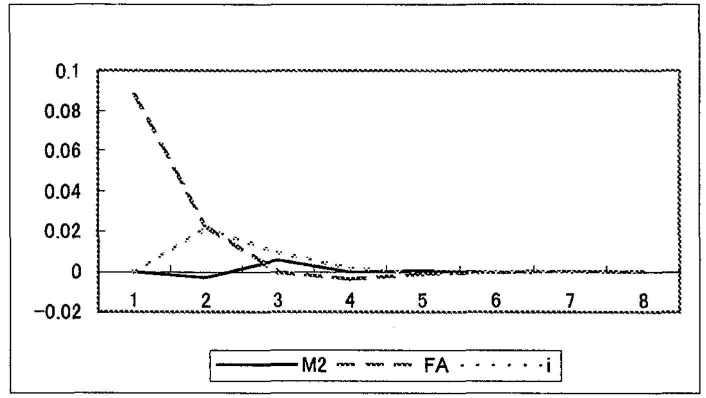


Source: IMF, *International Financial Statistics.*

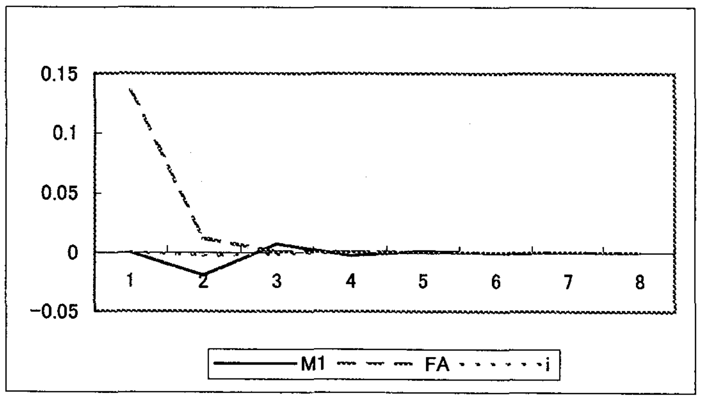
Figure 4.6.
The Responses of Monetary Aggregates and the Interest Rate to an Innovations in Foreign Assets

(a) Indonesia





(b) Korea



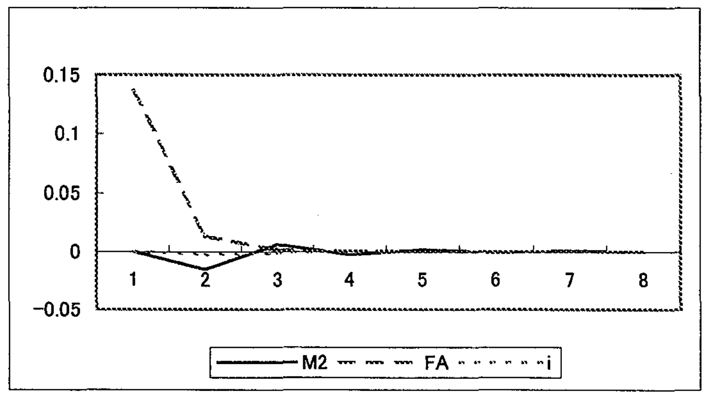
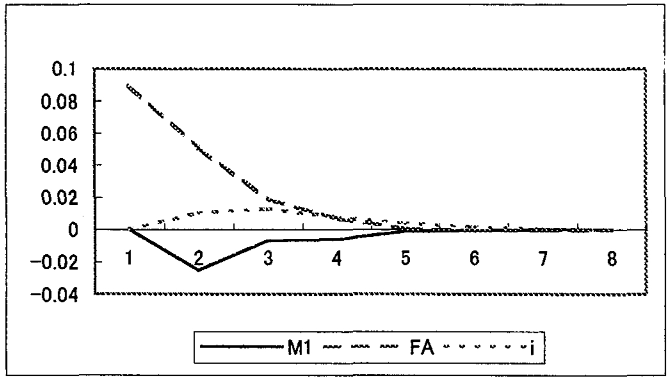
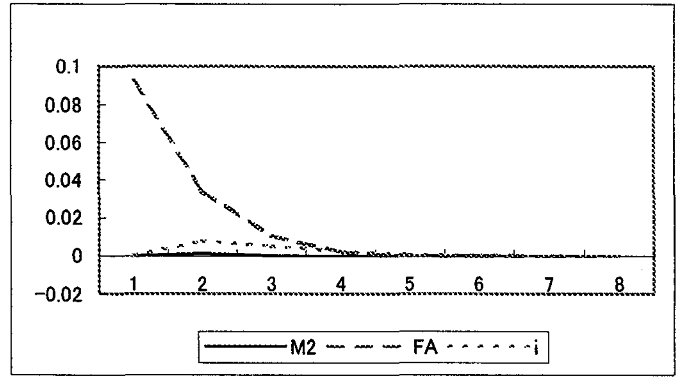


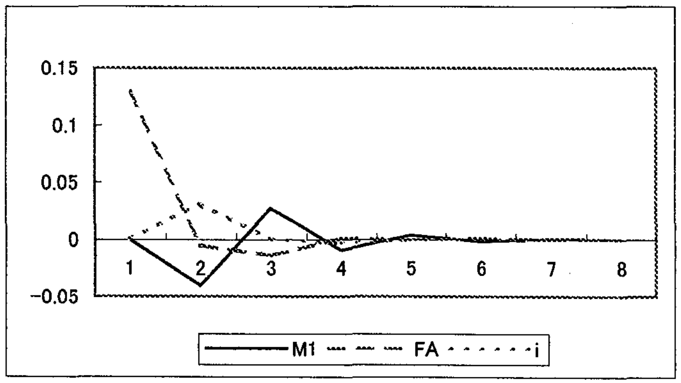
Figure 4.6 (Continued).

(c) Malaysia





(d) The Philippines



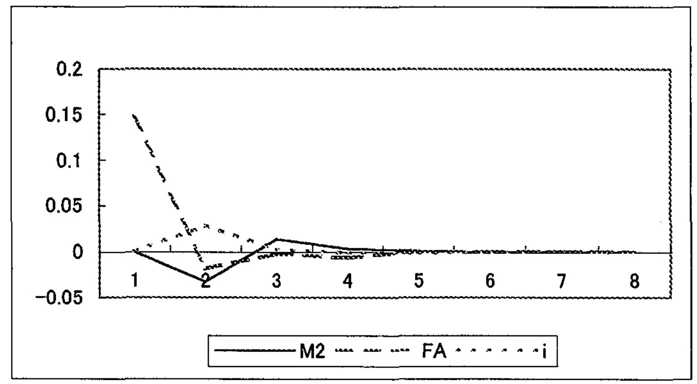
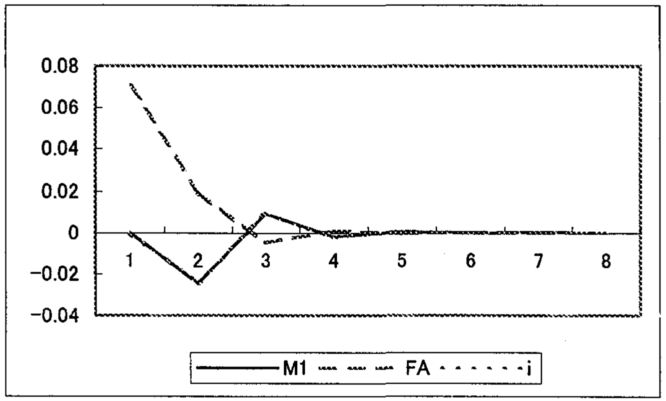
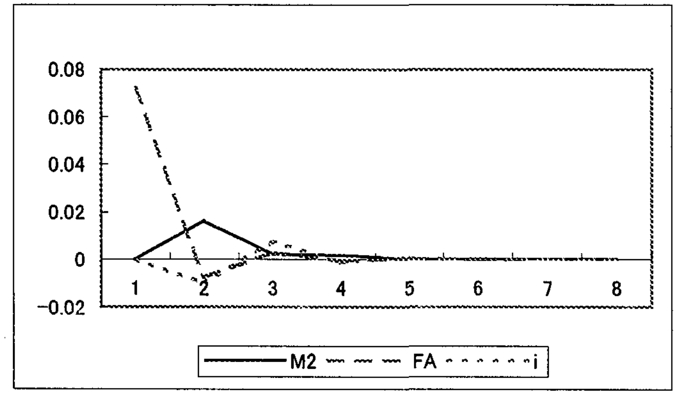


Figure 4.6 (Continued).

