INTERPRETING THE ASIAN CURRENCY CRISIS:

EMPIRICAL ANALYSIS AND PREDICTION.

By

Hoon Kim

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ABSTRACT

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This dissertation investigates the causes of the Asian crisis and improves implementation in predicting actual currency crises. Chapter II presents an overview of the inception and development of the Asian crisis with a focus on the movements of the macroeconomic variables and the structural conditions of the financial systems. Chapter II shows some evidence of deterioration of fundamentals. Yet the deterioration was not so severe as to make the outbreak of the Asian currency crisis an inescapable result.

Chapter III's survey of the currency crisis literature finds that most nonstructural empirical studies are limited by a lack of robustness to various sensitivity tests and poor performance in the prediction of actual crises. Therefore, to determine the uniqueness of the Asian crisis and to improve performance in predicting actual crises, structural model studies are used to model the currency crisis. Chapter IV then offers an analysis of the time series properties and forecasts of each variable of the structural currency crisis models introduced in Chapter III for the derivation of shadow exchange rates and probabilities of collapse. As a result of the addition of the ARFIMA(p,d,q)-FIGARCH(P,δ,Q) model to the analysis, it is found that some processes exhibit long memory in both their conditional mean and variances. In Chapter V, long and short-run

real money demand functions are estimated for the derivation of shadow exchange rates and probabilities of collapse. The empirical results of this chapter suggest that both long and short-run models can be specified in South Korea and in Malaysia. This justifies the monetary approach using the structural currency crisis model. Chapter VI estimates shadow exchange rates and probabilities of an exchange rate regime change for South Korea and Malaysia. Two countries experienced severe currency devaluation. This employs forecasts for the analyzed variables in Chapter IV and the estimates of real money demand function found in Chapter V. Both shadow exchange rates and probabilities of collapse reflecting the presence of weak fundamentals show that there were reasons to anticipate the 1997 Asian currency crisis.

In Chapter VII, a more extensive analysis of currency crisis with respect to the number of countries and the currency crisis episodes is performed using panel data. Here, the focus is on the role of contagion effects on the spread of currency crisis. The empirical results show that lending booms impact the currency crisis index much more among developing countries than industrial countries. In addition, contagion effects, represented by trade linkage and market sentiment, significantly improve the ability to predict the eruption of a currency crisis after controlling for other macroeconomic variables.

Based on the preceding empirical results, it appears that weak fundamentals and contagion effects can be indicators of upcoming currency crisis implying cumulative depreciation pressure. Nevertheless, a currency crisis cannot erupt without triggering events such as bank failure, corporate failure or political uncertainty that induce an equilibrium, currency crisis, to be an inescapable result among the multiple equilibria.

To my parents,

Yeon Tae Kim

And

Sung Ja Jung,

And to my wife,

Won Young

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TABLE OF CONTENTS

LIST OF TABLES

LIST OF FIGURES

CHAPTER I

INTRODUCTION

Episodes of speculative attacks on currencies in the 1990s (such as the 1992-93 crises in the European Monetary System and the 1994 Mexico peso collapse) have generated considerable debate on whether currency and financial instability should be attributed to arbitrary shifts in market expectations and confidence instead of weak economic fundamentals. These viewpoints about the underlying causes of a currency crisis are summarized by two main views. According to one view, advocated by 'fundamentalists', crises reflect a sustained deterioration in macroeconomic fundamentals and defective economic policies. Although market overreaction can exacerbate currency crises, fundamentalists stress that the cause of crises are due to structural factors. Another view, favored by 'non-fundamentalists', is that sudden shifts in market expectations and confidence are the crucial sources of initial financial turmoil, its propagation over time, and regional contagion. While the macroeconomic performance of some countries with currency crises was somewhat weak, the extent and depth of the crises should not be attributed to the sharp deterioration in fundamentals but rather to the panic of domestic and international investors. Yet, advocates of both the 'fundamentalist' and the 'non-fundamentalist' view, agree in principle that a deteriorating macroeconomic outlook is a necessary condition for an economy to be vulnerable to a crisis. In fact, it is well understood that multiple instantaneous equilibria, which provide the theoretical preconditions for self-fulfilling crises to occur as rational events, are only possible in a region in which the current or anticipated economic performance is sufficiently weak.

Identifying the source of currency crises has important implications for economic policy: if currency crises are indeed caused by fundamentals, the most effective way to prevent them is to support fiscal and monetary policies that stabilize exchange rates. If, on the other hand, self-fulfilling speculation can trigger crises regardless of fundamentals, there might be a case for specific measures to deter such speculation. One of the measures is a capital control that might help governments defend their currencies.

Whereas the speculative attacks on currencies in the early 1990s have been sufficiently analyzed to reach an agreement about the main source, the cause of the Asian currency crisis in 1997-98 is still under debate.

The main objective of this dissertation is to further study the issue of the causes of the Asian crisis and improve implementation in predicting actual currency crises. The analysis presented here examines the crisis using higher frequency data and more refined models than previous studies. The subsequent chapters are organized as follows:

Chapter II presents an overview of the Asian crisis with an emphasis on the movement of macroeconomic variables and the structural conditions of financial systems. The analysis concentrates on South Korea, Indonesia, Malaysia, the Philippines and Thailand. These countries experienced more severe currency depreciation than other Asian countries. Causal analysis of the currency crisis does not allow one to draw conclusions on their causes. That is, the fundamentalist view is not a more appealing explanation of the crises than the non-fundamentalist view, and vice versa.

In Chapter III, a survey of literature on currency crises is offered in addition to an extended model. The theoretical literature consists of two generations of models. First generation models, as in Krugman (1979), show how speculative attacks occur when the fundamentals are weak. Second generation models study the following two questions: "What happens when government policy reacts to changes in private behavior?" or "What happens when the government faces an explicit trade-off between a fixed exchange rate policy and other objectives such as economic growth, low unemployment or low inflation?"

The nonstructural studies such as the classic study by Frankel and Rose (1996) have attempted to exploit the high variability associated with multi-country information. Estimation results from their probit regression are largely consistent with the theoretical literature, however, the results are not robust and do not forecast crises well. Structural studies, beginning with Blanco and Garber's (1986), have presented strong evidence suggesting that domestic macroeconomic indicators play a key role in determining a currency crisis. Otker and Pazarbasioglu (1996, 1997b) also computed the probability of an exchange rate regime change from the European financial crises in 1992 and 1993 and the Mexican financial crisis in 1994. These studies focus on a particular country in a specific time period illustrating the uniqueness of each country's currency crisis.

An extension of the speculative attack model, suggested by Krugman (1979) and formalized by Flood and Garber (1984a), is derived to capture the uniqueness of the Asian crisis and to improve the performance in predicting actual crises in Chapter III. The model is a stochastic version of the monetary approach to exchange rate determination, in which the government and monetary authority of a small open economy are committed to maintaining the exchange rate by employing some form of a fixed exchange rate system.

Chapter IV provides an analysis of time series property and forecast of all of the variables introduced in the models of Chapter III. The analysis is used to derive shadow exchange rates and probabilities of an exchange rate regime change. Most of the previous studies on the structural analysis of currency crises, Blanco and Garber (1986), Cumby and Van Wijnbergen (1989) and Otker and Pazarbasioglu(1996, 1997b), do not estimate the properties of variables in the model but assume an AR(p) model. Unlike those studies, Goldberg (1994) estimates variables' time series properties and forecasts values one step ahead, using ARIMA models and Akaike tests. Whereas ARIMA models are able to capture autocorrelations that decay at an exponential rate, they cannot be applied to long memory processes where autocorrelations decay slowly. Therefore, the ARFIMA(p,d,q)-FIGARCH(P,δ,Q) model is used to capture the part of economic and financial time series that exhibit long memory in both their conditional mean and variances. Once the specific form of the model is determined and the parameters are estimated, the model is fit over each of the sample time periods to calculate forecasts. The forecasts are used to derive the shadow exchange rate and the probability of collapse.

In Chapter V, long and short-run real money demand functions of the structural currency crisis models introduced in Chapter III are estimated. The structural analyses of currency crises, Blanco and Garber (1986), Cumby and Van Wijnbergen (1989) and Goldberg (1994), estimated a real money demand function without consideration of the non-stationarities of the variables. Therefore, their results suffer from a spurious regression problem and the conventional t-ratio and F significance tests cannot be applied. Unlike previous studies, cointegration and error correction techniques are applied for the modeling of real money demand to remove the spurious regression problem and to use the t-ratio and F significance tests. First, a theoretical framework and an empirical model are presented. Next, unit-root tests are presented to detect the non-stationarity of variables. Lastly, a residual based tests, testing for the number of cointegration relations and estimating the cointegrating vectors, were performed.

Based upon forecasts of each variable and estimates of the real money demand function, shadow exchange rates and probabilities of an exchange rate regime change are derived for South Korea and Malaysia in Chapter VI. The derived shadow exchange rates and probabilities show that fundamentals were weak prior to the Asian crisis.

Chapter VII introduces an empirical model that performs an extensive analysis on currency crisis episodes using panel data. The model investigates a currency crisis focusing on the various variables or other external effects, e.g. contagion effects, whereas the traditional approach emphasizes the role played by declining international reserves in triggering the collapse of a fixed exchange rate. Sachs, Tornell and Velasco (1996) and Tornell (1999) seek to identify macroeconomic variables that can help explain which countries were vulnerable to "contagion effects", but only in emerging markets. Glick and Rose (1998) find that countries with important trade links to the country that initially experienced a crisis are more likely to experience a crisis themselves. Masson (1998) suggests that the contagion effect unexplained by the common external effects and trade links played a major role in the Mexican and Asian crises. I check if the macroeconomic variables that explain the cross-country variation in the severity of crises in emerging markets also have explanatory power in non-emerging markets. In addition, the extent of the contagion effect in all aspects of common external effects, trade linkage and market sentiment is examined. Finally, a currency crisis is predicted using the contagion effect as well as the weak fundamentals to make the predictions more precise than the previous studies'.

Chapter VIII summarizes all the results derived in this dissertation and suggests policies for the prevention of currency crises.

CHAPTER II

THE CAUSE OF THE ASIAN CURRENCY CRISIS

There have been two main alternative views about the causes of the Asian economic, currency and financial crisis that started in 1997.[1] One of the views focuses on the role of weak fundamentals such as growing current account deficits, real currency appreciation, bad loans, overinvestment, and foreign debt accumulation, as being contributors to the crisis.

In contrast, the other view stresses sudden arbitrary shifts in market expectations and confidence, i.e. financial panic, as the key cause of the crisis. Radelet and Sachs (1998b) admit that there were significant underlying fundamental problems in the Asian economies, but assert that these problems were less severe than financial panic in triggering a crisis of such magnitude.

This chapter presents an overview of the Asian currency crisis. For the purpose of finding evidence for either weak fundamentals or financial panics, the movement of macroeconomic variables and the structural conditions of the financial system are discussed. The analysis focuses on South Korea, Indonesia, Malaysia, the Philippines and Thailand. These countries experienced more severe currency depreciation than other Asian countries.

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[1] A list of recent studies are available at http://www.stern.nyu.edu/~nroubini/asia/AsiaHomepage.html

1. The inception and development of the Asian currency crisis

1.1 The period leading to the crisis: 1995-96

In Thailand, the macroeconomic and structural weakness that was growing throughout the 1990s became more serious in 1995-96. The real GDP growth rate slowed down to 5.5 percent in 1996 from 8.9 percent in 1994. In addition, the current account deficit worsened from 5.6 percent of GDP in 1994 to 8.1 percent of GDP in 1996. These deficits had been financed by short-term capital inflows that led to a sharp accumulation of short-term debt which increased from 95.98 percent of foreign reserves in 1993-94 to 106.95 percent in 1995-96. By the end of 1996, the macroeconomic indicators of Thailand already showed very unstable conditions: large current account deficits, accumulation of short-term foreign debt, and low profitability of real investment projects.

In Indonesia, a sharp increase in the GDP growth rate to 15.9 percent in 1994 and 8.2 percent in 1995 brought along worrisome signs of overheating. Inflation remained high, while the country's trade surplus suffered a steep drop. The government's response of a slightly deflationary budget and a modest tightening of monetary policy was initially cautious. The government did not want higher interest rates to fuel further capital inflows and appreciate the currency. The Bank of Indonesia also widened the rupiah's trading band from 2 percent to 3 percent around the daily mid-rate, hoping that the additional trading risk of holding the rupiah would balance the incentive to invest in domestic assets provided by the higher interest rates. The band was further widened from 3 percent to 5 percent in June 1996, and again from 5 percent to 8 percent in September 1996.

The current account deficit had widened between 1994 and 1995 in Malaysia, as well, reaching 8.4 percent of GDP in 1995. Notably, in 1994 and 1995 foreign direct investment failed to cover the full amount of the deficit. During the effort to restrain domestic demand, the Malaysian interest rate had become too attractive to be ignored by foreign fund managers. In 1996, short-term debt sharply increased to 40.9 percent of foreign reserve compared to that of 30.6 percent in 1995.

A serious worsening of macroeconomic conditions already had occurred in South Korea between 1995 and 1996. The current account deficit rapidly widened from 1.5 percent of GDP in 1994 to 4.8 percent in 1996, leading to a record-breaking accumulation of short-term foreign debt. The 1996 growth rate of GDP decreased to 7.1 percent from the previous year's 8.9 percent. Reflecting weak financial conditions of the conglomerates, the stock market fell sharply in the two-year period 1995-96, down by 36 percent relative to the 1994 peak. The won also weakened during 1996.

Relative to the other countries in the region, economic conditions were more stable in the Philippines. Under IMF supervision, the Philippines experienced a sustainable GDP growth rate in the 1990's although lower than some of the southeastern Asian countries. The government's budget was in surplus. However, the current account deficit was large, and the currency had severely appreciated in real terms.

1.2 The unfolding of the crisis in 1997

By early 1997, macroeconomic conditions had deteriorated in most of the region. In the government's effort to defend collapsing financial institutions, strong speculative attacks on the baht forced Thailand to let the currency float on July 2, a crucial date in the chronology of the Asian crisis. Before the change of the exchange rate regime, Thailand used a highly managed exchange rate system which allowed a narrow band for the float of the exchange rate. Following Thailand, the Philippine central bank allowed the peso to move in a wider range against the dollar. Subsequently, the peso started to depreciate sharply.

As with Thailand, Malaysia had over a decade of extremely large current account deficits. Bank Negara announced ceilings on lending to the property sector and for purposes of stocks and shares in order to regulate a booming speculative bubble in real estate and equity lending. This caused foreign investors, led by US fund managers, to start selling their stocks. Under depreciation pressure, the Malaysian central bank abandoned its defense of the ringitt on July 14.

The Indonesian rupiah began to come under severe depreciation pressure with heavily increasing external debt. Failing in its defense, Indonesia abolished its system of managing the exchange rate through the use of a band and allowed it to float on August 14.

In early 1997, South Korea was shaken by a series of bankruptcies by large conglomerates that had heavily borrowed in previous years to finance their investment projects. The bankrupt conglomerates included Hanbo steel, Sammi steel and Kia. The macroeconomic indicators in early 1997 fully reflected the extent of this crisis; the current account deficit was increasing, export growth was falling, and industrial production growth rates were below previous levels. The speculative attack started in early November and South Korea requested IMF assistance on November 21. Finally the government announced it would allow the Won to float on December 16.

2. Movement of macroeconomic variables

2.1 Current account imbalances

As shown in Table 1, several Asian countries whose currencies sharply depreciated in 1997 had experienced somewhat sizable current account deficits in the 1990s. Thailand and Malaysia, both of which experienced deficits for over a decade, exhibit the largest and most persistent current account imbalances in our sample. The current account deficits in the two countries were over 6 percent of GDP on average between 1995 and 1996. The Philippines also experienced long-term imbalances. The deficit problem worsened in 1996. Starting the decade with a large imbalance, the current account imbalance of Indonesia increased to 3 percent of GDP between 1995 and 1996 although it shrank in 1992-93. In South Korea, the current account deficit was low in the early 1990s (1-3 percent of GDP) and virtually negligible in 1993. However, since 1993 the imbalance grew very fast, approaching almost 5 percent of GDP in 1996.

Fast-growing current account deficits likely increase currency depreciation pressure. The expanding current account deficits in these five countries could be considered as one of the factors forewarning a coming currency crisis.

2.2 Output growth

Table 2 presents the growth data in our sample of Asian countries in the 1990s. As shown in Table 2, GDP growth rates were remarkably high in the 1990s. Growth rates averaging more than 7 percent were the norm. But the growth rate slowed down in 1996, a year before the crisis. Only the Philippines, where growth rates were low in the early 1990s, geared up its growth rate to 6 percent in 1996 from 5 percent in 1995.

Accepting the traditional view that a large current account deficit is likely to be sustainable when growth is high, the Asian countries did not appear to have a sustainability problem until 1995. But the consumption and investment boom, as well as large capital inflows driven by overly optimistic beliefs that the economic expansion would persist, added instability to the value of currencies with a slowdown in the growth rate. In such conditions, an external shock that leads to a sudden change in expectations can cause a rapid reversal of capital flows and trigger a currency collapse.

2.3 Inflation

Table 3 presents inflation rates in our sample of Asian countries. In all countries, inflation rates were relatively low in the 1990s. The only exception was the Philippines where inflation was close to 20 percent in 1990-91(but falling to 8 percent by 1995). It is believed that high inflation rates leave fixed or semi-fixed exchange rate regimes potentially exposed to speculative attacks. The low inflation rates observed signal sound macroeconomic policy and sustainability of the regime. However the banking and financial sector problems experienced by several Asian countries over the 1990s raised considerable doubt about their ability to keep inflation low in the near future. These doubts were related to the possibility that the cost of the banking sector bail-outs might induce increased usage of seigniorage, and would require infusions of liquidity to prevent systemic runs.