(e) Thailand





Notes: The impulse responses of real money (M1 or M2) and the interest rate (i) to a one standard deviation innovation in real foreign assets (FA) in a multivariate VAR system consisting of real M1 or M2, real FA , Y and i, with a constant term and seasonal dummies. In the case of the Philippines when M1 was used, a vector error correction model (VECM) was applied, given the possible presence of cointegration.

5. Risk Premiums and Exchange Rate Expectations: A Reassessment of the So-Called Dollar Peg Policies of Crisis East Asian Countries, 1994-97

5.1. Introduction

This chapter will use monthly interest rate data to assess the credibility of the so-called dollar peg policies of the four crisis East Asian countries (*i.e.,* Indonesia, Korea, Malaysia and Thailand) during the 42-month period immediately preceding the onset of the East Asian currency crisis in July 1997. In particular, it will use an unobserved components model to extract foreign exchange risk premiums from interest rate differentials with respect to US dollar-denominated assets, in order to obtain the estimates of expected exchange rate depreciation or appreciation, hence the degree of credibility of the policies of maintaining relative stability against the US dollar. Because the presence of an exchange rate premium, if any, suggests that East Asian currency-denominated assets and US dollar-denominated assets are imperfect substitutes, moreover, this exercise will also help assess the effectiveness of sterilization policies pursued in these countries in controlling the growth of monetary aggregates under pegged exchange rate regimes (see Chapter 4). It turns out that, during the period under consideration, (1) there was evidence of time-varying exchange risk premiums between the East Asian currencies and the US dollar (suggesting that East Asian currency assets and US dollar assets were imperfect substitutes, hence sterilization was effective) and that (2) markets apparently displayed systematic expectations of exchange rate depreciation or appreciation (suggesting that the so-called dollar peg policies were far from being credible).

It is well known that, from the latter part of the 1980s into early 1997, the emerging market economies of East Asia received a large volume of capital inflows (Calvo, Leiderman and Reinhart (1996), Chuhan, Claessens and Mamingi (1998) and Montiel (1998)). In fact, the inflows were massive indeed: in terms of GDP, the volume of cumulative inflows from 1988 to 1995, for example, amounted to 51.5 percent in Thailand, 45.8 percent in Malaysia, 9.3 percent in Korea and 8.3 percent in Indonesia (see Chapter 4). While responsible for this surge of capital inflows were both internal (or "pull") and external (or "push") factors of various types, the substantial interest rate differentials that existed in favor of assets denominated in East Asian currencies over those denominated in major industrial country currencies were undoubtedly an important contributing factor. For example, the average interest rate differentials favoring short-term money market instruments denominated in East Asian currencies over those denominated in the US dollar during January 1994-June 1997 were over 8 percent for Indonesia, about 8 percent for Korea about one percent for Malaysia and over 6 percent for Thailand (top panel, Table 5.1; and Figure 5.1).[1] Adjusted for actual exchange rate changes, the average excess returns over US dollar-denominated instruments remained substantial, amounting to 2.5-6.1 percent per year (bottom panel, Table 5.1).

In this connection, it is important to remember that these large positive excess returns on East Asian currency assets were observed against the background of the so-called dollar peg policies, the exchange rate policies of maintaining relative stability against the US dollar (Ito, Ogawa and Sasaki (1998), Takagi (1999) and Chapter 3). The presence of interest rate differentials favoring East Asian currency assets in this

[1] Negative interest rate differentials were observed for the Malaysian ringgit during the

environment means that there were risk premiums, expected depreciation, or some combination of both. Equivalently, the presence of positive ex post excess returns means that there were risk premiums, unexpected appreciation, or some combination of both on the part of East Asian currency assets. If the exchange rate policies had been credible in the sense that market participants expected the US dollar exchange rates to remain stable, one would have observed the expected rate of currency depreciation to be small. Then, most of the interest rate differentials would be explained by foreign exchange risk premiums. On the other hand, lack of credibility in the dollar peg policies would have meant that an important component of the interest rate differentials would reflect expected exchange rate change. An important task, therefore, is to decompose the observed interest rate differentials into risk premiums and expected rates of depreciation, which are both unobservable.

In attempting to make the decomposition, we will use the relatively simple statistical procedure called an unobserved components model, as previously employed by Wolff (1987, 2000a, 2000b), Nijman, Palm and Wolff (1993), and Cheung (1993). An important advantage of this approach is that it absolves us from making strong assumptions about the fundamental determinants of the risk premium, as would be required in the alternative regression-based approach. At a minimum, the regression-based approach would require the availability of consumption and other macroeconomic data and might even require us to specify the explicit form of utility functions (see Engel (1996) for a survey of empirical issues related to foreign exchange risk premiums). In contrast, the unobserved components approach will only require the assumption of rational expectations and the observation of the time-series property of the sum of unobservable components, the risk premium and a prediction error in

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period of appreciation (from late 1994 to mid-1995).

this case (see Section 5.2 for details). While this approach may or may not be the most reliable way of estimating the risk premium, it may well be the only feasible method for our purpose, given the limited availability of high-frequency data on macroeconomic variables and the need to restrict ourselves to a relatively short period of time immediately preceding the East Asian currency crisis, for which testing for policy credibility would be a relevant exercise.

The chapter is organized as follows. Section 5.2 explains the basic methodology of the unobserved components model in extracting the unobservable foreign exchange risk premium from the observable ex post excess return. Sections 5.3 and 5.4 report the results of applying the unobserved components model to East Asian data for 1994-97 and discuss their implications. In particular, Section 5.3 presents the estimates of risk premiums, while Section 5.4 discusses the credibility of the so-called dollar peg policies of the East Asian countries by estimating the expected future spot rates. Section 5.5 provides a summary and concluding remarks. Finally, Appendix 5.1 summarizes the source and nature of the data, and Appendix 5.2 provides an outline of the Kalman filter algorithm for the unobserved components model.

5.2. An Unobserved Components Model of the Foreign Exchange Risk Premium

5.2.1. Decomposing the Excess Return

For a given East Asian currency (*i.e.,* the Indonesian rupiah, the Korean won, the Malaysian ringgit, or the Thai baht), we begin by decomposing the forward exchange rate into two components, the expected future spot rate and the risk premium, as follows:

$$f(t,t+1) = E_t(s(t+1)) + \rho(t), \tag{1}$$

where $f(t,t+1)$ is the natural logarithm of the one period ahead forward rate at time t; $E_t(s(t+1))$ is the rational or efficient forecast of the natural logarithm of the spot rate at time $t+1$, based on all information available at time t; and $\rho(t)$ is the risk premium. Throughout the chapter, the exchange rate is defined as the price of the foreign currency (*i.e.,* the US dollar) in terms of the domestic (*i.e.,* East Asian) currency in question, such that an increase in f or s denotes a depreciation against the US dollar.

Subtracting $s(t+1)$, the realized future spot rate, from both sides of equation (1) and defining the forecast error at time t+1 as $v(t+1) = E_t s(t+1) - s(t+1)$, we obtain:

$$f(t,t+1) - s(t+1) = \rho(t) + v(t+1), \tag{2}$$

where $v(t)$, the forecast error, is serially uncorrelated with zero mean under the assumption of rational expectations. Assuming covered interest parity, equation (2) can equivalently be written as:

$$ER(t) = i(t) - i^*(t) - (s(t+1) - s(t)) = \rho(t) + v(t+1), \tag{3}$$

where i is the domestic (or East Asian) currency interest rate and i^* is the foreign currency (or US dollar) interest rate. Equation (3) shows that the excess return ($ER(t)$) consists of a risk premium and a white noise error, or a signal and noise. We prefer the form of equation (3) to the form of equation (2) because of the simple fact that data on forward exchange rates are not as readily available as data on money market interest rates for the East Asian countries in our sample. Our task is to extract the (unobservable) risk premium from the (observable) excess return.[2]

5.2.2. *Identifying the Foreign Exchange Risk Premium*

Now, let us first specify the motion of the risk premium so as to isolate it from its noisy environment. To do so, we investigate the time-series property of the excess return ($ER(t)$) by using the Augmented Dickey-Fuller (ADF) test. We can then use the observed motion of the excess return to derive information about the time-series property of the underlying risk premium by resorting to a summation theorem for moving-average processes (Wolff (1987)), inasmuch as the excess return is made up of the risk premium and a white noise error. Based on the ADF test statistics for excess returns (Table 5.2), we find that the excess returns were integrated of order one, i.e., I(1), suggesting that the risk premium ($\rho(t)$) was also non-stationary (I(1)), given that $v(t)$ is a white noise error. Hence, we may characterize the risk premium as following a random walk process.[3]

The resulting unobserved components model for the risk premium can be specified as the following system of equations:

[2] In contrast, Wolff (1987, 2000b) and Cheung (1993) have extracted the risk premium from the observed difference between the forward rate and the future spot rate, along the lines of equation (2). On the basis of the US dollar exchange rates of the pound sterling, the deutsche mark and the Japanese yen, these studies find that the risk premium was time-varying with a high degree of persistence.

[3] While some studies suggest that the risk premium is stationary (*e.g.,* Baillie and Bollerslev (1989) and Engel (1996)), others show that it evolves as a non-stationary process for major currency exchange rates (*e.g.,* Crowder (1994) and Evans and Lewis (1995)). In either case, the risk premium is shown to have a high degree of persistence (Cheung (1993) and Baillie and Bollerslev (1994)), so that we are justified in modeling the premium as an I(1) process, particularly given the

$$ER(t) = i(t) - i^*(t) - (s(t+1) - s(t)) = \rho(t) + v(t+1), \quad (4)$$

$$\rho(t) = \rho(t-1) + \varepsilon(t), \quad (5)$$

$$v(t) \sim i.i.d.N(0, R), \quad (6)$$

$$\varepsilon(t) \sim i.i.d.N(0, Q), \quad (7)$$

$v(t)$ and $\varepsilon(t)$ are independent for all t. (8)

This model assumes that $\rho(t)$ follows a random walk process, such that changes in the parameters are randomly driven by the disturbance $\varepsilon(t)$. The profile of $\rho(t)$, therefore, evolves over time according to equation (5).

In *state-space form*, equation (4) is known as a *measurement equation*, which relates the observed variable to the unobserved component $\rho(t)$, while equation (5) is a *transition equation*, which describes the evolution of the unobserved component over time. This model can be estimated by the maximum likelihood method through the application of the Kalman filter (*e.g.*, Harvey (1981, 1989) and Hamilton (1994)), which recursively updates the estimate of $\rho(t)$ by utilizing the new information in $ER(t)$. In this sense, the Kalman filter technique can be viewed as a Bayesian method.[4] An outline of the Kalman filter algorithm for this unobserved components model is presented in Appendix 5.1.

5.3. Estimating the Foreign Exchange Risk Premium

5.3.1. The Overall Performance of the Model

short time horizon of less than 4 years.

[4] The Kalman filter may be viewed as mimicking a sequential optimal learning process. The predictions are rational in the sense that the agent is assumed to optimally utilize current and past information when learning about his or her stochastic environment (see Harvey (1981) and

The results of applying the unobserved components model, given by equations (4)-(8), to monthly East Asian data for the period January 1994-June 1997 are summarized in Table 5.3. To begin with, the overall reasonableness of the fitted model may be examined by checking to see how much of the risk premium is captured. If the model has succeeded in adequately capturing the risk premium, the error terms should be serially uncorrelated. This can be checked by the Box-Pierce portmanteau Q-statistics (calculated from the residuals), which indicate that they are indeed serially uncorrelated in all four cases. We thus conclude that the unobserved components model was successful in capturing the essence of the foreign exchange risk premium in all four East Asian currencies.

5.3.2. Mean Foreign Exchange Risk Premiums

Next, we note that, for all currencies, the estimated mean premium was positive and significant at the one percent level, ranging from 0.0028 (for the Indonesian rupiah) to 0.0050 (for the Thai baht). These are large in magnitude because the estimated risk premium of 0.0050 (for the Thai baht), for example, means roughly 0.5 percent per month or 6 percent per year.[5] Likewise, the risk premium was roughly 3.4 percent for the Indonesian rupiah, 5 percent for the Korean won, and 3.8 percent for the Malaysian ringgit. Evidence thus seems to indicate that the interest differentials favoring East Asian currency assets during 1994-97 included substantial risk premium components and that substitutability between East Asian currency assets and US dollar assets were highly imperfect.

This evidence of imperfect asset substitutability is consistent with the evidence of

Hamilton (1994).

[5] It should be noted that the so-called peso problem may have biased upward the estimate of the risk premium. In other words, the estimated risk premium may include a more permanent component that reflects a large yet unrealized change in the exchange rate. See, for example,

weak causal or structural relationship between an increase in foreign exchange reserves and the growth of monetary aggregates, as documented by Chapter 4 for these and other East Asian countries during 1987-97. Chapter 4 interpreted this evidence to conjecture that the series of tight macroeconomic policy measures taken in these countries to counter the expansionary impact of the massive reserve inflows --- collectively called sterilization --- were effective in limiting the growth of narrow and broad. The evidence of imperfect asset substitutability, as indicated by the presence of a risk premium, certainly supports the conjectured efficacy of sterilization in the East Asian context. In turn, the apparent efficacy of sterilization may have worked to promote additional capital inflows by raising the level of domestic interest rates, to the extent that international investors willingly took open speculative positions on the risk premium through what became known as "carry trades".

The estimated variances of the risk premium and the forecast error may also be of some interest (Table 5.3). In the case of the Korean won and the Malaysian ringgit, the variance of the risk premium was larger than the variance of the forecast error, whereas the opposite is true for the Indonesian rupiah and the Thai baht. This means that the variation of the risk premium accounted for more than half the variation of the excess return for the Korean won and the Malaysian ringgit, while the variation of the forecast error accounted for more than half the variation of the excess return for the Indonesian rupiah and the Thai baht. This may be a simple reflection of the fact that the Indonesian rupiah (with a crawling peg to the US dollar) and the Thai baht (with a peg to a currency basket) fluctuated more than the other two currencies against the US dollar (Takagi (1999)).

Evans and Lewis (1995).

5.3.3. The Time Profiles of Risk Premiums

Depicting the time profiles of the estimated risk premiums for April 1994-May 1997, we find that they were time-varying with some degree of persistence (Figure 5.2). The persistent yet time-varying nature of the estimated risk premium is consistent with a wide range of studies finding similar evidence for major industrial country currencies (Wolff (1987, 2000a), Cheung (1993) and Nijman, Palm and Wolff (1993)). Several observations are in order. First, the estimated risk premium did vary over time throughout the sample period and took on both positive and negative values, as previously demonstrated for industrial country data (see Engel (1996)). The risk premium varied between -0.4 and 1.2 percent per month for the Indonesian rupiah, between -1.5 and 2.3 percent for the Korean won, between -1.2 and 2.6 percent for the Malaysian ringgit, and between -0.3 and 2.8 percent for the Thai baht.