2.4 Investment

Evidence on investment rates in Asian countries is shown in Table 4. Unlike the Latin American countries that experienced currency and financial crises in the recent past, the Asian countries were characterized by very high rates of investment throughout the 1990s. These rates were well above 30 percent of GDP in most countries, with the exception of the Philippines that had rates in the 20-25 percent range.

Despite the high investment rate of the Asian countries, the profitability of investment –the ratio between the investment rate and the rate of output growth- given by Table 5 suggests the efficiency of investment was falling in the three years, 1994-1996, prior to the 1997 crisis with the exception of the Philippines. Also the investment boom was confined to the non-traded sector (commercial and residential construction, as well as inward-oriental services) adding an unsustainable factor to the sharply growing current account deficit.

2.5 Savings

Data on saving rates in Asia are reported in Table 6, and to some extent stand for the mirror of the investment rates in Table 4. Asian countries were characterized by very high savings rates throughout the 1990s- in many cases above 30 percent of GDP and in some cases above 40 percent. Looking at the data in Table 7 before the crisis, there is little evidence of public dissaving so that the current account imbalances do not appear to be the result of increased public sector deficits. The absence of fiscal imbalances in the years preceding the crisis, however, should not be regarded as pervasive evidence against the fiscal roots of the Asian crisis. The pre-crisis years were a period of excessive credit growth in the banking system, leading to a large stock of non-performing loans and the eventual collapse of several financial institutions. The cost of restructuring the financial sector could have been an implicit fiscal liability for the Asian countries. Such a liability was not reflected by data on public deficits until the outbreak of the crisis, but affected the sustainability of the pre-crisis current account imbalances since it generated expectations of radical policy changes or currency devaluations.

2.6 Real exchange rate appreciation

A significant real exchange rate appreciation may be associated with a loss of competitiveness and a structural worsening of the trade balance, thus weakening the sustainability of the current account. Data on the real exchange rate of the Asian countries in Table 8[2] shows that the real exchange rate had appreciated by 15.7 percent in Malaysia, 26.1 percent in the Philippines, 8.0 percent in Indonesia, and 0.5 percent in Thailand by the end of 1996, taking 1990 as the base year. In South Korea, the currency depreciated in real terms by 9.2 percent. This suggests that, with the significant exception of South Korea, all the currencies that crashed in 1997 had experienced real appreciation. It should be stressed that in a number of countries, a large part of the real appreciation occurred after 1995, in parallel with the strengthening of the US dollar. The sharp increase of the US dollar relative to the Japanese yen and the European currencies since the second half of 1995 led to deteriorating competitiveness in most Asian countries whose currencies were effectively pegged to the dollar.

3. Structural conditions in the financial system

3.1 Weak banking system

In the 1990s the countries of East Asia performed a financial deregulation and capital liberalization. The financial liberalization involved loosening restrictions on both interest rate ceilings and the type of lending allowed. Bank lending increased sharply prior to the crisis, with much of it financed by inflows of international capital.

Of course, the problem was not that lending expanded, but rather that it expanded so rapidly that excessive risk-taking occurred. In fact, the increasing proportion of non-performing loans indicates that many of the loans made by banks were invested in risky and low profitable projects or used for real estate, non-traded area. Therefore, they weakened the banking system.

3.1.1 Lending boom

As shown in Table 9, domestic bank lending to the private sector shows a steep upward trend in the five countries prior to the crisis. The most extreme case was the Philippines, where banking claims on the private sector, as a percent of GDP, increased by more than 60 percent between 1994 and 1996. It was also large in Malaysia (25 percent) and Thailand (12 percent). Though more modest in Indonesia and South Korea, the magnitude of credit growth between 1994 and 1996 was much higher than in the early 1990s.

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[2] The source of these data is the JP Morgan RER series that go back to 1970; the base year for the trade weight is

3.1.2 Accumulation of bad loans

One of the main problems faced by the Asian countries was that many of the loans made by banks were invested in risky and unprofitable projects or used for real estate, property and the purchase of equity funds. A possible indicator of investment in risky and low profitable projects is the proportion of non-performing loans (NPLs) in total loans (Table 10). Since the 1997 crisis may have crippled otherwise healthy loans, it is appropriate to refer to data at the onset of the crisis. In 1996, the NPLs were estimated at 8-14 percent for the five afflicted countries. For the purpose of comparison, the estimated NPLs were 3-4 percent in Hong Kong, Singapore and Taiwan.

3.1.3 Loans financed by foreign liabilities

Large increases in the foreign liabilities of banks point out that much of the bank lending was mostly financed by borrowing from abroad (Table 11). In the Philippines, foreign liabilities soared from 5.5 percent of GDP at the end of 1993 to 17.4 percent of GDP three years later. In South Korea, the corresponding liabilities of the banking system more than doubled from 4.5 percent of GDP in December 1993 to 9.4 percent of GDP in December 1996. In Thailand, foreign liabilities jumped more sharply from 5.9 percent of GDP in 1992 to 26.6 percent in 1996. In Indonesia, though the liabilities remained at a more modest level, much of the offshore borrowing was undertaken directly by private firms. The only exception was Malaysia where foreign liabilities fell off sharply in 1994 and slightly increased to 11.4 percent in December 1996.

1990.

3.2 Imbalances in foreign debt accumulation and management

3.2.1 Rising share of short-term debt

If a large fraction of a country's external liabilities are short-term, a crisis may take the form of a pure liquidity shortfall – the inability by a country to roll over its short-term liabilities. The experiences of Mexico with its short-term public debt in 1994-95 and of several Asian countries with private external liabilities in 1997 provide examples of liquidity problems. The figures corresponding to the ratio of short-term debt to foreign reserves are presented in Table 12. All the countries except the Philippines have somewhat increasing ratio after 1993. In South Korea, the ratio sharply grew from 54.1 percent to 171.5 percent in 1995.

3.2.2 Foreign Exchange Reserves

Large foreign exchange reserves facilitate the financing of a current account deficit and enhance the credibility of a fixed exchange rate policy. Foreign exchange reserves and a small external debt burden reduce the risk of external crises, and enable a country to finance a current account deficit at lower costs. The real rate paid on the debt indicates the market's evaluation of the country's ability to sustain a current account deficit.

To measure the sufficiency of foreign exchange reserves, the ratio of money assets to foreign reserves is considered since in the event of an exchange rate crisis, all liquid money assets can potentially be converted into foreign exchange. Calvo(1998) suggests using the ratio of a broad measure of liquid monetary assets to foreign reserves, for instance, the ratio of M2 to foreign reserves.

Table 13 reports the ratio of M2 to foreign reserves. In most Asian countries the ratio was unusually high in 1996-97. In Indonesia, the ratio constantly rose throughout the 1990s and reached a peak as high as 7.1 in 1995. In South Korea, M2/FX was equal to 6.5 before1997, and rose to 10.5 by the end of 1997. In Malaysia, the ratio increased from 2.9 in 1990 to 3.7 at the end of 1996. In the Philippines, the ratio declined marginally from 4.8 in 1991 to 4.5 in 1996. In Thailand, the ratio went from 4.5 in 1990 to 3.9 in 1996.

Table 14 indicates that the ratios of most G7 countries are very high compared to those Asian countries. In addition, the ratios of all G7 countries except Japan also rose throughout the 1990s even though their currencies were not attacked by speculative agents in 1997. Therefore, the M2/FX needs further empirical study to verify whether it is an appropriate indicator of currency crisis.

4. A Debate between "weak fundamentals and financial panic" analysts

4.1 Evidence presented by financial panic analysts

Financial panic supporters accept that warning signs such as current account deficits, bad loans and overinvestment were reasons for financial weakness in the five countries. But they maintain that those signs were not enough to warrant the magnitude of the Asian crisis.

It is generally believed that some aspects of the real economy in at least some crisis economies were solid. Government budgets, which were at the center of economic crises in Latin America in the 1980s, indicated regular surpluses in each Asian country during the 1990s, as shown in Table 7. GDP growth rates were very high in the 1990s as well.

Although some analysts expected the possibility of a crisis[3], such warnings were unusual. Inflow of capital remained strong through 1996 and, in most cases, until mid 1997. The only exception is found in the stock markets in Thailand and South Korea, where foreign investors became uneasy in 1996, as shown in Table 15. In Malaysia, though stock markets began a rather steep decline in March 1997, bank lending continued to be very strong at least until mid-year. In Indonesia, both the stock market and bank lending remained strong until mid-1997.

Credit rating agencies, such as Standard & Poor's and Moody, provide an ongoing assessment of credit risk in emerging markets. If the market had expected a financial crisis and public sector bailouts, the ratings of sovereign bonds should have fallen in the pre-crisis period. However, the rating agencies did not signal any risk until after the onset of the Asian crisis. Long-term debt ratings remained unchanged throughout 1996 and the first half of 1997 for each of the Asian countries except the Philippines, where the debt rating was actually upgraded in early 1997. In each country, the outlook was described as "positive" or "stable" through June 1997. Only until weeks after the crisis started, did agencies downgrade the region's debt.

Another measure of expectations for the region may be found in IMF reports. The IMF gave very little indication that there was any macroeconomic risk to the Asian

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[3] See, for example, Park (1996)

region. For example, *World Economic Outlook* (IMF, December 1997) predicted 6 percent growth for South Korea in 1998, 7 percent for developing Asia (or 5 percent for developing Asia excluding China and India).

4.2 Evidence presented by weak fundamental analysts

Advocates of the weak fundamentals hypothesis challenge the financial panic view. First, they assert that credit ratings have had no informational value in forecasting currency crises for the last 20 years. As credit ratings failed to predict the crises of the early 1980s, in which fundamentals were obviously at work, one cannot suppose that their failure to predict the Asian crisis is evidence that the crisis was due to financial panic.

Second, IMF reports are not generally informative in predicting a crisis. Given the ability of IMF reports to sharply affect markets, such reports are always written in terms that express concern in very cautious terms.

Another piece of evidence is that countries with more sound fundamentals were spared the most serious collapses. In fact, Taiwan, Singapore, Hong Kong, and China were less affected by the regional turmoil.

5. Summary

Two main causes of the Asian currency crisis, financial panic and weak fundamentals, have emerged in recent debate.

Financial panic supporters admit that there were crucial underlying problems in the Asian economies, but assert that these problems were not severe enough to warrant a financial crisis of such large magnitude. In contrast, weak fundamental analysts stress the significant role of weak fundamentals in causing the crisis. Given the weakness of the macroeconomic stances in the afflicted countries, the crisis was an inescapable outcome rather than just a financial panic.

While the evidence given in this chapter reflects that deterioration in macroeconomic fundamentals and poor economic policies could be a root of the crisis, it does not strongly convince us that fundamentals had deteriorated severely enough to make the crisis inevitable.

In summation, the empirical evidence on the causes of the Asian crisis is not conclusive, demanding more formal studies on the issue.

Table 1. Current account (% of GDP), balance of payment definition

	'90	'91	'92	'93	'94	'95	'96	'97	'98	'99	'00
S. Korea	-0.69	-2.83	-1.28	0.30	-1.02	-1.86	-4.75	-1.85	12.73	6.01	2.65
Indonesia	-2.82	-3.65	-2.17	-1.33	-1.58	-3.18	-3.37	-2.24	4.35	4.23	4.88
Malaysia	-2.03	-8.69	-3.74	-4.66	-6.24	-8.43	-4.89	-4.85	13.14	16.01	9.33
Philippines	-6.08	-2.28	-1.89	-5.55	-4.60	-2.67	-4.77	-5.23	2.35	10.31	12.28
Thailand	-8.50	-7.71	-5.66	-5.08	-5.60	-8.06	-8.10	-1.90	12.71	10.61	7.24

Source: IMF, *International Financial Statistics*, available on CD-Rom.

Table 2. GDP growth rate (%)

	'90	'91	'92	'93	'94	'95	'96	'97	'98	'99	'00
S. Korea	9.51	9.13	5.06	5.75	8.58	8.94	7.10	5.47	-6.69	10.66	9.53
Indonesia	7.24	6.95	6.46	6.50	15.93	8.22	7.98	4.65	-13.2	0.13	5.61
Malaysia	9.74	8.48	7.80	8.35	9.24	9.46	8.58	7.81	-7.37	5.64	8.71
Philippines	3.03	-0.58	0.34	2.12	4.38	4.77	5.76	9.66	-0.59	3.32	3.95
Thailand	11.63	8.18	8.08	8.38	8.94	8.84	5.52	-0.43	-10.2	3.33	4.60

Source: IMF, *International Financial Statistics*, available on CD-Rom.

Table 3. Inflation rate

	'90	'91	'92	'93	'94	'95	'96	'97	'98	'99	'00
S. Korea	8.58	9.03	6.22	4.82	6.24	4.41	4.96	4.45	7.48	0.85	2.27
Indonesia	7.76	9.40	7.59	9.60	12.56	8.95	6.64	11.62	57.73	20.47	3.70
Malaysia	2.67	4.40	4.69	3.57	3.71	5.28	3.56	2.66	5.27	2.77	1.57
Philippines	14.16	18.70	8.93	7.58	9.06	8.11	8.41	5.01	9.71	6.71	4.29
Thailand	5.93	5.70	4.07	3.36	5.19	5.69	5.85	5.61	7.96	0.25	1.57

Source: IMF, *International Financial Statistics*, available on CD-Rom.

Table 4. Investment rate (% of GDP)

	'90	'91	'92	'93	'94	'95	'96	'97	'98	'99	'00
S. Korea	36.93	38.90	36.58	35.08	36.05	37.05	38.42	34.97	21.17	26.83	28.30
Indonesia	36.15	35.50	35.87	29.48	31.06	31.93	30.80	31.60	16.77	12.19	17.87
Malaysia	31.34	37.25	33.45	37.81	40.42	43.50	41.54	42.84	26.69	22.43	26.96
Philippines	24.16	20.22	21.34	23.98	24.06	22.22	24.02	24.84	20.24	18.63	17.61
Thailand	41.08	42.84	39.97	39.94	40.27	41.61	41.73	34.99	20.32	19.94	*NA*

Source: IMF, *International Financial Statistics*, available on CD-Rom.

Table 5. Incremental capital output ratio

	'90	'91	'92	'93	'94	'95	'96	'97	'98	'99	'00
S. Korea	3.88	4.26	7.22	6.10	4.20	4.14	5.41	6.38	-*	2.51	2.97
Indonesia	4.99	5.11	5.55	4.53	1.94	3.88	3.85	6.80	-*	93.8**	3.19
Malaysia	3.21	4.39	4.29	4.53	4.37	4.59	4.84	5.49	-*	3.98	3.10
Philippines	7.97	-*	62.8**	11.31	5.49	4.65	4.17	2.57	-*	5.61	4.46
Thailand	3.53	5.24	4.95	4.77	4.50	4.71	7.56	-*	-*	5.97	*NA*

Source: IMF, *International Financial Statistics*, available on CD-Rom.

* GDP growth rate is negative. ** Due to small GDP growth rates, these numbers are large.

Table 6. Saving rate (% of GDP)

	'90	'91	'92	'93	'94	'95	'96	'97	'98	'99	'00
S. Korea	35.69	35.74	34.88	34.91	34.60	35.14	33.60	33.06	34.46	33.65	32.40
Indonesia	31.75	31.10	33.41	28.66	29.52	27.65	27.50	27.98	26.52	20.20	25.72
Malaysia	29.07	23.24	30.06	27.70	33.81	34.65	37.81	39.34	48.48	47.20	46.85
Philippines	17.85	17.76	18.16	17.29	20.32	17.16	19.35	18.77	13.64	19.65	22.03
Thailand	32.33	34.83	33.73	34.26	33.89	33.25	33.22	32.64	36.20	32.57	*NA*

Source: IMF, *International Financial Statistics*, available on CD-Rom.

Table 7. Government budget balance (% of GDP)

	'90	'91	'92	'93	'94	'95	'96	'97	'98	'99	'00
S. Korea	-0.68	-1.63	-0.50	0.64	0.32	0.30	0.46	0.25	-3.82	-4.63	1.10
Indonesia	0.43	0.45	-0.44	0.64	1.03	2.44	1.26	-0.67	-2.82	-1.15	-1.10
Malaysia	-3.10	-2.10	-0.89	0.23	2.44	0.89	0.76	2.52	-1.75	-3.17	-4.20
Philippines	-3.47	-2.10	-1.16	-1.46	1.04	0.57	0.28	0.06	-1.87	-3.73	-5.72
Thailand	4.59	4.79	2.90	2.13	1.89	2.94	0.97	-0.32	-2.78	-3.28	-5.40

Source: IMF, *International Financial Statistics*, available on CD-Rom.

Table 8. Real exchange rate. End of year data

	'90	'91	'92	'93	'94	'95	'96	'97	'98	'99	'00	'01
S. Korea	97.2	92.0	89.2	87.1 (-10)*	86.7	90.1	88.3 (-9.2)	58.2	71.3	75.5	75.7	74.2
Indonesia	97.5	99.3	100.5	103.5 (6.2)	100.7	100.3	105.3 (8.0)	62.4	70.4	78.4	66.7	72.5
Malaysia	97.1	96.7	109.5	110.7 (14.0)	107.1	106.7	112.3 (15.7)	85.0	83.3	83.0	89.6	95.6
Philippines	92.4	103.1	107.1	97.4 (5.4)	111.6	109.6	116.5 (26.1)	89.8	83.2	85.3	80.7	96.2
Thailand	102.3	98.8	99.5	101.7 (-0.6)	98.1	98.6	102.8 (0.5)	74.2	87.1	85.5	83.3	84.6

Source: J.P. Morgan, **http://www.jpmorgan.com**

*The appreciation rate of the real exchange rate taking 1990 as a base year.