Second, unlike the industrial country results (in which the risk premium swings between positive and negative values), the risk premium for the East Asian currencies was for the most part positive, indicating that East Asian currency assets were generally considered to be riskier than US dollar assets. This, however, should not minimize the fact that the East Asian assets were sometimes considered to be safer as in the case of the Indonesian rupiah in early 1997, the Korean won in the second quarter and the latter part of 1996, and the Malaysian ringgit during the first half of 1995. At least on the surface, it appears that the negative risk premiums were associated with the presence of some appreciation pressure against the background of general depreciation (*e.g.,* the rupiah in early 1997, the won in 1996, and the ringgit from mid-1995 to late 1995).[6]

[6] During these periods, the currencies in question were generally depreciating. Even so, the Indonesian authorities were easing monetary policy to contain upward pressure on the rupiah, the Korean authorities were containing the appreciating pressure on the won arising from the sharp appreciation of the yen against the US dollar, and the Malaysian ringgit was under appreciating

Third, the estimated risk premium was considerably more volatile than the associated interest rate differential (*c.f.,* Figure 5.1). This means that the covariance of the expected rate of depreciation and the risk premium was negative, the result consistent with the evidence well known in industrial country data (Engel (1996)). Fourth, the premium of the Thai baht rose sharply (from 0 to 2.8 percent per month) during the two months preceding the onset of the currency crisis in July 1997, along with the sharply rising interest rate differential. This behavior of the risk premium indicates that the perceived risk of Thai baht assets rose sharply during this period in the minds of market participants. In contrast, the risk premium did not rise during this period in the case of the Indonesian rupiah and the Malaysian ringgit.

5.4. On the Credibility of the So-Called Dollar Peg Policies

5.4.1. Monetary, Exchange Rate and Other Policy Developments

Given the estimated risk premium, we can assess the credibility of the so-called dollar peg policies of the East Asian countries by obtaining the implied expected future spot rate. This can be done by making use of the fact that the interest rate differential is equal to the sum of the risk premium and the expected rate of depreciation. Before proceeding further, however, it may be helpful first to review the salient features of the monetary, exchange rate and other policy developments, which might have affected the market participants' exchange rate expectations, against the overall environment of the continued tight monetary policies designed to contain the expansionary impact of the reserve inflows.[7]

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pressure, following the upgrading of the credit rating of Malaysia's long-term and short-term external debt in late 1994 (see EIU, various issues).

[7] Unless otherwise indicated, the following information comes from EIU, various issues.

(i) Indonesia

Throughout the period, the Indonesian authorities maintained a tight monetary stance in order to contain inflationary pressure (with year-on-year consumer price inflation of 7.5-10.5 percent from early 1994 through mid-1996) against the recurrent fear of overheating. The measures included increases in the discount rate (*e.g.*, January 1995), increases in reserve requirements (*e.g.*, February 1996; April 1997), tightening of direct credit controls (*e.g.*, 1995/1996). However, when the inflationary pressure appeared contained towards the end of the period (when the year-on-year rate of inflation declined to around 5 percent), they somewhat eased monetary policy as in late 1996 and again in early 1997 in order to courage growth and to restrain the upward pressure on the rupiah, respectively. Given the higher rate of inflation relative to its most trading partners, the exchange rate was consistently adjusted downward to maintain the real effective exchange rate. The exchange rate depreciated from 2143 rupiah per US dollar in early 1994 to 2450 in mid-1997 (Figure 5.2a). There were temporary depreciating pressure on the rupiah from time to time, as in January 1995 (in connection with the Mexican crisis); there was temporary upward pressure, as in mid-1996 and in early 1997. In June and September 1996, the authorities widened the rupiah's fluctuating band against the US dollar, in order to increase the foreign exchange risk faced by foreign currency traders.

(ii) Korea

Throughout the period, the Korean authorities were faced with the dilemma of containing inflation and maintaining exchange rate stability against the US dollar, against the background of large capital inflows. Inflation was in the range of 3.6-6.9

percent, moderate by emerging market standards but high relative to the industrial country norm. The authorities generally maintained a tight monetary stance by raising interest rates; this stance was supplemented by some fiscal tightening. However, from late 1996, the monetary stance turned more accommodating, with interest rates beginning to decline. Although the won remained relatively stable against the US dollar (fluctuating between 758 and 820) through mid-1996, it began to depreciate sharply in late 1996, reaching almost 900 in the first quarter of 1997 (Figure 5.3b). This movement of the won was in part influenced by the downward movement of the yen against the dollar, which became evident from the summer of 1995; the authorities often intervened to maintain downward pressure on the won to make Korean exports remain competitive with Japanese exports.

(iii) Malaysia

Throughout this period, the Malaysian authorities pursued a tight monetary policy amid the continued concern about excess liquidity. The measures included increases in reserve requirements (*e.g.*, in January 1994, February and June 1996) and direct controls (*e.g.*, ceiling on external bank liabilities in January 1994; limit on bank lending in April 1997). Inflation was modest and falling, declining from the range of 5-6 percent in 1994 to the range of 2-3 percent in 1996 and early 1997. The exchange rate was generally stable, narrowly fluctuating around 2.5 ringgit per US dollar, although there sometimes was moderate pressure for appreciation, as from mid-1994 to mid-1995 (when the ringgit appreciated from 2.6 to 2.44), and from late 1996 to early 1997 (Figure 5.3c). From mid-1995 to early 1996, on the other hand, there was speculation against the ringgit, which was reversed with a recovery of capital inflows in late 1996.

(iv) Thailand

Particularly from mid-1994, the Thai authorities pursued a tight monetary policy amid the mounting concern about excess liquidity and inflation. Interest rates rose, reaching the highest level in 4 years in December 1995. Consumer price inflation remained in the rage of 4-6 percent, although there was a temporary pickup in late 1995 and early 1996 to a range exceeding 7 percent. During the 12-month period preceding the Thai crisis of July 1997, however, the stance of monetary policy was kept easy in part to support the deteriorating balance sheets of finance companies (Fane and McLeod (1999)). The exchange rate was relatively stable against the US dollar, fluctuating between 24.66 and 25.97 baht per dollar in quarterly average terms (Figure 5.3d). The baht, however, was subjected to periodic depreciating pressure, as during the Mexican crisis of January 1995, in the aftermath of the Bangkok Bank of Commerce (BBC) scandal and the subsequent downgrading of the credit rating of Thailand's external debt in the summer of 1996, in late 1996 and early 1997 (when there were rumors of an impending devaluation associated with weakening economic fundamentals), and during the financial crisis of March 1997.

5.4.2. Estimating the Expected Future Spot Exchange Rate

Figure 5.3 depicts, for each East Asian currency, the current US dollar exchange rate and the implied expected one-month ahead US dollar exchange rate during the period of April 1994-May 1997. Several interesting observations emerge from this set of figures. First, except for the Indonesian rupiah, it appears that whenever the domestic currency depreciated against the US dollar, the market participants apparently expected the currency to depreciate further over the coming month.

Likewise, whenever the currency appreciated against the US dollar, there were apparent expectations that the currency would further appreciate. This adaptive nature of exchange rate expectations is seen in the observation that the dotted line is above (below) the solid line when the latter moves upward (downward).

Second, in contrast, the dotted line for the Indonesian rupiah almost always exceeded the solid line, indicating that the market participants apparently expected the currency to display trend depreciation. This is consistent with the crawling peg arrangement, under which the rupiah was regularly adjusted downward against the US dollar. Third, for the Thai baht, the market participants were almost consistently expecting the currency to depreciate over the month, from mid-1995 to early 1997, when the expectations of depreciation rose sharply. It appears that the market participants were questioning the sustainability and hence the credibility of the so-called dollar peg policy of the Thai authorities through much of the sample period.

Finally, in general, the periods of expected (as well as actual) depreciation were associated with monetary easing, as in the case of the Indonesian rupiah and the Korean won in late 1996 and early 1997. In the case of the Thai baht, it appears that the expectations of depreciation were fueled by the reported rumors of an impending devaluation from mid-1996 to early 1997, in connection with the BBC scandal and the financial crisis of March 1997. In contrast, the periods of expected (as well as actual) appreciation were associated with the beginning of monetary tightening, as in the case of the Thai baht in early 1995. All in all, these observations suggest that the market participants held systematic expectations of either appreciation or depreciation for all four East Asian currencies throughout much of the sample period, so that the exchange rate policies of these countries can hardly be characterized as dollar peg policies, at least in terms of the perception of the market participants.

5.4. Summary and Concluding Remarks

In this chapter, we have used an unobserved components model to extract the foreign exchange risk premium from the ex post excess returns of East Asian currency assets over US dollar assets and derived the implied expected future spot rates of East Asian currencies against the US dollar. Empirical results, obtained from monthly data on the crisis East Asian countries of Indonesia, Malaysia, Korea and Thailand for the period January 1994-June 1997, have generally indicated that the risk premium was substantial and time-varying and that the market participants questioned the credibility of the so-called dollar peg policies by consistently forming expectations of either appreciation or depreciation over the coming month.

By way of conclusion, three implications can be stated. First, the presence of a risk premium means that East Asian currency assets and US dollar assets were imperfect substitutes, so that the sterilization and other tight monetary measures adopted in these countries were effective in limiting the growth of monetary aggregates arising from the large reserve inflows. Second, the consistently observed expectations of short-term depreciation or appreciation suggest that market participants were questioning the credibility of the exchange rate policies. In this sense, the so-called dollar peg policies were by no means promoting a stable external environment for these East Asian countries. Third, if capital inflows were magnified at all, it was not necessarily by the policy of exchange rate pegging, as frequently argued, but rather by the presence of substantial risk premiums. During much of the period under consideration, the existing interest rate differential over US dollar assets was to a large extent offset by expected depreciation, so that what counted for profit-motivated investors was the magnitude of the risk premiums.

Appendix 5.1. The Sources and Nature of Interest Rate Data

Data on dollar exchange rates and short-term market interest rates were obtained from the *Asian Wall Street Journal*, daily issues. They are end-of-month data for the period from January 1994 to June 1997 (or from January 1994 to February 1997 for Korea only), covering the Indonesian rupiah, the Korean won, the Malaysian ringgit and the Thai baht. The East Asian currency interest rates are the 30-day interbank settlements or nearest comparable rates, while the US dollar interest rate is the 1-month Eurodollar rate. Although both 30-day (or 1-month) and 90-day (or 3-month) rates were available (see Table 5.1 and Figure 5.1), we have only used the 30-day (or 1-month) rates in our empirical work so as to avoid autocorrelated residuals resulting from overlapping observations (with the sampling interval shorter than the forecasting horizon) when monthly data are used.

Appendix 5.2. An Outline of the Kalman Filter Algorithm

Following Harvey (1981, 1989), the Kalman filter algorithm for the unobserved components model can be presented as follows.

Let $a(t)$ denote the minimum mean square liner estimator (MMSLE) of $\rho(t)$ based on all available information, including the current observation of $ER(t)$. Let $a(t|t-1)$ denote the MMSLE of $\rho(t)$ at time t-1. It is assumed that all available information is incorporated in $a(t-1)$ at time t-1, with a covariance matrix $P(t-1)$ where $P(t-1)$ is known.

Then, the following set of equations (A1)-(A5) describes the Kalman filter algorithm for the unobserved components model:

$$a(t|t-1) = a(t-1), \quad \text{(A1)}$$

$$P(t|t-1) = P(t-1) + Q, \quad \text{(A2)}$$

$$a(t) = a(t|t-1) + P(t|t-1) \cdot (ER(t) - a(t|t-1)) / f(t), \quad \text{(A3)}$$

$$P(t) = P(t|t-1) - P^2(t|t-1) / f(t), \quad \text{(A4)}$$

$$f(t) = P(t|t-1) + R, \quad \text{(A5)}$$

where equations (A1) and (A2) are prediction equations, and equations (A3), (A4) and (A5) are updating equations. The former predicts the mean and variance of the coefficients, given the information set at time t-1. The latter updates the equation on the basis of current information, given the current value of $ER(t)$ as well as its history in Bayesian fashion (Harvey (1989) and Hamilton (1994)).

Here, $v(t) = ER(t) - a(t|t-1)$ is defined as the one period ahead prediction error. Then, the log likelihood function of the model can be written as,

$$\ln L = -\frac{T}{2}\ln(2\pi) - \frac{1}{2}\sum_{t=1}^{T}\ln f(t) - \frac{1}{2}\sum_{t=1}^{T}\frac{v^2(t)}{f(t)}. \quad \text{(A6)}$$

This log likelihood function is estimated by using the maximum likelihood method. As the starting values of ρ required in initializing the Kalman filter procedure, we use the OLS estimates obtained from the first five observations.

Table 5.1.
Summary Statistics: Interest Rate Differentials and Excess Returns on East Asian Currency Assets, 1994-97

(a) Interest rate differentials over US dollar assets (in percent per year)

	Indonesia		Korea		Malaysia		Thailand	
	1 month	3 month	1 month	3 month	1 month	3 month	1 month	3 month
Sample	94:1-97:6	94:1-97:4	94:1-97:2	94:1-97:4	94:1-97:6	94:1-97:4	94:1-97:6	94:1-97:4
Mean	8.288	8.706	7.882	8.256	0.939	0.923	6.133	6.013
Std Dev	1.855	1.902	1.273	0.778	1.101	1.133	2.000	1.755
Min	4.438	4.438	6.625	6.356	-0.900	-0.958	2.750	2.875
Max	10.28	11.13	12.25	10.36	3.436	3.323	14.75	13.91
t-statistic	28.95	29.67	27.66	68.80	5.526	5.279	19.88	22.21
p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000

(b) Excess returns over US dollar assets (in percent per year) [a]

Sample	94:1-97:6	94:1-97:4	94:1-97:2	94:1-97:4	94:1-97:6	94:1-97:4	94:1-97:6	94:1-97:4
Mean	3.621	4.002	5.053	4.642	3.524	2.476	6.102	5.368
Std Dev	6.139	3.562	12.45	9.216	12.65	7.955	9.559	5.200
Min	-9.074	-1.503	-28.19	-15.64	-17.41	-13.92	-9.516	-1.662
Max	18.59	12.90	36.48	25.65	46.60	22.29	54.78	22.35
t-statistic	3.777	7.016	2.469	3.146	1.784	1.944	4.087	6.446
p-value	0.001	0.000	0.018	0.003	0.081	0.059	0.000	0.000

Note: [a] Adjusted for actual exchange rate changes.
Sources: The *Asian Wall Street Journal*; authors' estimates.

Table 5.2.
The Time-Series Property of Excess Returns [$ER(t)$], 1994-97

	Indonesia	Korea	Malaysia	Thailand
	-2.564 (5) [0.296]	-2.663 (7) [0.252]	-3.066 (6) [0.115]	-2.285 (2) [0.442]

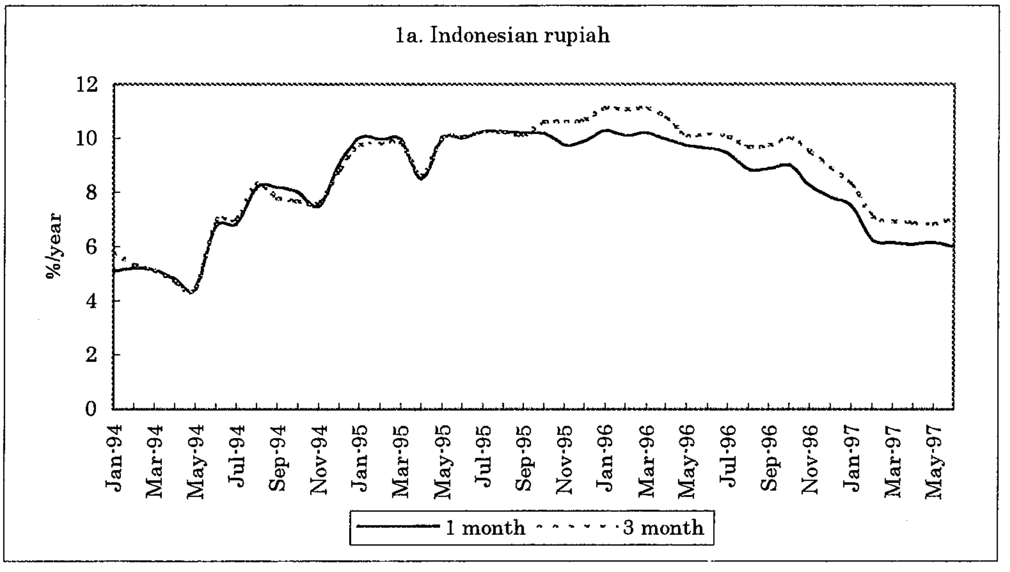
Note: The figures are Augmented Dickey-Fuller statistics obtained with a constant term and time trend. The figures in parentheses indicate lag length; and those in brackets are p-values.