Table 9. Bank lending to private sector (% of GDP)

	'90	'91	'92	'93	'94	'95	'96	'97	'98	'99	'00
S. Korea	52.54	52.81	53.34	54.21 (3.2)*	56.84	57.04	61.81 (8.7)**	69.79	71.71	79.35	87.60
Indonesia	49.67	50.32	49.45	48.90 (-1.6)	51.88	53.48	55.42 (6.8)	69.23	53.94	20.95	20.94
Malaysia	71.36	75.29	74.72	74.06 (3.8)	74.61	84.80	93.39 (25.2)	106.91	100.08	97.95	91.62
Philippines	19.17	17.76	20.44	26.37 (37.6)	29.06	37.52	48.98 (68.6)	56.53	47.76	41.70	39.63
Thailand	64.30	67.70	72.24	80.01 (24.4)	91.00	97.62	101.94 (12.0)	116.33	114.31	106.63	82.55

Source: IMF, *International Financial Statistics*, available on CD-Rom.

*Increasing rate between 1990 and 1993 ** Increasing rate between 1994 and 1996

Table 10. Non-performing loans (% of total lending)

	'96		'96
S. Korea	8%	Philippines	14%
Indonesia	13%	Singapore	4%
Malaysia	10%	Hong Kong	3%
Thailand	13%	Taiwan	4%

Source: Corsetti, Pesenti and Roubini (1998c)

Table 11. Foreign liabilities of the banking system (% of GDP)

	'90	'91	'92	'93	'94	'95	'96	'97	'98	'99	'00
S. Korea	4.06	4.86	4.81	4.48	5.46	6.93	9.35	11.26	7.98	6.48	6.01
Indonesia	6.46	5.25	5.74	6.20	6.51	5.93	5.59	11.22	10.37	9.34	7.18
Malaysia	7.02	8.88	12.58	19.06	8.93	7.26	11.39	17.44	12.24	9.27	7.79
Philippines	6.18	4.40	5.56	5.47	6.69	8.83	17.39	25.41	18.60	16.11	15.50
Thailand	5.01	4.94	5.92	11.08	21.46	27.75	26.64	40.73	23.00	15.27	11.09

Source: IMF, *International Financial Statistics*, available on CD-Rom.

Table 12. Short-term debt (% of foreign reserves)

	'90	'91	'92	'93	'94	'95	'96	'97	'98	'99	'00
S. Korea	72.13	81.75	69.62	60.31	54.06	171.5	203.2	307.5	58.71	33.67	24.55
Indonesia	149.3	154.6	172.8	159.7	160.4	189.4	176.6	188.0	84.51	66.53	82.77
Malaysia	19.54	19.05	21.12	25.51	24.34	30.60	40.98	75.57	39.42	21.93	16.76
Philippines	479.1	152.3	119.4	107.7	95.00	82.85	79.45	163.4	95.09	48.52	36.14
Thailand	62.55	71.31	72.34	92.49	99.48	114.2	99.69	179.5	97.23	34.96	25.32

Source: OECD, External Debt Statistics.

Table 13. Ratio of M2 to foreign reserves, Asian countries

	'90	'91	'92	'93	'94	'95	'96	'97	'98	'99	'00
S. Korea	6.48	8.33	7.20	6.91	6.45	6.11	6.51	10.50	3.55	3.74	3.80
Indonesia	6.16	5.51	5.61	6.09	6.55	7.09	6.50	7.37	2.50	3.08	3.91
Malaysia	2.91	2.99	2.64	2.09	2.47	3.33	3.66	4.99	2.70	2.73	3.10
Philippines	16.33	4.82	4.35	4.90	4.86	5.86	4.50	6.97	4.33	3.70	3.58
Thailand	4.49	4.10	4.10	4.05	3.84	3.69	3.90	5.29	3.99	3.89	4.04

Source: IMF, *International Financial Statistics*, available on CD-Rom.

Table 14. Ratio of M2 to foreign reserves, G7 countries

	'90	'96
Germany	15.76	18.29
France	22.31	38.43
United Kingdom	28.69	34.34
Italy	11.96	15.89
Canada	15.77	18.03
Japan	49.92	22.35
United States	52.14	70.35

Source: IMF, *International Financial Statistics*, available on CD-Rom.

Table 15. Stock market prices indices*

	'90	'91	'92	'93	'94	'95	'96	'97	'98	'99	'00	'01
S. Korea	696	610	678	866	1027	882	651	376	562	1028	505	694
Indonesia	417	247	274	588	469	513	637	401	398	677	416	392
Malaysia	505	556	643	1275	971	995	1237	594	586	812	680	696
Philippines	651	1151	1256	3196	2785	2594	3170	1869	1969	2143	1495	1140
Thailand	612	711	893	1682	1360	1280	831	372	356	482	269	304

Source: Corsetti, Pesenti and Roubini (1998c), Korea Stock Exchange, Jakarta Stock Exchange, Kuala Lumpur Stock Exchange, Philippine Stock Exchange and The Stock Exchange of Thailand

* End of period

Table 16. The change of macroeconomic conditions in the Asian countries before the currency crisis

Country	1995-96	1997
Thailand	- Real GDP growth rate slowed down to 5.5 % in 1996 from 8.9% in 1994. - Current account deficit worsened from 5.6% of GDP in 1994 to 8.1% of GDP in 1996. - Short-term debt (% of foreign reserves) increased from 95.98% in 1993-94 to 106.95% in 1995-96.	- The government increased its effort to defend collapsing financial institutions in early 1997. - Strong speculative attack on the bath forced Thailand to let the currency float on July 2.
Indonesia	- Real GDP growth rate decreased from 15.9% to 8.0% in 1996. - Inflation rate remained high at 8.0% in 1996. - Surplus of government budget balance (% of GDP) decreased from 2.4% in 1995 to 1.3% in 1996. - The rupiah's trading band was widened from 5% to 8% in September 1996	- The Indonesian rupiah began to come under severe depreciation pressure with heavily increasing short-term debt in early 1997. - Indonesia allows the managed exchange rate to float on August 14.
Malaysia	- Current account deficit widened in 1994-95, reaching 8.4% of GDP in 1995. - Short-term debt (% of foreign reserves) increased to 40.9% compared to that of 30.6 % in 1995.	- Under the severe depreciation pressure caused by lending boom and current account deficit, the Malaysian central bank abandoned the defense of the ringitt on July 14.

Country	1995-96	1997
South Korea	- Current account deficit widened from 1.5% of GDP in 1994 to 4.8% in 1996. - Short-term debt (% of foreign reserves) increased from 54.1% in 1994 to 203.2% in 1996. - Growth rate of real GDP decreased from 8.9% in 1995 to 7.1% in 1996 - The stock market fall sharply in 1995-96, down by 36% relative to the 1994 peak	- Current account deficit was increasing, export growth was falling, and industrial production growth rates were below previous levels in early 1997. - The government announced it allowed Korean Won to float on December 16.
Philippines	- Current account deficit continued to be high, 4.8% in 1996. - Real exchange rate severely appreciated by 26.2 % relative to that in 1990.	- Following Thailand, the Philippine central bank allowed the peso to move in a wider range against the dollar.

CHAPTER III

SURVEY OF LITERATURE AND EXTENDED MODEL

Currency crises in Europe, Mexico and Asia in the 1990's have drawn global attention to speculative attacks on government-controlled exchange rates. Research has proceeded on both theoretical and empirical fronts to find the roots and results of the crises. This chapter will review the research and extend the currency crisis model for further research.

1. Theoretical literature

1.1 First generation models

Initial models, now called first generation research, were developed in response to currency crises in developing countries such as Mexico (1973-82) and Argentina (1978-81).

The first model came from Krugman(1979). Krugman showed that, under a fixed exchange rate regime, domestic credit creation caused by the fiscal deficit of government in excess of money demand growth leads to a gradual loss of reserves. Therefore, this loss leads to a speculative attack against the currency that forces the abandonment of the fixed exchange rate and the adoption of a flexible rate regime- a phenomenon known as the "peso problem". Because of the nonlinearities involved in his model, however, Krugman was unable to explicitly derive a solution for the time of collapse in a fixed exchange rate regime. Later work done by Flood and Garber (1984a) provided an example of how such a solution can be derived in a linear model, with or without arbitrary speculative behavior.

The basic framework in the following subsection used a simple continuous time, perfect foresight model. This framework explains the Krugman-Flood-Garber insight clearly. The model in the framework is a log-linear formulation that allows us to solve explicitly for the time of occurrence of the crisis, by assuming initially that the exchange rate is allowed to float permanently in the postcollapse regime.

1.1.1 Basic framework

In this subsection, Flood and Garber's model (1984a) is introduced as the basic framework to analyze the first generation approach to a balance of payments crisis.

Consider a small open economy where residents consume a single, tradable good whose domestic supply is exogenously fixed at $\bar{y}$. The good is perishable and its foreign currency price is fixed (at unity). Purchasing power parity holds, so that the domestic price level is equal to the nominal exchange rate. Three assets are available: domestic money (held by domestic residents only), domestic bonds and foreign bonds. Domestic and foreign bonds are perfect substitutes. There are no private banks, so that the money supply is equal to the sum of domestic credit issued by the central bank and the domestic currency value of foreign reserves held by the central bank, which earn no interest. Domestic credit is assumed to expand at a constant growth rate. Finally, agents have perfect foresight.

Formally, the model is defined as follows.

$$m_t - p_t = \alpha_0 - \alpha_1 i_t, \ \alpha_1 > 0 \qquad (1)$$

$$m_t = \gamma d_t + (1-\gamma) r_t, \ 0 < \gamma < 1 \quad (2)$$

$$\dot{d}_t = \mu, \ \mu > 0, \qquad (3)$$

$$p_t = p_t^* - s_t \qquad (4)$$

$$i_t = i^* + E_t \dot{s}_{t+1} . \qquad (5)$$

All variables, except interest rates, are measured in logarithms. m_t, p_t and i_t are the domestic money stock, price level and interest rate, respectively. d_t and r_t denote domestic credit and the domestic government book value of foreign money holdings, respectively. s_t is the spot exchange rate, i.e. the domestic money price of foreign money. An asterisk (*) indicates "foreign variables", assumed to be constant. A dot over a variable denotes a time derivative.

Equation (1) defines real money demand as a negative function of the domestic interest rate. Equation (2) is a log-linear approximation of the identity linking the money stock to reserves and domestic credit. Equation (3) assumes that domestic credit grows at the rate μ. Purchasing power parity and uncovered interest rate parity are defined in equations (4) and (5), respectively.

Under perfect foresight, $E_t \dot{s}_{t+1} = \dot{s}_{t+1}$. Setting $\alpha_0 = i^* = 0$ and substituting (2), (4), and (5) into (1) yields the following money market equilibrium condition:

$$m_t = \gamma d_t + (1-\gamma) r_t = s_t - \alpha_1 \dot{s}_{t+1} . \qquad (6)$$

Equation (6) states that the demand for money can be satisfied either from reserves or domestic credit.

If the exchange rate is fixed at $\bar{s}$, then $\dot{s}_t = 0$. When domestic credit grows, reserves, at any time t, adjust to maintain money market equilibrium, according to the following rule:

$$r_t = [\bar{s} - \gamma d_t]/(1-\gamma) \ . \qquad (7)$$

The rate of change in reserves is obtained from (3) and (7):

$$\dot{r}_{t+1} = -\mu/\theta, \quad \theta = (1-\gamma)/\gamma \ . \quad (8)$$

Equation (8) shows that if domestic credit grows while money demand remains unchanged, reserves are run down at a rate proportional to the rate of credit expansion. Clearly the country will run out of reserves eventually and a fixed exchange rate regime cannot survive forever.

To find the time of the attack, Flood and Garber introduce the idea of the shadow exchange rate, which is defined as the floating exchange rate that would prevail if speculators purchased the remaining government reserves committed to the fixed rate. After reserves reach the lower bound, the government refrains from foreign exchange market intervention and allows the exchange rate to float freely and permanently thereafter. The shadow exchange rate $\tilde{s}$, therefore, is the exchange rate that balances the money market following an attack in which foreign exchange reserves are exhausted.

Substituting the trial solution $\tilde{s}_t = \lambda_0 + \lambda_1 m_t$ into equation (6), they find that $\lambda_0 = \alpha_1 \gamma \mu$ and $\lambda_1 = 1$. Thus,

$$\tilde{s}_t = \alpha_1 \gamma \mu + m_t . \qquad (9)$$

Equation (9) implies that the shadow exchange rate depreciates steadily and proportionally to the rate of growth of domestic credit. For simplicity, they assume that

r_f=0 when the fixed exchange rate regime is abandoned. Noting that $d_t = d_0 + \mu t = m_t / \gamma$, they obtain

$$\tilde{s}_t = \gamma(d_0 + \alpha_1 \mu) + \gamma \mu t. \quad (10)$$

Figure 1 plots $\tilde{s}$ in equation (9) and the pre-attack fixed exchange rate, $\bar{s}$. Let d^E denote the domestic credit level at point A where $\tilde{s}$ is equal to $\bar{s}$, i.e. the two lines intersect when $d = d^E$.

Suppose that d is smaller than d^E. If speculators attack at level d, then the currency will appreciate and the speculators will experience a capital loss on the reserves they purchase from the government. Thus, there will be no attack when $d < d^E$. Suppose instead that $d > d^E$, so $\tilde{s} > \bar{s}$. Now there is a capital gain to speculators for every unit of reserves purchased from the government. Speculators can forecast when that capital gain will be acquired and compete against each other for the profit. The way they compete in this framework is to get a jump on each other and attack earlier. Such competition continues until the attack is driven back in time to the point where $d = d^E$. As a consequence, arbitrage in the foreign exchange market fixes the exchange rate immediately after the attack to equal the fixed rate prevailing at the time of the attack. Exchange rate jumps are ruled out by speculative competition.

The condition that $\tilde{s} = \bar{s}$ is used to find both the timing of the attack and the extent of government reserve holdings at the time of the attack. Substituting $\bar{s}$ for $\tilde{s}$ into equation (10) and $\dot{s} = 0$ produces the timing of the attack T:

$$T = \frac{\bar{s} - \gamma d_0}{\gamma \mu} - \alpha_1 = \frac{\theta r_0}{\mu} - \alpha_1. \quad (11)$$

Equation (11) shows that the higher the initial stock of reserves, R_0, or the lower the rate of credit expansion, the longer it takes before the fixed exchange rate regime collapses. The (semi-) interest rate elasticity of money demand determines the size of the downward shift in money demand and reserves when the fixed exchange rate regime collapses and the nominal interest rate jumps to reflect an expected depreciation of the domestic currency. The larger α is, the earlier the crisis. Finally, the larger the initial proportion of domestic credit in the money stock (the higher γ), the sooner the collapse. This result is due to the fact that Flood and Garber's model is based on the monetary approach. According to the approach, a rise in domestic money supply, with demand for money remaining unchanged would ultimately be offset by an equal and opposite change in the international reserves through the balance of payments. When reserves run out, the fixed exchange rate is abandoned.

1.1.2 Extensions to the basic framework

The basic theory of balance of payment crises presented above has been extended in various directions. I emphasize two models that are strongly related to the empirical work in the following chapters: One introduces uncertainty into the above context and the other examines the real effects of a currency crisis in a model with endogenous output, sticky forward-looking wage contracts, and external trade.[4]

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[4] For details of other major extensions, refer to the survey of first generation models by Agenor, Bhandari and Flood (1992).

Uncertainty and the probability of attack

In the basic model developed above, it has been assumed that there is some binding threshold level, known by all agents, below which foreign reserves are not allowed to be depleted. The attainment of this critical level implies a regime shift from a fixed exchange rate regime to a floating rate regime. In practice, however, agents are only imperfectly informed of central bank policies. They may not exactly know the threshold level of reserves that triggers the regime shift. If uncertainty about current and future government policy is prevalent, the assumption of perfect foresight may be improper.

An implication of the perfect foresight model developed above, which is contradicted empirically, concerns the behavior of the domestic nominal interest rate. In the model the nominal interest rate stays constant until the moment the attack occurs-at which point it jumps to a new level consistent with the postcollapse regime. Uncertainty over the depreciation rate, as modeled below, may help to account for a rising interest rate in the transition period. Indeed, while specific results are sensitive to arbitrary specifications regarding distributional assumptions of random terms, only stochastic models are consistent with the large interest rate fluctuations observed in actual cases.

Uncertainty about domestic credit growth was first introduced by Flood and Garber (1984a) in a discrete time stochastic model. In their framework, domestic credit is assumed to depend on a random component. In the basic model, equations (3) and (5) are modified as follows:

$$d_t = d_{t-1} + \mu + \varepsilon_t \qquad (3)'$$

$$i_t = i_t^* + E_t s_{t+1} - s_t . \qquad (5)'$$

Variables common to equations (1)-(5) are defined as before, except that t is now an integer. In equation (3)', ε_t represents a random disturbance in domestic credit growth. Equation (5)' introduces notation $E_t(\cdot)$, the mathematical expectation operator conditional on the information set available at time t.

Let $\bar{s}$ and $\tilde{s}_t$ denote the fixed exchange rate and the shadow exchange rate as before. In each period, the probability of collapse in the next period is found by evaluating the probability that domestic credit in the next period will be sufficiently large to result in a discrete depreciation, should a speculative attack occur. In the Flood-Garber framework a fixed rate regime will collapse whenever it is profitable to attack it. The condition for profitable attack is, as in the model developed above, that the postcollapse exchange rate, s_t, be larger than the prevailing fixed rate, $\bar{s}$. Profits of speculators are equal to the exchange rate differential multiplied by the reserve stock used to defend the fixed rate regime. Since these are risk-free profits earned at an infinite rate (speculators could always sell foreign exchange back to the central bank at the fixed rate if the attack is unsuccessful), the system will be attacked if and only if $\tilde{s}_{t+1} > \bar{s}$. Therefore, the probability at time t of an attack at time $t+1$, ${}_t\pi_{t+1}$, is given by

$$ {}_t\pi_{t+1} = prob(\tilde{s}_{t+1} > \bar{s}). \qquad (12)$$

The unconditional expected future exchange rate is a probability-weighted average of $\bar{s}$ and $\tilde{s}_{t+1}$.

$$E_t s_{t+1} = [1 - {}_t\pi_{t+1}]\bar{s} + {}_t\pi_{t+1} E_t(\tilde{s}_{t+1} | \tilde{s}_{t+1} > \bar{s}). \quad (13)$$

Rearranging (13) yields:

$$E_t s_{t+1} - \bar{s} = {}_t\pi_{t+1}[E_t(\tilde{s}_{t+1} | \tilde{s}_{t+1} > \bar{s}) - \bar{s}] . \qquad (14)$$

This equation provides the economic intuition of rising interest rates prior to a crisis.[5] According to the uncovered interest rate parity condition, the left side in equation (14) equals the interest rate differential between the domestic and foreign interest rate. Since the foreign interest rate is assumed to be constant, increases in $[E_t s_{t+1} - \bar{s}]$ would correspond to higher domestic interest rates.