Table 5.3.
Summary Statistics from the Estimated Unobserved Components Model

	Indonesia	Korea	Malaysia	Thailand
Mean (ρ)	0.0028***	0.0042***	0.0032***	0.0050***
Var (ρ)	1.26E-05	6.41E-05	5.99E-05	3.02E-05
t-statistic	5.079	3.219	2.675	5.799
Var (ν)	1.39E-05	4.23E-05	5.34E-05	3.14E-05
Q (7)	10.1	10.1	11.2	6.15
log L	147.6	112.8	120.6	131.3

Note: The unobserved components models were estimated by the maximum likelihood method through the Kalman filter technique. Mean (ρ) and Var (ρ) are the mean and the variance of the estimated risk premium. Var (ν) is the variance of the derived noise. Q (7) is the Box-Pierce portmanteau test statistic calculated from residuals. The critical value of the Q (7) statistic is 12.02 at the 10 percent level. *** indicates the statistic is significant at the 1 percent level.

Figure 5.1.
1-month and 3-month Interest Rate Differentials:
East Asian Currency Assets over US Dollar Assets, 1994-97
(in percent per year)



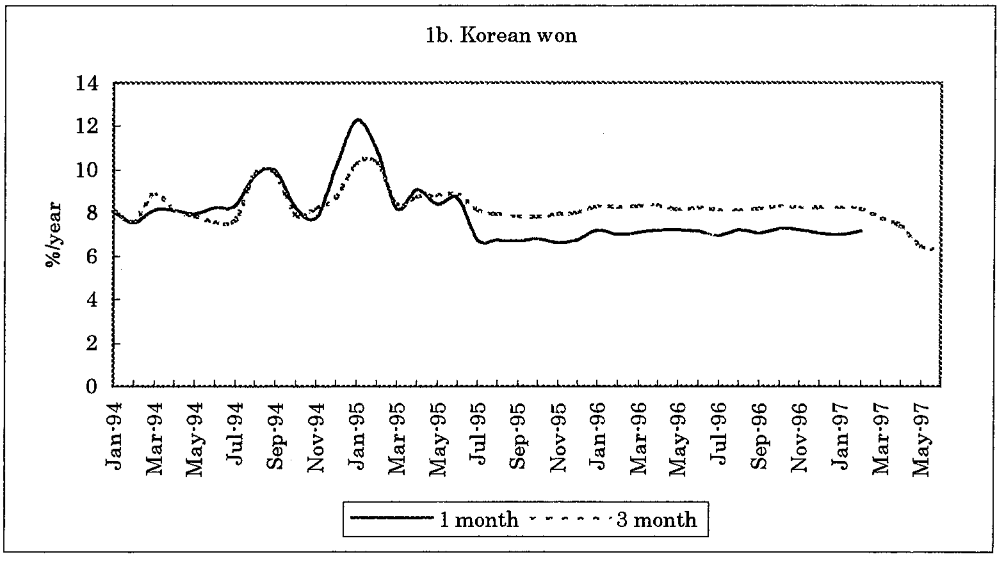
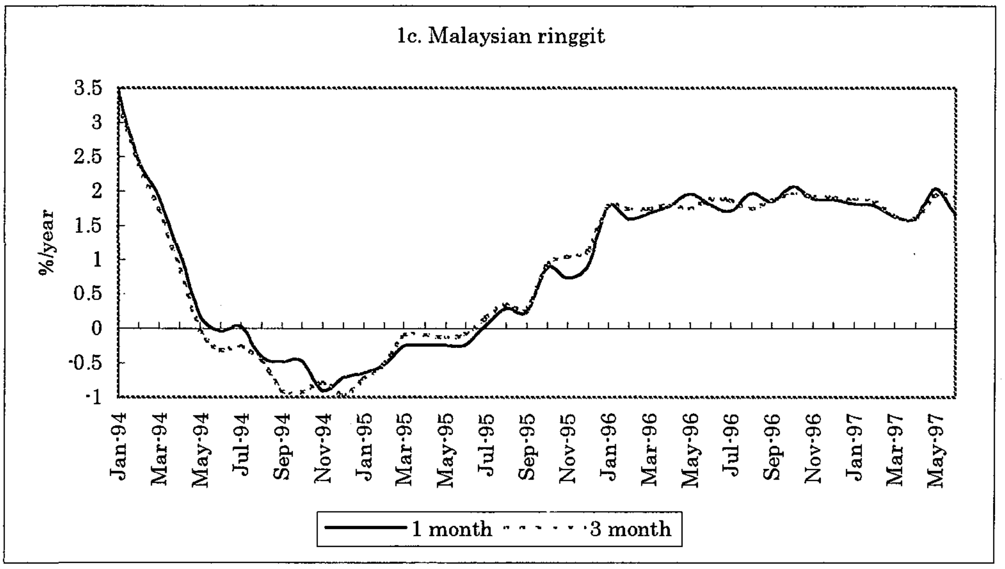
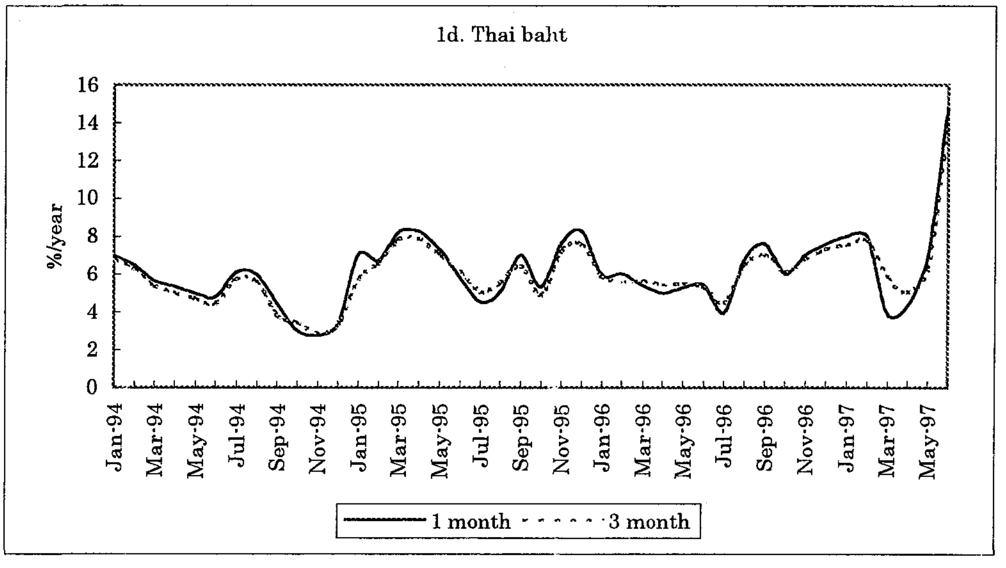


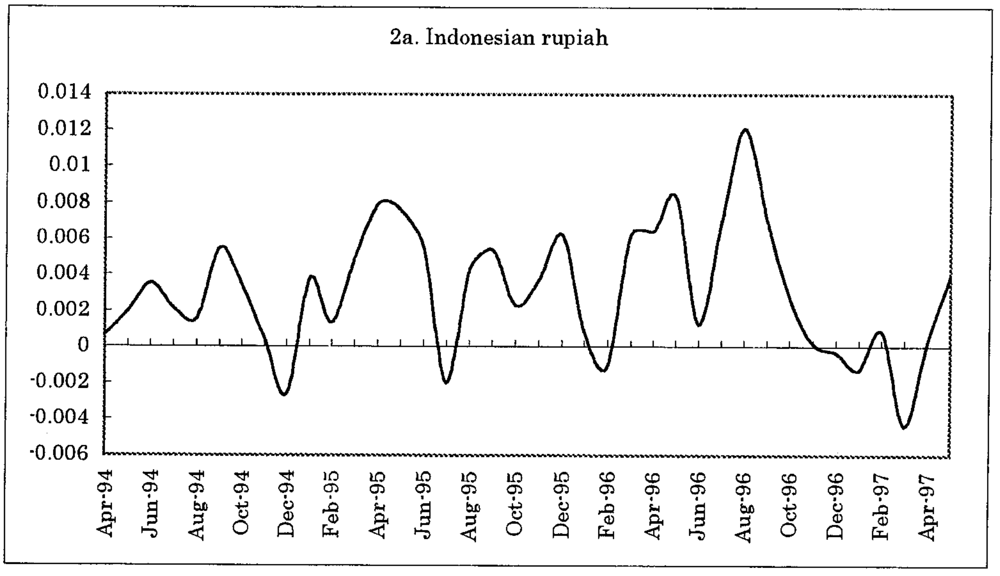
Figure 5.1 (Continued).





Source: The *Asian Wall Street Journal.*

Figure 5.2.
1-month Foreign Exchange Risk Premiums relative to the US dollar, 1994-97



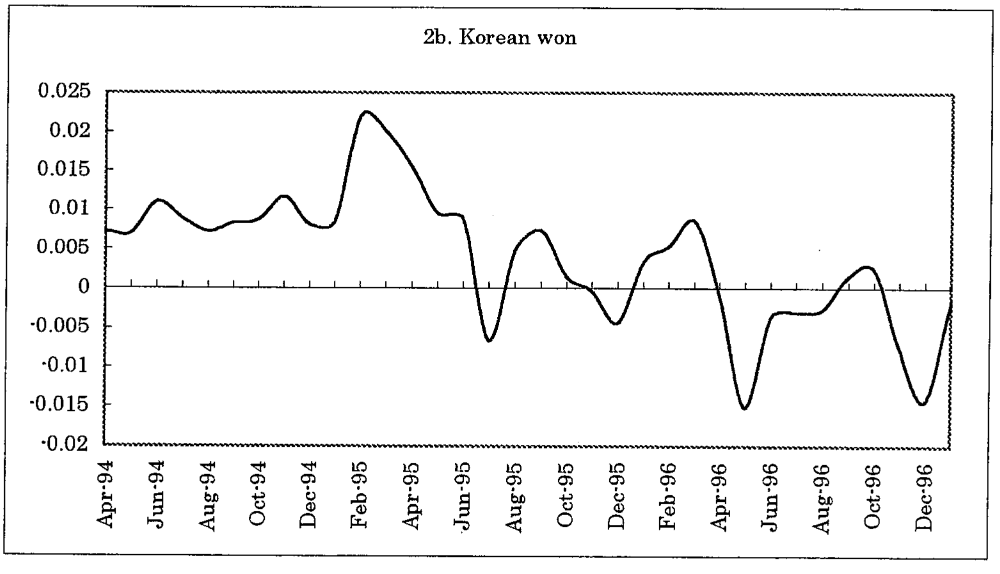
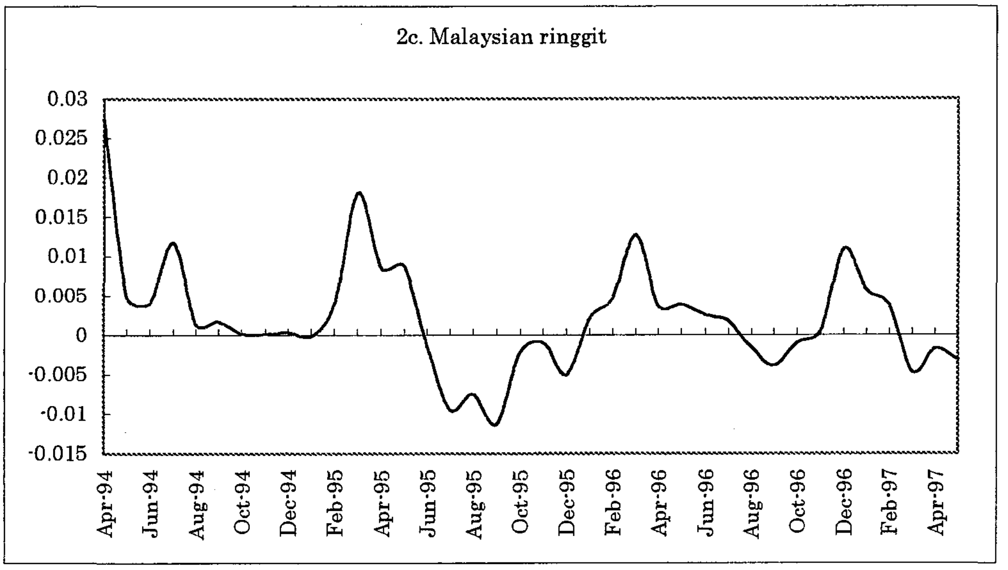
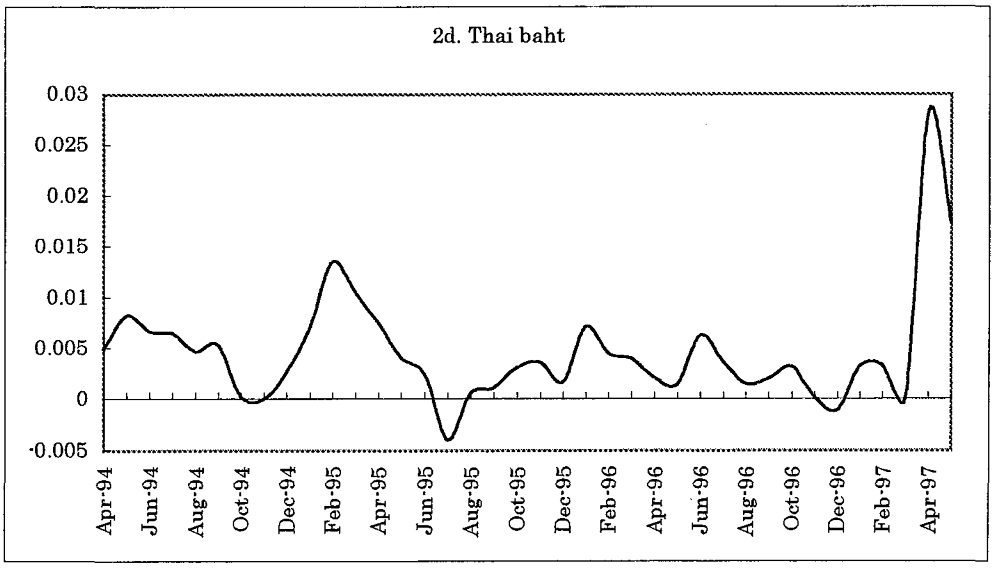