The expected rate of exchange rate depreciation, $[E_t s_{t+1} - \bar{s}]$, increases prior to the collapse because both ${}_t\pi_{t+1}$ and $[E_t(\tilde{s}_{t+1} | \tilde{s}_{t+1} > \bar{s}) - \bar{s}]$ rise with the approach of the crisis. The probability of an attack next period ${}_t\pi_{t+1}$ rises because the increasing value of the state variable (domestic credit) makes it increasingly likely that an attack will take place at $t+1$. The quantity $[E_t(\tilde{s}_{t+1} | \tilde{s}_{t+1} > \bar{s}) - \bar{s}]$ gives the gain that agents may expect given that there will be a speculative attack at $t+1$. In turn, that gain depends on the value agents expect for the state variable next period, given that an attack will occur at $t+1$. As the value of the state variable rises from period to period, its conditional expectation also rises.

The introduction of uncertainty has important implications. First, the transition to a floating regime is stochastic, rather than certain. The collapse time becomes a random variable and cannot be determined explicitly, since the timing of a possible future speculative attack is unknown. Second, there is always a nonzero probability of a

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[5] For an explicit solution, refer to Flood and Garber (1984a).

speculative attack in the next period, which, in turn, produces a forward premium in foreign exchange markets. Third, the degree of uncertainty about the central bank's credit policy plays an important role in the speed at which reserves of the central bank are depleted. In the stochastic setting, reserve losses exceed increases in domestic credit because of a rising probability of regime shift, so that reserve depletion accelerates on the way to a regime change- a pattern that has often been observed in actual crises.

Real effects of crises

The early literature on currency crises emphasized the financial aspects of crises and overlooked real events that were occurring at the same time. Evidence suggests, however, that currency crises have been often preceded by large current account deficits or economic depression.

The real effects of a potential exchange rate crisis have been investigated by Flood and Hodrick (1986) in economies with sticky prices and contractually predetermined wages, and by Willman (1988) in the context of a model with endogenous output and foreign trade. Willman shows that crises are preceded by weak fundamentals such as economic recession and a current account deficit. A crucial feature of Willman's model is the existence of forward-looking wage contracts. Under perfect foresight, an anticipated future collapse will affect wages, which, in turn, will influence prices, the real exchange rate, and therefore, output and the trade balance. At the moment the collapse occurs, the real interest rate falls because of the jump in the rate of depreciation of the exchange rate. Output therefore increases, while the trade balance deteriorates. But since wage contracts are forward looking, anticipated future increases in prices are discounted back to the present and affect current wages. As a result, prices start adjusting before the collapse occurs. The steady rise in domestic prices is associated with appreciation in the real exchange rate and a negative impact on real output.

The continuous loss of competitiveness caused by the real appreciation, unless it is outweighed by effects from a fall in output, implies that the trade balance deteriorates in the period before the collapse of the fixed exchange rate regime.

1.2 Second generation models

Newer models, second generation research, are designed to capture features of the speculative attacks in Europe and in Mexico in the 1990s. Second generation models focus on potentially important nonlinearities in government behavior. They study what happens when government policy reacts to changes in private behavior or when the government faces an explicit trade-off between the fixed exchange rate policy and other objectives such as economic growth, low unemployment, or low inflation rate. Two examples of second generation research are introduced in this section.

1.2.1 Attack-conditional policy changes

Assume a conditional shift occurs in the growth rate of domestic credit from μ_0 to μ_1. If there is no attack on the fixed exchange rate, domestic credit grows at the rate μ_0; if there is an attack, domestic credit grows at the faster rate μ_1.

Figure 2 has the same shadow line as Figure 1, but it has an additional line representing the rate of credit expansion μ_1. The shadow rate line for $\mu = \mu_0$ intersects the

$\bar{s}$ line at point E and the shadow rate line for $\mu = \mu_1$, at point A. d^E and d^A indicate the domestic credit level at points E and A, respectively.

Suppose now that domestic credit lies in the range to the left of d^E. If there is no attack, the shadow rate is on the $\tilde{s}_{\mu 0}$ line. If speculators attack, the shadow rate moves to the $\tilde{s}_{\mu 1}$ line, which is still below the fixed exchange rate. Since any attack leads to capital losses, there is no incentive for the speculators to attack the fixed exchange rate if domestic credit is less than d^E.

If domestic credit is in the range between d^E and d^A, then multiple equilibria could be possible assuming speculators are small and uncoordinated as a group or face costs in confronting the government. The economy could reside on the lower shadow rate line indefinitely if agents believe it is impossible that the market will be attacked. On the other hand, the economy could jump to the higher shadow rate line if agents are confident there will be a run. Convinced of a run, no individual agent will find it profitable to hold domestic currency since this would result in a sure capital loss when the run occurs. Consequently, all agents will participate in an attack, leading to a collapse of the fixed rate and a more expansionary credit policy.

Suppose there is a large trader who can take a massive position against the fixed exchange rate, as George Soros supposedly did against the sterling in 1992. Then, there is a unique equilibrium. The economy faces only the attack equilibrium since a well-financed speculator always moves to exploit available profit opportunities. But suppose there is no large trader in the foreign exchange market, only many small credit-constrained traders. Without anything to coordinate their expectations and actions, they cannot mount an attack of sufficient size to move the economy from the no-attack equilibrium to the attack equilibrium. Then as suggested in Obstfeld (1986), there are multiple equilibria. The economy can maintain the fixed exchange rate indefinitely unless something coordinates expectations and actions to cause an attack.

Morris and Shin (1995) show how some types of uncertainty can eliminate multiple equilibria and make the attack outcome the unique equilibrium. They describe a speculative game in which each economic agent obtains information about the state of the economy, but with a small amount of error. Specifically, if the true state of the economy is $\bar{d}$, the agent observes a message that lies in the interval $[\bar{d} - \varepsilon, \bar{d} + \varepsilon]$, where ε is a small positive number. Messages are independent across agents. With noisy differential information, it is never common knowledge that the fixed exchange rate is sustainable. Accordingly, each investor should consider the full range of possible beliefs held by others and should think of what to do if the rate is unsustainable. If there is a good chance other speculators believe the fixed exchange rate is unsustainable, and if it is not too costly to take a position against the currency, then it makes sense for the individual investor to speculate, even knowing the peg is otherwise viable. Holding onto the currency may yield a bigger gain if everyone else holds on as well, but it is a riskier course of action because it relies on everyone else behaving similarly. Consequently, the only equilibrium in the region bounded between d^A and d^B is the attack equilibrium.

1.2.2 Escape clause

The second example comes from Obstfeld (1997). The model's basic framework is drawn from Kydland and Prescott (1977) and Barro and Gordon (1983). Here, a policymaker desires to raise employment above its natural rate through surprise currency depreciation. The model assumes an open economy and identifies the (log) nominal exchange rate, e_t (the price of foreign money in terms of domestic money) with the domestic price level. In this model, devaluations are triggered by the government's desire to offset negative output shocks, but a sudden shift in market expectations on the change in the exchange rate can trigger a devaluation that would not have occurred under different private expectations.

The government minimizes the loss function

$$L_t = (n_t - n^*)^2 + \theta(e_t - e_{t-1}), \qquad (15)$$

where n_t is employment, n^* is the government's employment target, $e_t - e_{t-1}$ is home inflation, and $\theta > 0$. Employment is determined by

$$n_t = n^* + \sqrt{a}[(e_t - E\{e_t | I_{t-1}\}) - u_t - k]. \qquad (16)$$

where I_{t-1} is the information set, u_t is an i.i.d mean zero employment shock, and $k>0$ is a fixed distortion in the economy that causes employment systematically to fall short of n^*, the target employment level. While labor markets pre-set wages in ignorance of the realized value of u_t, the policymaker is assumed to set the exchange rate after having observed the shock. In general, the policymaker will want to use the exchange rate to offset some of the effect of u_t on employment by unexpectedly depreciating the currency.

There are at least two distinct policymaking processes that might govern management of the exchange rate. Under discretion authorities choose e_t to minimize L_t given $E\{e_t|I_{t-1}\}$ and u_t. The exchange rate change a policymaker chooses under discretion is

$$e_t - e_{t-1} = \lambda\{E\{e_t|I_{t-1}\} - e_{t-1}) + \lambda(k + u_t), \quad \lambda \equiv \frac{\alpha}{\alpha+\theta} .$$

In addition, rational expectations in the labor market imply an expected loss of

$$EL^D = \gamma E(\frac{\lambda k}{1-\lambda} + k + u)^2, \quad \gamma \equiv (1-\lambda)\alpha .$$

Under the other rule, a fixed exchange rate: $e_t = e_{t-1}$, expected loss is

$$EL^F = \alpha E(k+u)^2 .$$

Assume the policymaker faces a personal cost $\underline{c}$ of revaluing the currency and a cost $\bar{c}$ of devaluing. Under discretion the policymaker takes the market's expected devaluation or revaluation rate, $\delta(\underline{u},\bar{u})$[6], as given. Then if the choice is to realign, the ex post social loss is

$$L^D\{\delta(\underline{u},\bar{u}),u\} = \gamma\{\delta(\underline{u},\bar{u}) + k + u\}^2 ,$$

and if the fixed exchange rate is maintained, the loss is

$$L^F\{\delta(\underline{u},\bar{u}),u\} = \alpha\{\delta(\underline{u},\bar{u}) + k + u\}^2 .$$

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[6] The expected exchange devaluation or revaluation rate when $u_t > \bar{u}$ or $u_t < \underline{u}$ where $\bar{u}$ and $\underline{u}$ are optimal policy switch points.

Without substantive loss of generality, assume that revaluation is ruled out from the start: only a large positive realization of u induces discretion, in which case devaluation occurs. The case of a single equilibrium boundary, $\bar{u}$, accurately depicts devaluation-prone countries while simplifying the algebra. In addition, to make matters simple the specific distribution assumed for u is the tent-shaped density function

$$g(u)=\begin{cases} (\mu-|u|)/\mu^2 & for\ u \in [-\mu, \mu] \\ 0 & for\ u \notin [-\mu, \mu] \end{cases}.$$

Since the policymaker's sole concern is the social-cost differential, $L^F - L^D$, her optimal decision rule is to devalue the currency for $u \geq \bar{u}$, where $\bar{u}$ is the solution to

$$L^F\{\delta(\bar{u}),\bar{u}\} - L^D\{\delta(\bar{u}),\bar{u}\} = (\alpha-\gamma)\{\delta(\bar{u})+k+\bar{u}\}^2 = \bar{c}.$$

More simply, interior equilibria correspond to values of $\bar{u}$ that solve

$$\delta(\bar{u})+k+\bar{u} \equiv \Phi(\bar{u}) = \sqrt{\bar{c}/(\alpha-\gamma)} \equiv K.$$

Alternative equilibria are most easily found by changing the shape of the function $\Phi(\bar{u})$ that emerges as the parameters k and λ - which respectively measure the severity of the time–inconsistency problem and the willingness to accommodate - are varied.

Consider an economy with a relatively large time-inconsistency problem k=0.015 and a non-extreme λ=0.75. The graph in Figure 3 shows the expected policy loss implied by different possible switch points $u \in [-0.03, 0.03]$ (right-hand vertical axis). The bold graph shows the $\Phi(\bar{u})$ function that arises in this case (left-hand vertical axis). The best equilibrium is at $\bar{u}^* = 0.0145$ (Figure 3), with loss $L(\bar{u}^*) = 0.867$, expected depreciation $\delta^* = 0.39\%$, and a 0.133 chance of devaluation. This equilibrium dominates a fixed rate because $L(0.03) = 1$. Imposing the fixed devaluation cost $K=\Phi(u^*)$ might not suffice to produce this relatively attractive equilibrium. There are two additional interior equilibria, associated with the boundaries $\bar{u}' = -0.0123$ and $\bar{u}'' = -0.0256$, and with the expected depreciation rates $\delta' = 3.0\%$ and $\delta'' = 4.4\%$, respectively. The implied losses, $L(\bar{u}') = 2.402$ and $L(\bar{u}'') = 3.291$, are much higher than that under a pure fixed-rate regime.

If there is a substantial risk of ending up at a bad equilibrium, then it might be best to go for an irrevocably fixed exchange rate. This could be achieved by confronting the policymaker with a prohibitively high devaluation cost on entering a common currency area. Uncertainty about the $\Phi(u)$ function would reinforce this conclusion, since a very small mistake in setting $\bar{c}$ could open the door to a catastrophe in the form of an additional, undesirable equilibrium. Unfortunately, imposing the appropriate cost on the policymaker is necessary, but not sufficient, for reaching a socially preferred equilibrium. Market expectations can be self-fulfilling, leading in plausible cases to any number of equilibria, most of which are dominated by the original simple rule.

1.2.3 Other objectives and early evidence

The two examples given above provide the essence of the second generation research, but cover only a part of the many mechanisms that have been discussed. Any economic objective that is reasonably part of the government's social welfare function and whose attainment involves a trade-off with a fixed exchange rate is a potential fundamentals for predicting crises. Alternative mechanisms introduced in this section mainly depend on the higher nominal interest rates associated with market expectation of depreciation.

This mechanism was especially evident in Italy's 1992 dilemma. Highly indebted governments with mostly short-term or floating-rate nominal debts will see their fiscal burden increase sharply if market expectations of depreciation drive up domestic interest rates. A government chooses depreciation of the domestic currency and/or an increase in the tax rate to repay its sharply increased debt. The government that wants to minimize the social loss function under this fiscal constraint will find that a certain degree of currency depreciation is optimal (Obstfeld, 1994). Therefore, *public debt* could be one of the objectives whose achievement includes a trade-off with the fixed exchange rate.

Banks

Many financial intermediaries come under pressures when market interest rates rise unexpectedly. Non-performing loans rise quickly, and depositors withdraw their funds out of concern over the safety of the banking system. The government's desire to sidestep a costly bailout at public expense results in a currency devaluation. (Tornell, 1999)

Real interest rates

The analysis in section 1.2.2 allowed for an effect of reduced competitiveness on output and employment[7], but not for real interest rate effects. With sticky domestic prices

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[7] An i.i.d. mean-zero shock, u, in equation (16) can be interpreted as changes in competitiveness.

and hikes in the nominal interest rate, the real interest rate rises and adversely affects employment. A slowdown of the economy may generate self-fulfilling devaluation pressures.

Pure speculation

Crises may simply occur as a consequence of pure speculation against the currency. For example, models of herding behavior emphasize that information costs may lead foreign investors to make decisions based on limited information and be more sensitive to rumors (Calvo and Mendoza, 1997). Contagion effects suggest that a crisis in one country may raise the odds of a crisis elsewhere by signaling that devaluation is more likely as a result of the initial crisis. This signal may lead to a self-fulfilling speculative attack (Masson, 1998).

2. Empirical literature

2.1 Nonstructural empirical analyses

Nonstructural approaches for analyzing crisis episodes in a set of countries before the 1990s confirm the role of traditional fundamentals in predicting crises. For example, Klein and Marion (1997) used panel data for 80 devaluation episodes in Latin American countries during 1957-91 and found the monthly probability of abandoning a pegged exchange rate increased with real overvaluation and declined with the level of foreign assets. Structural factors, such as the openness of the economy and its geographical trade concentration, political variables, such as changes in the executive, and time already spent on the peg also influenced the monthly probability of ending a fixed exchange rate.

The speculative attacks of the 1990s, particularly those in Europe, challenged the view that currency crises were due mainly to the government's inconsistent policy to achieve fiscal and monetary discipline. Numerous researchers, motivated by the crises in the 1990s and the theoretical achievements by the second generation models, have turned to empirical studies that use an extensive variety of informative variables to distinguish between periods leading up to currency crises and tranquil periods.

These studies differ from one another in the definition of a crisis, estimation methodology, explanatory variables included, and data coverage. Three examples are provided below.

2.1.1 Examples

Example 1

Frankel and Rose (1996) apply a probit model to analyze the determinants of currency crises using a panel of annual data for 105 developing countries, from 1971 through 1992. Their work is non-structural, and takes the form of univariate graphical analysis and multivariate statistical analysis. They define a currency crisis as a nominal depreciation of at least 25 percent in the bilateral exchange rate against the US dollar with respect to the previous year.

They analyze variables relevant to the theoretical literature. Informed by the first generation models, they choose the rate of growth of domestic credit, the government budget as a fraction of GDP, the ratio of reserves to imports, the degree of overvaluation of the real exchange rate, differentials between domestic and foreign interest rates, and the current account as a percentage of GDP. In the spirit of the second generation models, they also examine the effects of the growth rate of real output, the ratio of short-term debt to total external debt and the ratio of external debt to GDP.

Most estimation results from their probit regression are consistent with the theoretical literature. They find that a high short-term debt ratio (as a share of total external debt), low international reserves, high domestic credit growth and overvaluation of the real exchange rate increase the probability of currency crises.

The significance of their results, however, is limited by the lack of robustness to various sensitivity tests and poor performance in predicting actual crises.[8]

Example 2

Sachs, Tornell and Velasco (1996) analyze the Mexican crisis and its reverberations in the financial markets of 20 developing countries in 1995. They attempt to determine whether there exists some set of fundamentals that help explain the variation in financial crises across countries or whether the variation just reflects contagion.

They measure the extent of financial crisis in 1995 with a crisis index (denoted IND) that measures pressures on the foreign exchange market. IND is a weighted average of the devaluation rate with respect to the U.S. dollar and the percentage change in foreign exchange reserves between the end of November 1994 and the end of the first six months of 1995. The weights given to the loss in reserves and the devaluation are designed to equalize volatilities of the components of the index.

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[8] Only five of sixty nine actual crises during the sample period are predicted by their model.

The rationale for this index is as follows. If capital inflows reverse, the government can let the exchange rate depreciate. Alternatively, it can defend the currency by running down reserves or by increasing interest rates. Since there are no reliable and comparable cross-country interest rate data, they construct the index using levels of reserves and exchange rates.

Using the crisis index, IND, as a dependent variable, they use an ordinary least square regression and find that low international reserves relative to broad money, real exchange rate appreciation, and a weak banking system explain about 70 percent of the variation of their IND. Cooper (1996), however, has warned against overemphasizing these results, since they are drawn from a single crisis episode.

Example 3

Kaminsky and Reinhart (1999) examine the links between banking and currency crises and the stylized facts of the period leading up to and immediately following crises. Initially they provide a definition and chronology of the crises and their links. Then, they review the stylized facts around the periods surrounding the crises, while the next step addresses the issues of the vulnerability of economics around the time of the crisis and the issue of predictability.

The definition of a crisis is a mix of those in Frankel and Rose (1996) and Sachs, Tornell and Velasco (1996). They construct an index of currency crisis

$$I = \frac{\Delta_e}{e} - \frac{\sigma_e}{\sigma_R} \cdot \frac{\Delta_R}{R}$$

where Δ_e/e is the rate of change in the exchange rate, Δ_R/R is that of reserves, σ_e is the standard deviation of Δ_e/e, and σ_R is the standard deviation of Δ_R/R. Since changes in the exchange rate enter with a positive weight and changes in reserves have a negative weight attached, readings of this index that were three standard deviations or more above the mean were cataloged as crises.

Using monthly data of 16 macroeconomic and financial variables around the time of crises, they find evidence suggesting that several economic variables behave quite differently in tranquil periods as compared to crises period through graphical analysis: The real exchange rate is overvalued; exports and the terms of trade are deteriorated; and economic growth is slowed down. The periods just before a crisis are also characterized by a highly expansionary monetary policy, low reserves and increases in interest rate differentials. Based on the results, they propose an early warning system to prevent future currency crises. However, a graphical approach has disadvantages. The graphs are informal. More importantly, they are intrinsically univariate.

2.2 Structural empirical analyses

Since Krugman (1979), currency crises have been thought to have a significant predictable component, with first generation models identifying fundamentals useful for prediction. A fiscal deficit financed by domestic credit creation is considered to be the root cause of a speculative attack. As the monetary authority monetizes the budget deficit, oversupply of money causes a gradual decline in international reserves. Accordingly, investors attack the fixed exchange rate, depleting the government's reserve holdings used for a defense.

Following empirical studies along this line have focused on Mexico between 1973 and 1982 (Blanco and Garber, 1986), Argentina (Cumby and Van Wijnbergen, 1989), and Mexico in the 1980s (Goldberg, 1994).

The currency crises in Europe and Mexico in the early 1990s, however, rejected the major role of traditional factors in crises. Otker and Pazarbasioglu (1996) computed the probability of an exchange rate regime change using Blanco and Garber's model. They found the Mexican financial crisis in 1994 was not the result of fiscal imbalances, which had played a major role in Mexico's previous balance of payments crises; rather, it was the rise in private sector indebtedness and the corresponding increase in credit to the banking system that augmented the pressures building up in the exchange market in mid 1994. Moreover, the experiences of several European countries in the context of the European Monetary System (Otker and Pazarbasioglu, 1997b) show that there are other triggering determinants of crises - e.g. pure speculation - unexplained by the fundamentals. As these studies focus on a particular country in a specific time period, they show the uniqueness of each country's currency crisis.

2.2.1 Examples

Example 1

Blanco and Garber (1986), in their study of repeated devaluations of the Mexican peso, present the first empirical study of the speculative attack model.

A money market equation provides the essential part of their model:

$$m_t - p_t = \beta + \Omega y_t - \alpha i_t + \omega_t \qquad (1)$$

where m_t, p_t and y_t are the logarithms of the money stock, the domestic price level and the aggregate output level, respectively; i_t is the domestic interest rate; and ω_t is a stochastic disturbance to the money demand. Equation (1) represents the demand for real money balances. They further assume that the price level and the interest rate are determined by

$$i_t = i_t^* + Es_{t+1} - s_t \qquad (2)$$

$$p_t = p_t^* + s_t + u_t \qquad (3)$$

where an asterisk signifies an exogenous foreign variable, and s_t and u_t are the logarithms of the nominal and the real exchange rate, respectively. The operator E represents expectations conditional on information through time t. Equation (2) reflects the assumption of uncovered interest rate parity. Equation (3) comes from the definition of the real exchange rate.

Using the money market clearing condition, the flexible exchange rate can be determined. Substituting (2) and (3) into (1) yields

$$h_t = -\alpha E\tilde{s}_{t+1} + (1+\alpha)\tilde{s}_t \qquad (4)$$

where $h_t \equiv \log[D_t + \overline{R}] - \beta - \Omega y_t + \alpha i_t^* - p_t^* - u_t - \omega_t$, D_t is the domestic credit component of the monetary base at time t and $\tilde{s}_t$ represents the permanently floating exchange rate. $\overline{R}$ represents a lower bound on net reserves where the central bank, having fixed the exchange rate, stops intervening in the foreign exchange market. Equation (4) says that the floating exchange rates $\tilde{s}_t$ and $E\tilde{s}_{t+1}$ are determined by the economic fundamentals h_t.

h_t is assumed to be a first order autoregressive process exogenous to the exchange rate. Specifically, the process h_t is

$$h_t = \theta_1 + \theta_2 h_{t-1} + v_t \qquad (5)$$

where v_t is a white noise process with a normal density function $g(v)$, with zero mean and standard deviation σ.

The flexible exchange rate is obtained by solving the difference equations in (4) and (5). The solution is

$$\tilde{s}_t = \mu\alpha\theta_1 + \mu h_t \qquad (6)$$

where $\mu = 1/[(1+\alpha) - \alpha\theta_2]$.

They assume that the new fixed rate is a simple linear function

$$\hat{s}_t = \tilde{s}_t + \delta v_t \qquad (7)$$

where δ is a nonnegative parameter, and $\hat{s}_t$ is the new fixed exchange rate that will be established if the level of reserves attains the value $\overline{R}$ at time t.

$\hat{s}_t$'s exceeding the current fixed rate is equivalent to a devaluation at time t. Therefore, the probability of devaluation at time $t+1$ based on information available at t is

$$pr(\mu r(_1 + \mu h_{t+1} + \delta v_{t+1} > \bar{s}),$$

where $\bar{s}$ is the time t value of the fixed rate. Alternatively, the devaluation probability is

$$1\text{-}F(k_t) \equiv pr(v_{t+1} > k_t) \qquad (8)$$

where $k_t \equiv [1/(\mu+\delta)][\bar{s} - \mu\alpha\theta_1 - \mu(\theta_1 + \theta_2 h_t)]$, and $F(k_t)$ is the cumulative distribution function associated with $g(v)$.

Knowing this density function, agents can form expectations of future exchange rates from the average of the current fixed exchange rate and the expected rate conditional on a devaluation, both weighted by the respective probabilities of occurrence:

$$Es_{t+1} = F(k_t)\bar{s} + [1 - F(k_t)]E(\hat{s}_{t+1}|v_{t+1} > k_t). \qquad (9)$$

Using (7), the conditional expectation can be expressed as

$$E(\hat{s}_{t+1}|v_{t+1} > k_t) = \mu\theta_1(1+\alpha) + \mu\theta_2 h_t + (\mu+\delta)E(v_{t+1}|v_{t+1} > k_t) \quad (10)$$

where $E(v_{t+1}|v_{t+1} > k_t) = \int_{k_t}^{\infty} \frac{vg(v)}{1 - F(k_t)} dv$. Since g($v$) is a normal density function, the unconditional forecast of the exchange rate for $t+1$ is

$$Es_{t+1} = F(k_t)\bar{s} + [1 - F(k_t)][\mu]_1(1+\alpha) + \mu\theta_2 h_t] + \frac{\sigma(\mu+\delta)\exp[-.5(k_t/\sigma\sigma^2]}{\sqrt{2\pi}} \quad (11)$$

The one-step-ahead devaluation probability (8) and the conditional and unconditional exchange rate forecasts (10) and (11) are the main products of this model. It is expected that [1- $F(k_t)$] should reach a peak immediately before a devaluation and Es_{t+1} should be closely correlated with the appropriate forward rates. Finally, the conditional forecast should approximate the exchange rate when devaluation occurs.

The Mexican crisis over the 1973-1982 is analyzed in this study. Their estimated probabilities of devaluation in the next quarter, which range from highs of more than 20 percent in late 1976 and late 1981, to lows of less than 5 percent in early 1974 and late 1977, reach local peaks in the period of devaluation and reach local minima in the periods following devaluation as predicted by the theory. Furthermore, the expected exchange rates conditional on devaluation are close to the values that actually materialized in the major episodes.

Example 2

Goldberg (1994) applied a discrete time model of a collapsing exchange rate regime to the experience of Mexico between 1980 and 1986. The model was used to predict the probability that the existing fixed exchange rate regime would collapse due to a speculative attack on central bank foreign exchange reserves.

She applied almost the same framework as Blanco and Garber (1986). The model is provided by equations (1)-(7) below:

$$\frac{M_t^d}{Q_t} = a_0 - a_1 I_t + a_2 Y_t - a_3\left(\frac{Es_{t+1} - s_t}{s_t}\right) \quad (1)$$

$$Q_t = \alpha P_t + (1-\alpha) s_t P_t^* \quad (2)$$

$$\frac{P_t}{s_t} = P_t^* + \rho_t + \eta_t \quad (3)$$

$$I_t = I_t^* + \frac{Es_{t+1} - s_t}{s_t} \quad (4)$$

$$M_t^s = D_t + R_t \quad (5)$$

$$D_t = D_{t-1} + \mu_t + \varepsilon_t \quad (6)$$

$$\varepsilon_t = \gamma_t - \varphi_t \quad (7)$$

While the variables in the equations do not take logarithms, equation (1) reflects real money demand and equation (2) defines the aggregate price index as the weighted sum of domestic goods prices P_t and traded goods prices $s_t P_t^*$. The weight α corresponds to the share of domestic goods in consumer expenditure.

By equation (3) the deviation of domestic good's price from the foreign good's due to medium-term systematic deviations from PPP, denoted by ρ_t, and due to stochastic shocks to relative prices, denoted by η_t, could be modeled and equation (4) shows the uncovered interest rate parity.

Equation (5) is the money supply equation. R_t represents foreign exchange reserve and D_t is total domestic credit.

The domestic credit component of the money supply is modeled in equation (6) as evolving according to a trend that reflects the mean basic government budget deficit, μ_t, with some period-by-period stochastic component ε_t.

She added a currency substitution impetus, $(Es_{t+1} - s_t)/s_t$, associated with an expected devaluation of the nominal exchange rate to the demand for real balances. Also, in equation (7), she decomposed the shock to domestic credit expansion by source: (i) random revenue or expenditure affecting the need to monetize government deficits, γ_t; and (ii) random and constrained access to external credit, φ_t, that makes uncertain the share of government deficits to be financed by external borrowing instead of inflationary finance.

Using the money market clearing condition, the period t+1 shadow exchange rate is derived as

$$\tilde{s}_{t+1} = \left[\frac{\left(\frac{a_{t+1} + a_1 + a_3}{a_{t+1}} \right) \mu_{t+1} + D_t + \overline{R} + \gamma_{t+1} - \varphi_{t+1}}{a_{t+1} \{ P^*_{t+1} + \alpha_{t+1} (\eta_{t+1} + \rho_{t+1}) \}} \right] \qquad (6)$$

where $a_{t+1} \equiv a_0 + a_2 Y_{t+1} - a_1 I^*_{t+1}$.

Compared to Blanco and Garber's model, her framework explicitly shows the effects of the fundamental variables and parameters on the shadow exchange rate.

She used an ARIMA process to describe the evolution of each variable for which forecasts are required and instrument variables to avoid simultaneity problems for the estimation of real money demand. Using an iterative estimation procedure, she re-estimated estimated parameters to yield new parameter values for estimation of the next pass estimates of collapse probabilities and expected shadow exchange rates. The iterative estimation procedure is completed until parameter convergence occurs.

By applying her model to the Mexican currency crisis, she found that Mexico's monetary and fiscal policies, rather than anticipated external credit shocks, were the driving forces in triggering speculative attacks on the Mexican peso in the 1980s.

Example 3

Otker and Pazarbasioglu (1997b) evaluated the role of macroeconomic fundamentals in generating episodes of speculative pressures on six currencies of the European Exchange Rate Mechanism (ERM) in 1992 and 1993.

The study proceeds in two steps. First, it identifies whether the observed regime changes can be predicted by the presence of speculative pressures. Second, in order to identity the contribution of deterioration in economic fundamentals to such pressures, it estimates the probability of a regime change as a function of such fundamentals by using a monetary model of speculative attacks. The latter outlines a process in which fiscal or financial imbalances may lead to an eventual collapse of the exchange rate peg by generating domestic credit expansions that initially cause a gradual erosion of the foreign exchange reserves. The erosion of reserves is followed by generally self-fulfilling currency attacks as forward-looking investors engage in one-sided bets, anticipating that the central bank will exhaust its reserves in defending its currency. Eventually, the peg can no longer be sustained and the prevailing exchange rate peg collapses, involving either a discrete devaluation or a switch to flexible rates.

In order to identify the episodes of speculative pressures and the associated regime changes, they estimate the one-step-ahead probability of a regime change as a function of pressure indicators below

$$\pi_t = \Pr ob(Y-1) = \pi[(i_t - i_t^*),(E-C)_t, \log R_t, \Delta \log R_t] \qquad (1)$$

where π_t denote the one-step-ahead probability of a regime change, Y is the central bank's decision regarding a change in its exchange rate regime as a discrete variable which can take only two values; one, if there is either a devaluation or a switch to flexible rates, and zero, when, existing parity is maintained, i and i^* are short-term interest rates in the domestic and anchor country, $(E - C)$ is the deviation of the spot rate from the central parity, R is official foreign reserves, and Δ is the first difference operator. For Belgium, Denmark, and Italy, a loss of foreign reserves and increased depreciation of the currency within the band appear to indicate a build-up of speculative pressures, while for France, Ireland and Spain the existence of pressures appear to be mainly associated with the depreciation of the spot rate within the band and hikes in domestic interest rates.

They also studied the speculative pressure by the monetary model. If we define $\widetilde{s}_{t+1}$ to be a shadow exchange rate at time $t+1$, the one-step-ahead probability, π_t, of a regime change at $t+1$ based on information available at t can then be written as a function of the prevailing fixed rate and a set of economic fundamentals, h_t, that influence the shadow exchange rate:

$$\pi_t \equiv \Pr(\widetilde{s}_{t+1} \rangle \bar{s}_t) \equiv \pi(h_t;\bar{s}_t), \quad h_t = h(D_t, y_t, u_t, i_t^*, p_t^*) \qquad (2)$$

where D_t is the central bank's domestic credit, y_t is real output, i_t^* and p_t^* are the short-term interest rate and price level of the anchor country, respectively, and u_t is the real exchange rate. For the French franc, the expansion of central bank credit appears to have contributed to pressure. In addition, the positive coefficient on the unemployment rate for France and Italy and of the loss of competitiveness for France and Ireland are consistent with explanations that adverse economic conditions can make it costly for the government to defend the fixed exchange rate. This market perception may set off speculative pressures and result in an adjustment in the exchange rate.

3. Extended currency crisis model

To evaluate the influence of macroeconomic fundamentals on the Asian currencies without the contagion effects, this section introduces basic and extended currency crisis models using the implications of speculative attack model, first suggested by Krugman (1979) and formalized by Flood and Garber (1984a).[9]

3.1 Basic model

A number of small country assumptions are used in setting up the model. The country described by the model is a small developing open economy. The foreign price level is taken as an exogenous contributor to the randomness in the purchasing power relationship. The country also lacks well-developed financial markets. Therefore, its government cannot engage in open market operations through bond sales. Throughout this section, the transition or 'collapse' studied is one in which a fixed exchange rate gives way to a flexible rate or a developed new fixed rate. This model applies *M2* instead of *M1* as a money supply to support the particular aspects of the Asian currency crisis in which the domestic credit by private banks reflected in *M2* played a crucial role. In particular, previous works used *M1* as a money supply to represent the domestic credit of central bank. However, considering that *M2* is an account in the debit of monetary survey which shows integrated accounts of the central bank and domestic banks and that domestic credits of central bank and domestic banks are accounts in the asset of monetary

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[9] The survey article by Agenor, Bhandari, and Flood (1992) provides a detailed description of the currency crises models. Blanco and Garber (1986), Cumby and van Wijnbergen (1989), Goldberg (1994), Otker and Pazarbasioglu (1996, 1997a, 1997b) find empirical support for the basic currency crisis model.

survey as well, we can regard *M2* as an monetary aggregate which explains the domestic credits of central bank and domestic banks at the same time. Therefore, the use of *M2* as a monetary aggregate should contribute to the explanation of how a "Lending Boom" by domestic banks became one of the major causes in the Asian currency crisis. In addition, the model allows currency substitution impetus associated with an expected devaluation of the nominal exchange rate and a risk premium.