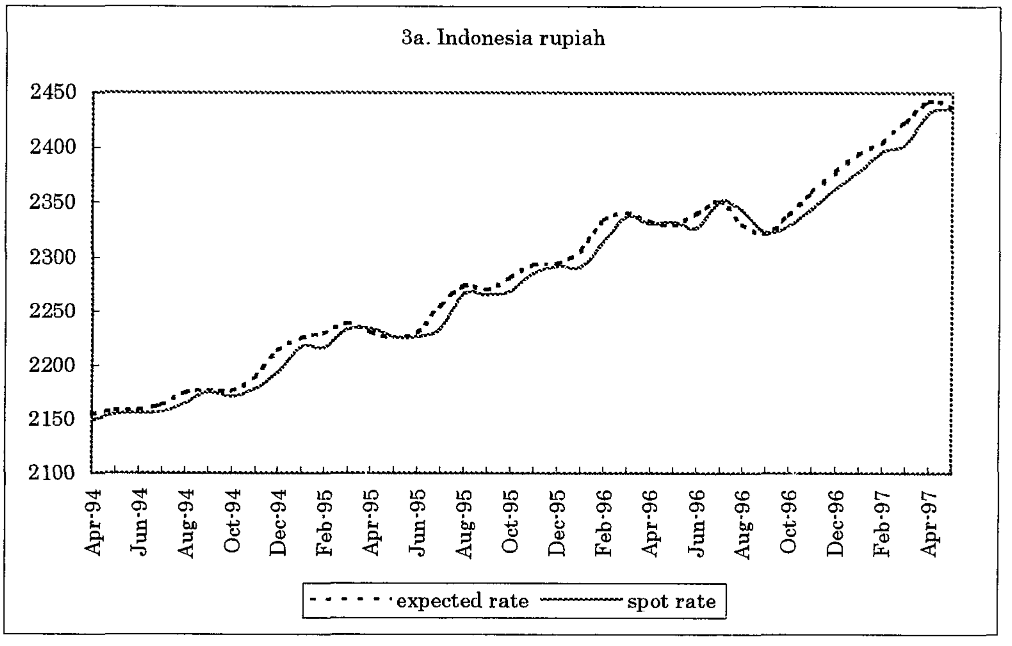
Figure 5.2 (Continued).





Sources: The *Asian Wall Street Journal*; authors' estimates based on the unobserved components model.

Figure 5.3.
Current and Expected 1-Month Future Exchange Rates against the US Dollar, 1994-97



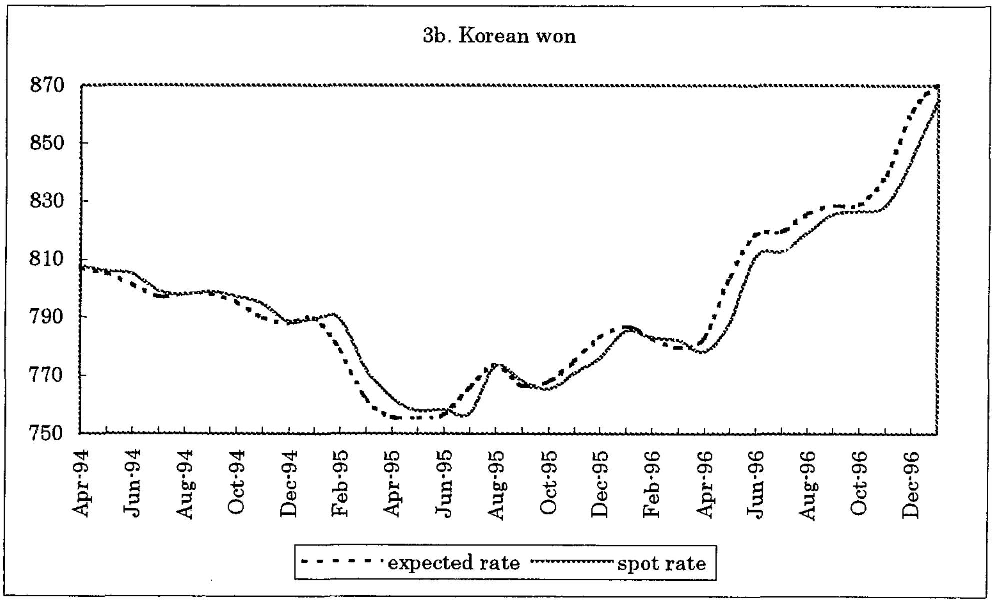
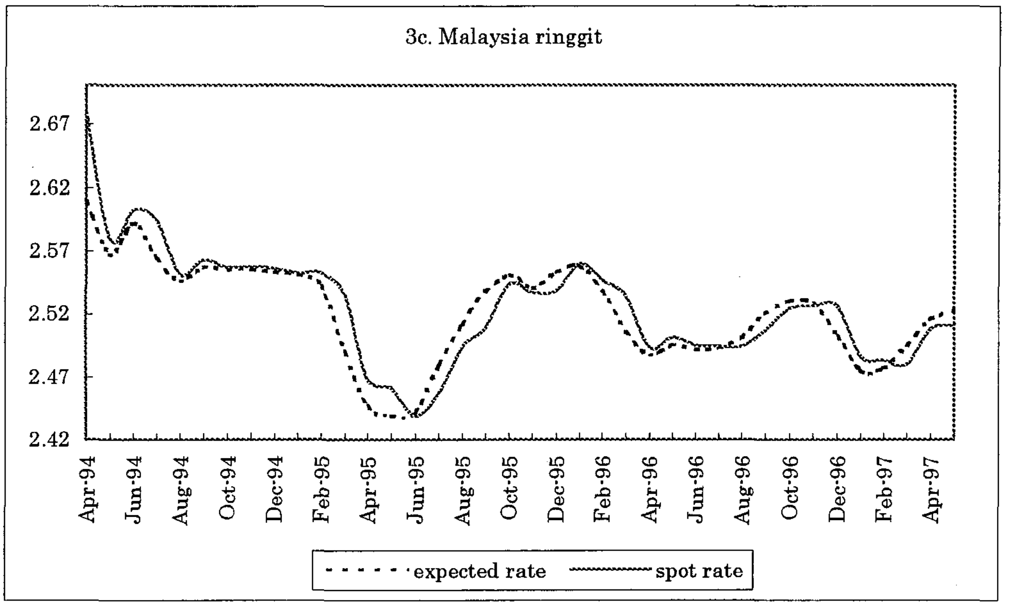
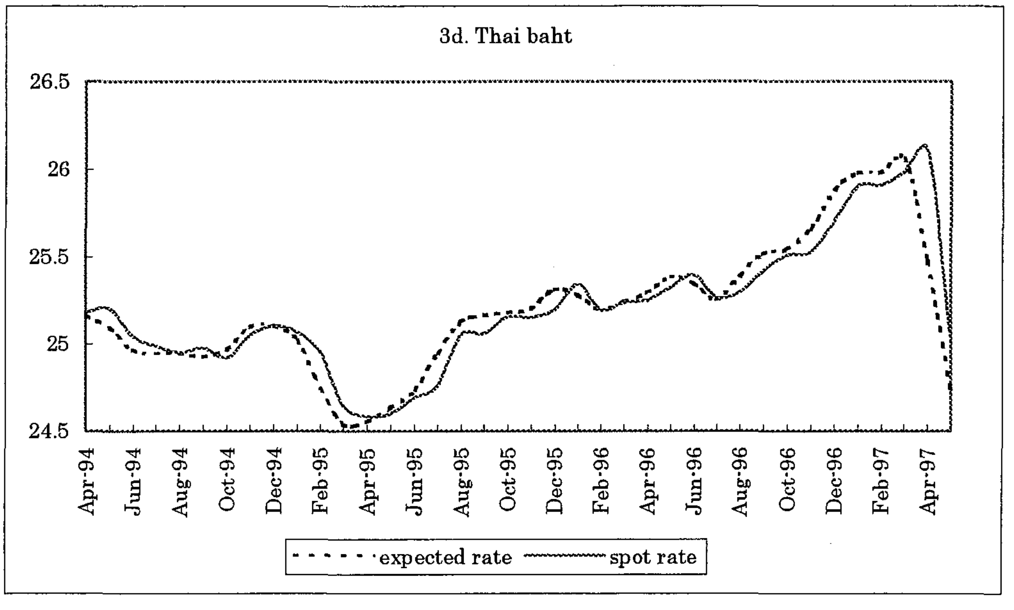


Figure 5.3 (Continued).





Sources: The *Asian Wall Street Journal*; authors' estimates based on the unobserved components model.

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