The following equations describe the basic model.

$$m_t^d - p_t = a_0 - a_1 i_t + a_2 y_t - a_3(E_t s_{t+1} - s_t + \rho_t) \quad (1)$$

$$i_t = i_t^* + (E_t s_{t+1} - s_t) + \rho_t \quad (2)$$

$$p_t = p_t^* + s_t + u_t \quad (3)$$

$$m_t^s = d_t + r_t \quad (4)$$

$$m_t^d = m_t^s \quad (5)$$

$$d_t = d_{t-1} + E_{t-1}\mu_t^d + \varepsilon_t^d \quad (6)$$

where m, d, r, p, p^*,and y are the logarithms of the money stock, domestic credit extended by the domestic banks, central bank foreign reserves, domestic price level, foreign price level, and real output, respectively, i is the domestic interest rate, i^* is the foreign nominal interest rate, ρ is the risk premium on domestic assets, s and u are the logarithms of the nominal and real exchange rates, respectively. E_t represents the expectation conditional on information available in the current period.

Equation (1) specifies the transaction and asset motives for real money balances. In addition to the standard variables, real money balances are reduced by the opportunity for currency substitution. The impact of currency substitution on real money balances is proportional to the sum of expected rate of domestic currency depreciation, $E_t s_{t+1} - s_t$, and a risk premium, ρ. An increase in $E_t s_{t+1} - s_t + \rho_t$ reflects a decrease in the desirability of holding the domestic currency.

Equation (2) is the interest parity condition, which states that the interest rate differential between the domestic and foreign country reflects the expected rate of depreciation of the domestic currency and a risk premium on domestic assets. The risk premium for domestic investments reflects the standard increased compensation for more risky investments in domestic assets.

Equation (3) allows for deviations from purchasing power parity.

Equation (4) defines the money supply as the sum of logarithm of domestic credit extended by the domestic banks and logarithm of central bank foreign reserve. The currency crises in the 1970s, 1980s, and early 1990s were rooted in the dynamics of the domestic credit extended by the central bank to the government. When there is an excess of domestic credit creation, a new money market equilibrium can be achieved by a reduction in the central bank's foreign exchange reserves or by an exchange rate adjustment. However, the crucial role of the domestic credit to the government in the crisis vanished in the Asian currency crisis in 1997. Instead, the domestic credit to the private sector enhanced by the domestic banks fueled by the foreign liabilities took its role.

Equation (5) is the money market equilibrium condition. The money market equilibrium condition determines the path of foreign reserves of the central bank under a fixed exchange rate system. When the reserves needed to maintain this equilibrium are exhausted, or when they reach a critical level, R_c, the exchange rate must be adjusted.

Equation (6) assumes that d evolves according to a period-by-period systematic stationary component, $E_{t-1}\mu_t^d$, and a stochastic element ε_t^d. In addition, p^*, y, i, i^*, and u are assumed to evolve by following the same process as d.

The probability at the end of period t that the fixed exchange rate regime will be abandoned at the end of $t+1$ is denoted as π_t. Therefore, the probability that the fixed exchange rate regime will continue is $(1-\pi_t)$. This implies that the expected exchange rate at $t+1$ is

$$E_t s_{t+1} = \pi_t E_t \widetilde{s}_{t+1} + (1-\pi_t)s_t \qquad (7)$$

where $\widetilde{s}_{t+1}$ is the shadow, or floating, exchange rate that would clear the market when the central bank stops defending its fixed parity.

By combining equations, (1), (2), (3), (5), and (7) and assuming that the fixed exchange rate regime collapses or $\pi_t = 1$,

$$m_t = a_0 - a_1 i_t^* + a_2 y_t + p_t^* + s_t + u_t - (a_1 + a_3)(E_t \widetilde{s}_{t+1} - s_t + \rho_t). \qquad (8)$$

By following Flood and Garber (1984)'s method of undetermined coefficients positing without the possibility of a bubble path, the shadow exchange rate expressed as

$$\widetilde{s}_t = \lambda_0 + \lambda_1 m_t \; . \qquad (9)$$

Then, by using equation (6) to depict the growth rate of the domestic money supply, the derived coefficients are

$$\lambda_0 = -a_0 + a_1 i_t^* - a_2 y_t - p_t^* - u_t + (a_1 + a_3)(E_{t-1}\mu_t^d + \rho_t), \text{ and}$$

$$\lambda_1 = 1$$

Therefore, by substituting λ_0 and λ_1 into equation (9), the shadow exchange rate is derived as

$$\widetilde{s}_t = m_t - a_0 + a_1 i_t^* - a_2 y_t - p_t^* - u_t + (a_1 + a_3)(E_{t-1}\mu_t^d + \rho_t). \quad (10)$$

We can use equation (10) to derive the probability of a collapse, π_t, occurring at the end of period $(t+1)$.[10]

3.2 Extended model

As indicated by recent literature, virtually all of the variables in the monetary model can be expected to be non-stationary. Hence individual economic variables may wander extensively when shocked. However, previous studies did not explicitly take into account the non-stationarity of the variables. For example, in equation (1), the coefficient of the each variable should be estimated first in order to obtain the probability of collapse. However, OLS estimates of these coefficients will display the spurious regression problem and the conventional t-ratio and F significance tests cannot be applied. Therefore, an extension to the basic model is necessary.

As far as the demand for money is concerned as in equation (1), economic theory suggests that the non-stationary variables in the function are expected to move so that they do not drift too far apart in the long run. The long run equilibrium can be interpreted with the concept of cointegration in econometric literature (Engle and Granger, 1987). If

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[10] A process of derivation of π_t is explained in section 4.

each element of a vector series X_t becomes stationary after first-differencing, but there exists a linear combination $\alpha' X_t$ that already is stationary, then the X_t are said to be cointegrated with a cointegrating vector α. By interpreting $\alpha' X_t$ as reflecting the long-run equilibrium, cointegration implies that deviations from the long run equilibrium are stationary, with a finite variance. This is so even though the series, themselves, are non-stationary and have infinite variance.

Although economic theory suggested[11], there is no prior reason to believe that the I(1) variables observed in this study necessarily obey the functional form in equation (1),

$$m_t^d - p_t = a_0 - a_1 i_t + a_2 y_t - a_3 (E_t s_{t+1} - s_t + \rho_t).$$

Hence to avoid an invalid restriction of the real money demand function, the likelihood ratio tests of cointegration rank of Johansen (1988b, 1991), Johansen and Juselius (1990), and a common stochastic trends test of Stock and Watson (1988) have been performed in the previous studies. Since Johansen's (1988b, 1991) maximum likelihood methods for the analysis of cointegration can simultaneously detect the number of the cointegration rank in the system, estimate, and test for linear hypotheses about the cointegrating vectors and their adjustment coefficients, this is the most favored technique in recent research for the reliable form of real money demand function.

To test for the number of cointegration relations and estimate the cointegrating vectors require that one begins with a VAR(p) representation expressed in first order difference and lagged levels,

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[11] Refer to Chapter V for the theoretical background of real money demand function.

$$H_1{:}\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-1} + \mu_0 + \eta_t \quad (11)$$

where x_t is a p-dimensional vector of I(1) variables, $\eta_1, \cdots, \eta_T$ are $IIN_p(0, \Lambda)$ and $x_{-k+1} \cdots x_0$ are fixed. The Π matrix conveys the long-run information in the data. The hypothesis of r cointegrating vectors is formulated as a reduced rank of the Π matrix,

$$H_2(r): \Pi = \alpha\beta' \qquad (12)$$

where α and β are p x r matrices of full rank.

Under $H_2(r): \Pi = \alpha\beta'$, (11) can be interpreted as an error correction model (see Engel and Granger 1987, and Johansen 1988a). Therefore, equation (1) in the previous subsection can be modified as an error correction model (ECM),

$$\Delta(m_t - p_t) = \Delta rm_t = \mu_0 + \alpha(L)\Delta i_{t-1} + \beta(L)\Delta y_{t-1} + \gamma(L)\Delta if_{t-1} + \Pi_1 x_{t-1} + \eta_t \qquad (13)$$

where $if_t = i_t - i_t^* = E_{t-1}\mu_t^d + \rho_t$, Π_1 is the first row of the Π, and where $x'_{t-1} = (rm_{t-1}, a_0, t, i_{t-1}, y_{t-1}, if_{t-1})$. Then, the shadow exchange rate,

$$\tilde{s}_t = m_t - \mu_0 - \alpha(L)\Delta i_{t-1} - \beta(L)\Delta y_{t-1} - \gamma(L)\Delta if_{t-1} - \Pi_1 x_{t-1} - rm_{t-1} - p_t^* - u_t \qquad (14),$$

can be explained as driven by the long-run disequilibrium shocks, short-run shocks, and random innovative shocks respectively using the error correction model (ECM).

4. The probability of currency crisis

4.1 Explicit form of probability in the basic model

The probability of a currency crisis, π_t, is the probability that the shadow exchange rate, $\tilde{s}_t$, will exceed $\bar{s}_t$, the time t value of the fixed rate, in period *(t+1)*. It is therefore defined by:

$$\pi_t = \Pr[\tilde{s}_{t+1} - \bar{s}_t \rangle 0]$$

$$= \Pr[m_{t+1} - a_0 + a_1 i^*_{t+1} - a_2 y_{t+1} - p^*_{t+1} - u_{t+1} + (a_1 + a_3)(E_t \mu^d_{t+1} + \rho_{t+1}) - \bar{s}_t \rangle 0]. \quad (15)$$

The rationale for this formulation is straightforward. Since a government's commitment to a fixed rate gives speculators unrestricted access to central bank foreign reserves, speculators who perceive that the shadow rate will exceed the fixed rate will purchase reserves at the fixed exchange rate. With the opportunity to resell the reserves at the higher market rate (equal to the shadow rate), their speculative purchases would yield a profit of $(\tilde{s}_{t+1} - \bar{s}_t) \rangle 0$ per unit of reserves. Barring interim intervention using capital controls and trade restrictions, which would alter the access to and the speed of decline of reserves, the speculation may draw the central bank reserves down to their critical minimum level. In this way, the prediction of a currency crisis can become self-fulfilling. A forced collapse would not occur at shadow exchange rates below the fixed rate since there would be no opportunity for profit on the purchase of foreign exchange reserves.

Since the forced currency crisis can only occur when a speculative attack is capable of driving reserves at or below their minimum or critically low level, (15) can be rewritten with m_{t+1} replaced by $d_{t+1} + r_c$.

$$\pi_t = \Pr[d_{t+1} + r_c - a_0 + a_1 i^*_{t+1} - a_2 y_{t+1} - p^*_{t+1} - u_{t+1}$$

$$+ (a_1 + a_3)(E_t \mu_{t+1} + \rho_{t+1}) - \bar{s}_t \rangle 0] \quad (16)$$

The probability of a crisis in period *(t+1)*, viewed from period *t* given the available information set, is composed of both random influences and components known with certainty in period *t*. The random influences come about through period-by-period stochastic components of *d*, p^*, *y*, *i*, i^*, and *u*.

Before continuing with the derivation of the probability of collapse, it is necessary to attach statistical distributions to the random components of d produced by the money supply and the random components of p^*, y, i, i^*, and u caused by fluctuations in money demand. The assumption that only d and u have random influences, that is other money demand's randomness is neglected as Goldberg (1994), makes the model excessively simplified and unrealistic in the rigorous aspects of time series analysis. In addition, stochastic properties of variables should not be discarded in making forecasts of variables needed to derive the probability of collapse. To do this, the stochastic parts, $\varepsilon_t^d, \varepsilon_t^i, \varepsilon_t^{i^*}, \varepsilon_t^y, \varepsilon_t^{p^*}$ and ε_t^u are assumed to be uncorrelated with each other and their linear combination is assumed to be normally distributed, $\varepsilon_t|\Omega_{t-1} \sim N(0, \sigma_t^2)$ where

$$\varepsilon_t = \varepsilon_t^d + (a_1 + a_3)\varepsilon_t^i - a_3\varepsilon_t^{i^*} - a_2\varepsilon_t^y - \varepsilon_t^{p^*} - \varepsilon_t^u \quad \text{and}$$

$$\sigma_t^2 = (\sigma_t^d)^2 + (a_1 + a_3)^2(\sigma_t^i)^2 + a_3^2(\sigma_t^{i^*})^2 + a_2^2(\sigma_t^y)^2 + (\sigma_t^{p^*})^2 + (\sigma_t^u)^2$$ [12].

The normal distribution is chosen for analytical convenience. The systematic stationary components of the variables are summed into $E_{t-1}\mu_t$ where

$$E_{t-1}\mu_t = E_{t-1}\mu_t^d + (a_1 + a_3)E_{t-1}\mu_t^i - a_3E_{t-1}\mu_t^{i^*} - a_2E_{t-1}\mu_t^y - E_{t-1}\mu_t^{p^*} - E_{t-1}\mu_t^u.$$

Then, the probability of devaluation at time $t+1$ based on information available at t is

$$\pi_t = \Pr[d_t + r_c - a_0 + (a_1 + a_3)i_t - a_3i_t^* - a_2y_t - p_t^* - u_t + E_t\mu_{t+1} + \varepsilon_{t+1} - \bar{s}_t \rangle 0]. \quad (17)$$

For ease of notation, define a variable k_t as

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[12] The covariance terms are removed since $\varepsilon_t^d, \varepsilon_t^i, \varepsilon_t^{i^*}, \varepsilon_t^y, \varepsilon_t^{p^*}$ and ε_t^u are assumed to be uncorrelated with each other.

$$k_t = \bar{s}_t - d_t - r_c + a_0 - (a_1 + a_3)i_t + a_3 i_t^* + a_2 y_t + p_t^* + u_t - E_t \mu_{t+1}.$$

Subsequently, the π_t can be rearranged as

$$\pi_t = \Pr[\varepsilon_{t+1} \rangle k_t] = \int_{k_t}^{\infty} \frac{1}{E_t \sigma_{t+1} \sqrt{2\pi}} e^{-\frac{\varepsilon_{t+1}^2}{2E_t \sigma_{t+1}^2}} d\varepsilon. \quad (18)$$

where $E_t \sigma_{t+1}^2$ is a forecasted volatility of ε_{t+1}.

The probability of devaluation in equation (17) is expected to be positively influenced by the previous level of domestic credit and the interest rate. As the foreign interest rate, GDP, foreign price, and real exchange rate's past levels rise, the probability of devaluation is predicted to decline. Alternatively, an increase in the floor level of reserves, r_c, and systematic stationary components of domestic credit and the interest rate increase the probability of a devaluation.

4.2 Explicit form of probability in the extended model

The probability of a currency crisis, π_t, in the extended model has the same validation as the basic model. Therefore, the probability is

$$\pi_t = \Pr[\tilde{s}_{t+1} - \bar{s}_t \rangle 0]$$

$$= \Pr[d_{t+1} + r_c - u_0 - \alpha(L)\Delta i_t - \beta(L)\Delta y_t - \gamma(L)\Delta if_t - \Pi_1 x_t - rm_t - p_{t+1}^* - u_{t+1} - \bar{s}_t \rangle 0]. (19)$$

As before, the randomness comes about through period-by-period stochastic components of *d*, *p*•, and *u*.

Suppose the systematic stationary component of each variable and the stochastic part of the processes *d*, *p*•, and *u* are justified in the same way as was the case in the basic model. Then, the probability of devaluation at time $t+1$ based on information available at t is,

$$\pi_t = \Pr[d_t + r_c - u_0 - \alpha(L)\Delta i_t - \beta(L)\Delta y_t - \gamma(L)\Delta if_t - \Pi_1 x_t - rm_t - p_t^* - u_t$$

$$+ E_t \mu_{t+1} + \varepsilon_{t+1} - \bar{s}_t \rangle 0] \qquad (20)$$

Once more, if we define k_t as

$$k_t = \bar{s}_t - d_t - r_c + u_0 + \alpha(L)\Delta i_t + \beta(L)\Delta y_t + \gamma(L)\Delta if_t + \Pi_1 x_t + rm_t + p_t^* + u_t - E_t \mu_{t+1},$$

then π_t can be rearranged as,

$$\pi_t = \Pr[\varepsilon_{t+1} \rangle k_t] = \int_{k_t}^{\infty} \frac{1}{E_t \sigma_{t+1} \sqrt{2\pi}} e^{-\frac{\varepsilon_{t+1}^2}{2E_t \sigma_{t+1}^2}} d\varepsilon . \qquad (21)$$

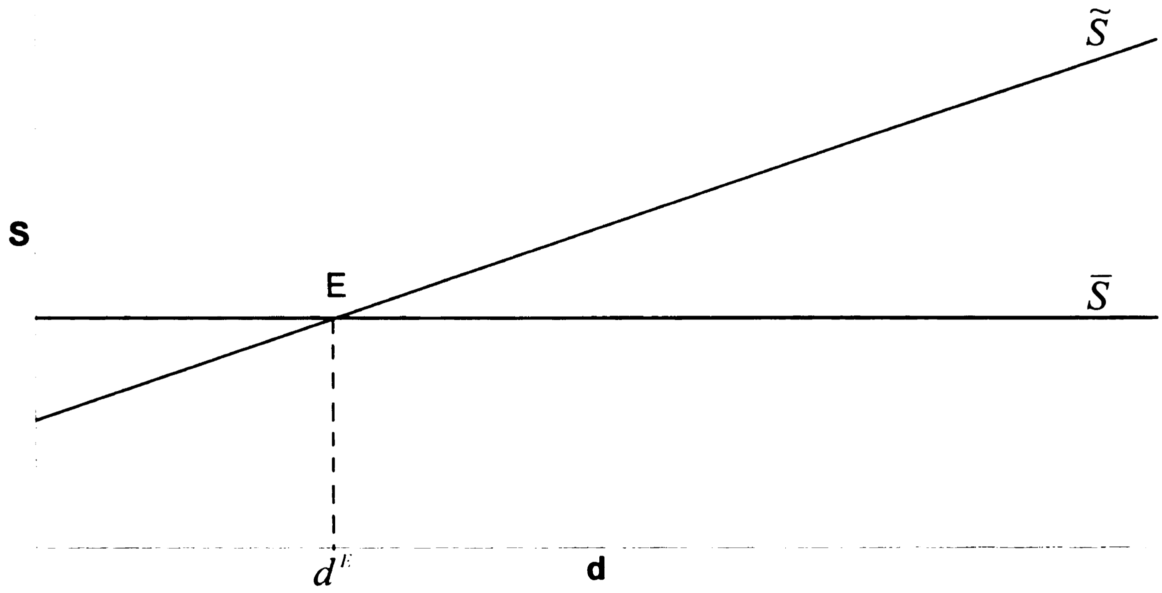

Figure 1. Attack time in a certainty model

When domestic credit d is less than d^E, there is no attack; when $d > d^E$, speculators attack the currency.

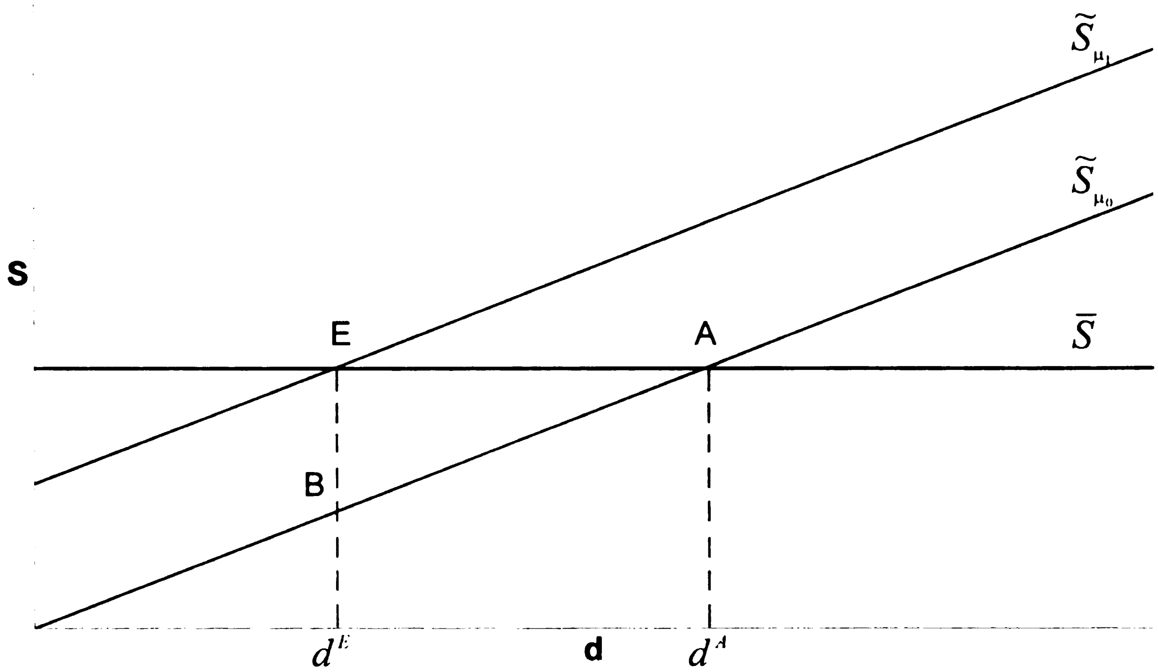

Figure 2. Attack times with attack-conditional policy shift

When $d < d^E$, there is no attack; when $d^E < d < d^A$, multiple equilibria are possible; when $d > d^A$ speculators attack the currency.

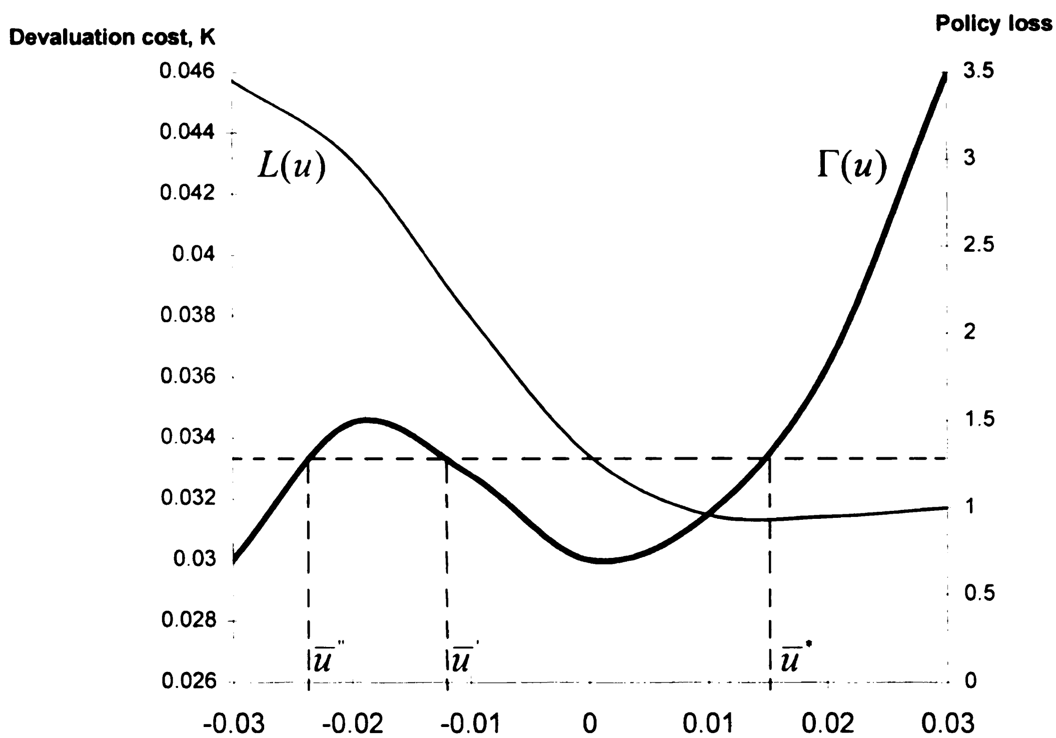

Figure 3. Devaluation cost and policy loss

CHAPTER IV

TIME SERIES PROPERTIES OF VARIABLES IN CURRENCY CRISIS MODEL

1. Introduction

Application of the currency crisis model to the Asian experience in 1997-98 requires the analysis of the time series properties and the making of forecasts for the variables, p^{*}, y, i, i^{*}, d, and u to derive a shadow exchange rate and an one-step-ahead probability of currency crisis.

However, most of the previous studies about the structural analysis of currency crisis, Blanco and Garber (1986), Cumby and Van Wijnbergen (1989), Otker and Pazarbasioglu (1996, 1997b) do not estimate the properties of variables in the model. Instead, they assume an AR(p) model. Unlike those studies, Goldberg (1994) uses ARIMA models and applies Akaike tests to determine which ARIMA process should be used to describe the variables for which forecasts are required. However, while ARIMA models capture autocorrelations that decay at an exponential rate as associated with stationary and invertible ARMA(p,q) models of the first differences of stochastic processes, ARIMA models could not be applied to long memory processes where the autocorrelations decay more slowly than the exponential rate.

Therefore, the first objective of this chapter is to explain the time series properties of macroeconomic variables in our model that exhibit long memory in both their conditional mean and variances. To this end, the ARFIMA(p,d,q)-FIGARCH(P,δ,Q) models are added to the families of models to be selected from by the Wald tests. By making this inclusion, we detect that some processes exhibit dual long memory behavior.

The second objective of this chapter is to forecast of the one-step-ahead levels and volatilities of d, p^*, y, i, i^*, and u which will be used for the derivation of the probability of collapse. Since the ARFIMA(p,d,q)-FIGARCH(P,δ,Q) model lets the stochastic part, ε_t, of each variable be assumed to have a distribution of $\varepsilon_t|\Omega_{t-1} \sim N(0, \sigma_t^2)$, the expected conditional variance, $E_t\sigma_{t+1}^2$, as well as the expected conditional mean, $E_t\mu_{t+1}$, can be obtained to forecast the levels and volatilities of all of the relevant variables[13]. To this end, we estimate the coefficient of each variable in ARFIMA(p,d,q)-FIGARCH(P,δ,Q) model from the analysis of the time series properties of the variables. Then, the estimated coefficients are used to calculate the expected conditional mean and variance.

In this chapter, Wald tests are initially applied to select the appropriate model explaining the behaviors of the variables. Then, once the specific form of the model is determined, the coefficients are estimated and the model is fit over each of the sample time periods to calculate one-step- ahead forecasts.

2. Data set

Monthly data are collected from 1970:01 through 2000:12 for Indonesia, South Korea, Malaysia, Philippines, and Thailand. All of these countries mainly experienced the

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[13] Please see Chapter III for the details.

currency crisis from 1997 to 1998. All of the variables are taken from the CD-ROM version of the International Monetary Fund's *International Financial Statistics* (IFS).[14] Table 17 reports the description and sources of the data. The sample size is dictated by the availability of data on the variables.

The analysis in this chapter differs from previous empirical studies[15] in the use of longer sample time period than preceding studies. In addition, monthly data is expected to capture all the variation in some of the variables for both the months before and after the collapse. A careful identification of all of the activities is crucial for the analysis of the Asian crisis since less frequent data could hide the rapid movements in the second half of 1997.

3. Analysis of time series property and forecast of variables

3.1 ARFIMA-FIGARCH model

Several recent articles have discussed the property of long memory in either the conditional mean or variance of a process.

Granger and Joyeux (1980), Granger (1980, 1981), and Hosking (1981) introduced discrete time representations of fractional Brownian motion known as ARFIMA(p,d,q) processes, which combine the stationary and invertible ARMA model with the fractional difference operator. The model is,

$$\Phi(L)(1-L)^d(y_t-\mu)=\theta(L)\varepsilon_t \qquad (1)$$

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[14] Monthly data for GDP are missing in Indonesia, Thailand, and Philippines. For those countries, quarterly and annual data are used for the analysis.

where, d is a fractional differencing parameter; and $\Phi(L) = 1 - \Phi_1 L - \ldots - \Phi_p L^p$, $\theta(L) = 1 + \theta_1 L + \ldots + \theta_q L^q$, and all the roots of $\varphi(L)$ and $\theta(L)$ lie outside the unit circle for stationarity and invertibility; and $E(\varepsilon_t) = 0$, $E(\varepsilon_t^2) = \sigma^2$ and $E(\varepsilon_t \varepsilon_s) = 0$ for $s \neq t$. The Wold decomposition, or infinite order moving average representation of this process is given by $y_t = \sum_{j=0,\infty} \psi_j \varepsilon_{t-j}$; and the infinite order autoregressive representation is given by $y_t = \sum_{j=1,\infty} \pi_j y_{t-j} + \varepsilon_t$. For high lag j, these coefficients decay at a very slow hyperbolic rate, i.e. $\psi_j \approx c_1 j^{d-1}$ and $\pi_j \approx c_2 j^{-d-1}$, where c_1 and c_2 are constants. For $-0.5 < d < 0.5$, the process is stationary and invertible and y_t is said to be fractionally integrated of order d, or I(d). Therefore, the parameter d represents the degree of "long memory" behavior for the series. For $0.5 \leq d < 1.0$, the process does not have a finite variance, but for $d < 1.0$ the impulse response weights are finite, which implies that shocks to the level of the series are mean reverting.

Time dependent heteroskedasticity in conditional variance is a well-known feature of many asset pricing series and also it is considered useful for some macroeconomic series. Usually the ARCH model of Engle (1982) is introduced as

$$\varepsilon_t = z_t \sigma_t \qquad (2)$$

Where $E_{t-1} z_t = 0$ and $Var_{t-1} z_t = 1$. Throughout this chapter, E_{t-1} and Var_{t-1} refer to the conditional expectation and variance with respect to this same information set. Thus, by

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[15] See appendix 2 for the summary of previous studies.

definition, the $\{\varepsilon_t\}$ process is serially uncorrelated with mean zero. However, the conditional variance of the process, σ_t^2 is a time-varying, positive and measurable function of the information set at time $t-1$. A useful extension to the ARCH model is the GARCH(P,Q) specification of Bollerslev (1986). This model is defined by

$$\sigma_t^2 = \omega + \alpha(L)\varepsilon_{t-1}^2 + \beta(L)\sigma_{t-1}^2 \qquad (3)$$

where, L denotes the lag operator; and $\alpha(L) \equiv \alpha_1 L + \alpha_2 L + \cdots + \alpha_Q L^Q$ and $\beta(L) \equiv \beta_1 L + \beta_2 L + \cdots + \beta_P L^P$. In this model, one can guarantee the stability and covariance stationarity of the $\{\varepsilon_t\}$ process by assuming that all the roots of $\{1-\alpha(L)-\beta(L)\}$ and $\{1-\beta(L)\}$ are constrained to lie outside the unit circle. Otherwise, the GARCH(P,Q) model in equation (3) may also be expressed as an ARMA(M,P) process in ε_t^2,

$$\{1-\alpha(L)-\beta(L)\}\varepsilon_t^2 = \omega + \{1-\beta(L)\}v_t \quad (4)$$

where $M \equiv \max\{P,Q\}$ and $v_t \equiv \varepsilon_t^2 - \sigma_t^2$ is mean zero serially uncorrelated. This model has been useful in describing many volatility processes. However, many high frequency asset pricing series have very persistent volatility. In other words, the autoregressive lag polynomial, $\{1-\alpha(L)-\beta(L)\}$, has a root that is very close or even indistinguishable from unity. This led to the integrated GARCH model or IGARCH model of Engle and Bollerslev (1986). The IGARCH(P,Q) process is defined succinctly by

$$\varphi(L)(1-L)\varepsilon_t^2 = \omega + \{1-\beta(L)\}v_t \quad (5)$$

where, $\varphi(L) \equiv \{1-\alpha(L)-\beta(L)\}(1-L)^{-1}$ is of order $M-1$.

Baillie, Bollerslev and Mikkelsen (1996) introduced the FIGARCH process to model very slow hyperbolic decay in terms of how a shock affects the conditional variance process. The FIGARCH process has the benefit of being a midpoint between the extremes of stationary GARCH and IGARCH. The FIGARCH(P,δ,Q) process is naturally given by,

$$\{1-\beta(L)\}\sigma_t^2 = \omega + \{1-\beta(L)-\varphi(L)(1-L)^{\delta}\}\varepsilon_t^2 \qquad (6)$$

where $0<\delta<1$, all the roots of $\varphi(L)$ and $\{1-\beta(L)\}$ lie outside the unit circle. The FIGARCH process can also be rewritten as,

$$\sigma_t^2 = \omega\{1-\beta(1)\}^{-1} + \{1-[1-\beta(L)]^{-1}\varphi(L)(1-L)^{\delta}\}\varepsilon_t^2 \qquad (7)$$

The FIGARCH(P,δ,Q) model nests the covariance stationary GARCH(P,Q) for $\delta=0$ and IGARCH(P,Q) model for $\delta=1$. Allowing for $0<\delta<1$ gives far more flexibility in modeling long term dependence in the conditional variance. This process implies a slow hyperbolic rate of decay for lagged squared innovations. For instance, the FIGARCH(1,δ,0) process can be expressed as

$$\sigma_t^2 = \omega(1-\beta)^{-1} + \lambda(L)\varepsilon_t^2 \qquad (8)$$

where $\lambda_k = \Gamma(k+\delta-1)/\{\Gamma/\{\Gamma(k)\}\}.\{(1-\beta)-(1-\delta)/k\}$, and for large lags k, $\lambda_k = \{(1-\beta)/\Gamma(\delta)\}\, k^{\delta-1}$, which generates slow hyperbolic rate of decay on the impulse response weights. The process is strictly stationary and ergodic for $0\le\delta\le1$, and shocks will have no permanent effect. It is also worth noting that the FIGARCH model can be expressed as an ARFIMA(P,δ,Q) model in ε_t^2.

$$\varphi(L)(1-L)^{\delta}\varepsilon_t^2 = \omega + \{1-\beta(L)\}v_t \qquad (9)$$

where, $0 < \delta < 1$, and all the roots of $\varphi(L)$ and $\{1-\beta(L)\}$ lie outside the unit circle. $v_t = \varepsilon_t^2 - \sigma_t^2$ is mean zero serially uncorrelated.

To capture the property of economic and financial time series that exhibit long memory in both their conditional mean and variances, the ARFIMA(p,d,q)-FIGARCH(P,δ,Q) model is applied following the result of the Wald tests. The model is given by

$$\Phi(L)(1-L)^d(y_t - \mu) = \theta(L)\varepsilon_t \qquad (10)$$

$$\varepsilon_t = z_t\sigma_t \qquad (11)$$

$$\sigma_t^2 = \omega\{1-\beta(1)\}^{-1} + \{1-[1-\beta(L)]^{-1}\varphi(L)(1-L)^{\delta}\}\varepsilon_t^2 \qquad (12)$$

where $\Phi(L)$, $\theta(L)$, $\beta(L)$, and $\varphi(L)$ are as defined earlier in equation (1), (3), and (5). As noted above, the pure ARFIMA(p,d,q)-homoskedastic process will have a finite variance for $-0.5 < d < 0.5$. However, ARFIMA-FIGARCH process will have an infinite unconditional variance for all d given $\delta \neq 0$. This fact is discussed in the context of the pure FIGARCH model by Baillie, Bollerslev, and Mikkelsen (1996).

Assuming conditional normality, the logarithm of the likelihood can be expressed in the time domain as

$$L(\lambda;\varepsilon_1,\varepsilon_2,\cdots,\varepsilon_T) = -(T/2)\ln(2\pi) - (1/2)\sum_{t=1,T}\{\ln(\sigma_t^2) + \varepsilon_t^2\sigma_t^2\} \qquad (13)$$

where $\lambda' = (\mu,\Phi_1,\cdots,\Phi_p,\theta_1,\cdots,\theta_q,\beta_1,\cdots,\beta_P,\varphi_1,\cdots,\varphi_Q,d,\omega,\delta)$. The QMLE of the parameters are estimated by a similar methodology to that described by Baillie, Bollerslev and Mikkelsen (1996), where the likelihood function is maximized conditional on initial conditions and the pre-sample values of $\varepsilon_t^2, t = 0, -1, -2, \ldots$ are fixed at the sample unconditional variance. The initial observations $y_0, y_{-1}, y_{-2}, \ldots$ are also assumed fixed, in which case minimizing the conditional sum of squares function will be asymptotically equivalent to MLE.

After the specific form of process is selected and the parameters are estimated, the process is fit over each of the samples contained in the entire interval. Following Baillie and Bollerslev (1992), the minimum MSE predictors using the expected conditional means, $E_t\mu_{t+1}$, and conditional variances, $E_t\sigma^2_{t+1}$, are applied to obtain one-step-ahead forecasts of levels and volatilities of variables, d, p^*, y, i, i^*, and u, where

$$E_t\mu_{t+1} = E_t\mu^d_{t+1} + (a_1 + a_3)E_t\mu^i_{t+1} - a_3E_t\mu^{i^*}_{t+1} - a_2E_t\mu^y_{t+1} - E_t\mu^{p^*}_{t+1} - E_t\mu^u_{t+1} \text{ and}$$

$$E_t\sigma^2_{t+1} = E_t(\sigma^d_{t+1})^2 + (a_1 + a_3)^2E_t(\sigma^i_{t+1})^2 + a_3{}^2E_t(\sigma^{i^*}_{t+1})^2$$

$$+ a_2{}^2E_t(\sigma^y_{t+1})^2 + E_t(\sigma^{p^*}_{t+1})^2 + E_t(\sigma^u_{t+1})^2$$ [16]

3.2 Empirical results for US inflation

Table 18 reports the estimation of various univariate models to represent the US CPI inflation data described earlier. The final column of Table 18 estimates an ARFIMA(0,d,1) model with d estimated at around 0.42. The autocorrelations of the standardized residuals are relatively smaller than other models but the squared residuals exhibit autocorrelation that is consistent with very persistent ARCH effects. Figure 4 represents the long run ARCH effects in the squared residuals. In order to allow for ARCH effects, a range of volatility models conditional on the selection of specification for the conditional variance process is considered. The conditional variance processes considered are the FIGARCH(1,δ,1), IGARCH(1,1), and GARCH(1,1) models. The estimates from these models are reported in column 1, 2, and 3 of Table 18. As for the first column, the estimated long memory conditional mean, d, is around 0.43, and significantly different from zero or one. In addition, robust Wald tests can reject the hypothesis that the long memory conditional variance $\delta = 0$ with δ estimated around 0.67. Therefore, we concluded that the estimated ARFIMA(0,d,1)-FIGARCH(1,δ,1) model is the most appropriate specification for accounting for the dynamics of the conditional mean and variance.

Figure 5 provides visual evidence that the persistent ARCH effect of the process vanished after the estimation by the ARFIMA(0,d,1)-FIGARCH(1,δ,1) model. In addition, the $Q^2(20)$ of ARFIMA model, 622.15, significantly decreased to 15.91. This also could be an another confirmation that ARFIMA(0,d,1)-FIGARCH(1,δ,1) model is valuable for eliminating the persistent ARCH effect.

Figure 6 illustrates how well the ARFIMA(0,d,1)-FIGARCH(1,δ,1) model explain the US CPI inflation data. Even though some extreme values are not fitted well by the model, the ARCH effect of the process is effectively demonstrated. Figure 7 shows a one-step-ahead forecast using the model estimated with sample period ending in 1996. It is observed that some of the extreme values are not forecasted appropriately, as are some of the fitted values, but the trend of forecasted values obviously follows that of the actual values of the US CPI inflation data.

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[16] See Chapter III for the details

3.3 Empirical results for deviations from PPP

Table 19 contains the estimation of assorted univariate models to describe the deviations from PPP of Indonesia, South Korea, Malaysia, Philippines, and Thailand. Percentage changes for the deviations from PPP are used for the y_t in the model. This value explains the volatility of deviations more successfully than other values when MLE estimators are chosen to estimate the parameters in the model. In addition, using percentage change for the y_t does not influence significantly any of the forecasted values.

The first column of Table 19 estimates an ARFIMA(0,d,0)-FIGARCH(1,δ,0) model with d estimated around 0.13 and δ assessed around 0.30 for Indonesia. Wald tests significantly reject the hypotheses that $\delta = 0$, indicating strong evidence of long memory in the conditional variance as well as the conditional mean. Figure 8 provides the autocorrelations of the standardized residuals and squared residuals after the estimation. The long memory in conditional mean and variance are noticeably captured. This is shown both in Figure 8 and by how the $Q^2(20)$ estimated around 26.30 could not reject the null hypothesis that there is no autocorrelation in the squared residuals. The rest of columns in Table 19, however, do not include long memory in the conditional mean or in the conditional variance. The relatively small sample size, 324, makes it impossible to estimate long memory for the South Korea, Malaysia, Philippines, and Thailand. Despite the size of the data sets, the relative absence of a long memory property in the conditional mean and variance does not significantly influence the ability of estimation to eliminate autocorrelations from the process. The $Q(20)$ and $Q^2(20)$ of each country's estimation explain that the autocorrelation of the standardized residuals and squared residuals are sufficiently captured by the estimation of the model. Figures 9 to12 offer graphical confirmations of the statistics mentioned above.

3.4 Empirical results for domestic credit

The growth of domestic credit, one of the main causes in the currency crisis, shows various aspects in the estimation of ARFIMA(p,d,q)-FIGARCH(P,δ,Q) models for the domestic credit of each country: Indonesia, South Korea, Malaysia, Philippines, and Thailand. Percentage change of domestic credit is chosen for the y_t.

Table 20 offers the result of the estimation. ARFIMA(0,d,0)-FIGARCH(1,δ,1) and ARFIMA(1,d,0)-FIGARCH(1,δ,1) are chosen for the estimation of Malaysia and Thailand based on the Wald tests. Malaysian percentage change of domestic credit has an estimated d of 0.06 and an estimated δ of 0.52. For Thailand, the estimated long memory conditional mean parameter, d, is –0.06 and the estimated long memory volatility parameter, δ, is 0.71. Indonesia, Philippines, and South Korea, however, do not show any long memory property in the conditional mean or variance. Failure to find evidence for long memory property could be because of the small size of the data set. The autocorrelations of residuals presented in Figures 13-17 indicate that the estimated models did reduce the autocorrelations of the process.

3.5 Empirical results for interest rate

ARFIMA(p,d,q)-FIGARCH(P,δ,Q) models are then estimated for the changes of interest rate in Indonesia, South Korea, Malaysia, Philippines, Thailand, and the US as shown in Table 21. The change of interest rate is chosen for the y_t based on the same criterion used above.

For Indonesia, South Korea, Malaysia, Philippines and US, Wald tests detect a long memory conditional mean parameter, *d*, in the process. However, for South Korea, Malaysia and Philippines, the estimated d lies in the range of $-0.5 < d < 0$. Therefore, the processes have 'intermediate memory', and all their autocorrelations, excluding lag zero, are negative and decay hyperbolically to zero. Whereas Wald tests can detect '*d*' in all country's processes apart from Thailand's, they do not indicate any strong evidence of long memory in the conditional variance for all countries. For Malaysia, Philippines, Thailand and US, the changes of interest rate show a highly persistent volatility shocks represented by IGARCH models. For Indonesia and South Korea, ARFIMA with homoskedasticity models are more representative of the changes of interest rate. As shown by Figures 18 to 23, the autocorrelations of residuals imply that the estimated models do eliminate the autocorrelations of the process for all the countries considered.

3.6 Empirical results for real GDP

ARFIMA(p,d,q)-FIGARCH(P,δ,Q) models are then estimated for real GDP of Indonesia, South Korea, Malaysia, Philippines, and Thailand as shown in Table 22. The rate of change of real GDP[17] is selected for the y_t and monthly, quarterly, or annual data are used to the extent that the highest frequency data are available.

The third column of Table 22 estimates an ARFIMA(1,d,0)-FIGARCH(1,δ,1) model with d and δ estimated around -0.32 and 0.53 for Malaysia. Wald tests significantly reject the hypotheses that $\delta = 0$, indicating strong evidence of long memory in the conditional variance. Figure 26 provides the graphical evidence that the autocorrelations of the standardized residuals and squared residuals are reduced after the estimation for Malaysia. Similarly, the *Q(20)* and *Q²(20)* estimated around 26.94 and 14.02 indicate that the autocorrelation of the standardized residuals and squared residuals are sufficiently captured by the estimation of the model as well.

The rest of columns in Table 22, however, do not contain any long memory in the conditional mean or variance. While the processes of Indonesia and Thailand do not show any long memory for both the conditional mean and variance, Wald tests reveal a long memory conditional mean parameter, *d*, for both South Korean and the Philippines. The autocorrelations of residuals presented in Figures 24-25 and 27-28 indicate that the estimated models did reduce the autocorrelations of the process.

4. Conclusion

This chapter analyzes the time series properties of the macroeconomic variables in Indonesia, South Korea, Malaysia, Philippines, and Thailand which mainly experienced the currency crisis in 1997-98.

[17] 1/10 of the change rate is applied in case of Indonesia, Thailand, and Philippines.

As given in the above sections, some hybrid ARFIMA-FIGARCH models are estimated for the macroeconomic variables. Interestingly, the Wald tests support the conclusion that the U.S. inflation rate, the percentage changes of deviations from PPP in Indonesia, the percentage changes of domestic credits in Malaysia and Thailand and the change rates of real GDP in Malaysia appear to have both estimated long memory parameters d and δ which lie in the range of $-0.5 < d < 0.5$ and $0 < \delta < 1.0$, respectively. However, as shown in Tables 15 to 19, Wald tests do not indicate any dual long memory behavior in other processes.

Following the analysis of time series properties, the expected conditional mean, $E_t\mu_{t+1}$, and conditional variance, $E_t\sigma^2_{t+1}$, obtained using the estimated coefficients in the analysis, are applied for the forecast of levels and volatilities of variables. In particular, the forecasted levels of d, i, i^*, y, p^* and u[18] are substituted for the d_{t+1}, i_{t+1}, i^*_{t+1}, y_{t+1}, p^*_{t+1}, and u_{t+1} to obtain the shadow exchange rate,

$$\tilde{s}_{t+1} = d_{t+1} + r_c - a_0 + (a_1 + a_3)i_{t+1} - a_3 i^*_{t+1} - a_2 y_{t+1} - p^*_{t+1} - u_{t+1} - \bar{s}_t .$$

In addition, the forecasted volatilities, $E_t\sigma^2_{t+1}$, of d, i, i^*, y, p* and u are used for the derivation of probability of collapse,

$$\pi_t = \Pr[\varepsilon_{t+1} \rangle\, k_t] = \int_{k_t}^{\infty} \frac{1}{E_t\sigma_{t+1}\sqrt{2\pi}} e^{-\frac{\varepsilon^2}{2E_t\sigma^2_{t+1}}} d\varepsilon .$$

[18] d, i, i^*, y, p^* and u are assumed to be uncorrelated with each other.

Table 17. Data sources and definitions

Variables	Definition	Sources
d_t	Log of the M2 minus log of the central bank total reserves	IFS(line11.d,14.a,14.b),International Monetary Fund
r_t	Log of the central bank total reserves minus gold in millions of US dollars	IFS(line11.d), International Monetary Fund
i_t	Money market rate*	IFS(line60b,60c), International Monetary Fund
i_t^*	Treasury bill rate for the United States**	IFS(line60c), International Monetary Fund
p_t	Log of the consumer price index	IFS(line64), International Monetary Fund
p_t^*	Log of the U.S. consumer price index	IFS(line64), International Monetary Fund
y_t	Log of the nominal gross domestic product deflated using consumer price index*** in domestic currency	IFS(line99b), International Monetary Fund
s_t	Log of the end-of-period spot exchange rate	IFS(line60c), International Monetary Fund

Notes. * For Indonesia, monthly data from 1983:01 through 2000:10 are collected. IFS defines the maturity of the money market rate to be short-term. For South Korea, monthly data from 1976:08 through 2000:11are collected. The maturity of the money market rate is defined to be short-term. For Malaysia, monthly data from 1970:01 through 2000:11 are collected. The maturity of money market rate is not defined. For Philippines, monthly data from 1976:01 through 2000:12 are collected. The maturity of treasury bill rate is defined to be 91 days. For Thailand, monthly data from 1976:01 through 2000:12 are collected. The maturity of money market rate is not defined.

** For U.S., monthly data from 1970:01 through 2000:12 are collected. The maturity of treasury bill rate is defined to be 91 days.

*** For Indonesia, annual data from 1970 through 1998 are collected. For South Korea, monthly data of log of the industrial production index from 1970:01 through 2000:11 are collected. For Malaysia, monthly data of log of the industrial production index from 1971:01 through 2000:11 are collected. For Philippines, quarterly data from 1981: 1Q through 2000: 3Q are collected. For Thailand, annual data from 1970 through 1998 are collected.

Table 18. Estimated ARFIMA(p,d,q)-FIGARCH(P,δ,Q) models for US monthly inflation rate

$$(1-L)^d(y_t-\mu)=(1+\theta L)(1+\Theta L^{12})\varepsilon_t,$$

$$\varepsilon_t|\Omega_{t-1}\sim N(0,\sigma_t^2),$$

$$(1-\beta L)\sigma_t^2=\omega+[1-\beta L-\varphi L(1-L)^{\delta}]\varepsilon_t^2$$

(p,d,q)- (P,δ,Q)	(0,d,1)- (1,δ,1)	(0,d,1)- (1,1,0)	(0,d,1)- (1,0,1)	(0,d,1)- (0,0,0)
μ	0.5477 (0.2925)	0.5697 (0.4864)	0.5704 (0.5727)	0.3848 (0.2666)
d	0.4324 (0.1210)	0.4573 (0.1791)	0.4612 (0.2055)	0.4222 (0.0920)
θ	-0.2250 (0.1333)	-0.2501 (0.2005)	-0.2496 (0.2266)	-0.1754 (0.1275)
Θ	0.1295 (0.0365)	0.1451 (0.0362)	0.1449 (0.0383)	0.0428 (0.0591)
δ	0.6696 (0.3390)	1.0000 (-)	- (-)	- (-)
ω	0.0012 (0.0012)	0.0019 (0.0016)	0.0038 (0.0022)	0.1054 (0.0107)
β	0.8085 (0.1291)	0.8499 (0.0684)	0.8343 (0.0685)	- (-)
φ	0.4416 (0.1254)	- (-)	0.9548 (0.0224)	- (-)
LL	-79.942	-86.929	-83.494	-176.190
$Q(20)$	34.0424	33.2755	32.7610	30.8235
$Q^2(20)$	15.9059	17.9971	19.5660	622.1546
Skewness	0.105	0.164	0.175	0.137
Kurtosis	4.419	4.611	4.456	7.150
$W_{d=1}$	21.989	9.186	6.876	39.397
$W_{\delta=0}$	3.902	-	-	-

Key: The table reports the Quasi Maximum Likelihood Estimates (QMLE) for various ARFIMA-FIGARCH models. The QMLE are calculated using the normal likelihood function. Robust standard errors are reported in parentheses. LL is the value of the maximized Gaussian log likelihood; and Q(20) and $Q^2(20)$ are the Ljung-Box test statistics with 20 degrees of freedom based on the standardized residuals and squared standardized residuals respectively. $W_{d=1}$ and $W_{\delta=0}$ are both Wald test statistics for testing long memory property in the conditional mean and variance.

Table 19. Estimated ARFIMA(p,d,q)-FIGARCH(P,δ,Q) models for deviations from PPP

$$(1-\Phi L)(1-L)^d(y_t-\mu)=(1+\theta L)\varepsilon_t,$$

$$\varepsilon_t|\Omega_{t-1}\sim N(0,\sigma_t^2),$$

$$(1-\beta L)\sigma_t^2=\omega+[1-\beta L-\varphi L(1-L)^\delta]\varepsilon_t^2$$

Country	Indonesia	South Korea	Malaysia	Philippines	Thailand
(p,d,q)- (P,δ,Q)	(0,d,0)- (1,δ,0)	(0,d,1)- (1,0,1)	(0,d,0)- (1,0,1)	(1,0,1)- (1,δ,0)	(0,0,1)- (1,0,1)
μ	0.1266 (0.1278)	0.1231 (0.2069)	-0.1409 (0.2244)	0.1080 (0.1544)	0.0588 (0.0522)
d	0.1305 (0.0488)	0.2458 (0.0589)	0.1461 (0.0667)	- (-)	- (-)
Φ	- (-)	- (-)	- (-)	0.7937 (0.2515)	- (-)
θ	- (-)	0.1798 (0.0729)	- (-)	-0.6902 (0.2088)	0.1215 (0.0580)
δ	0.2977 (0.1025)	- (-)	- (-)	0.7421 (0.1316)	- (-)
ω	0.1460 (0.1140)	0.0572 (0.0322)	0.8711 (0.7754)	0.1241 (0.1352)	0.0727 (0.0367)
β	0.0221 (0.1358)	0.8363 (0.0447)	0.4375 (0.4442)	0.6266 (0.1265)	0.8236 (0.0542)
φ	- (-)	0.9354 (0.0412)	0.7081 (0.2135)	- (-)	0.9015 (0.0539)
LL	-468.249	-419.207	-602.231	-700.643	-402.756
$Q(20)$	45.0589	14.6086	34.6525	44.1258	29.5997
$Q^2(20)$	26.3045	19.8895	14.3141	19.1193	21.9693
Skewness	1.094	0.287	0.355	-2.426	0.326
Kurtosis	5.542	3.743	8.022	14.604	3.668
$W_{d=1}$	317.057	163.865	164.009	-	-
$W_{\delta=0}$	8.439	-	-	31.805	-

Key: The table reports the Quasi Maximum Likelihood Estimates (QMLE) for various ARFIMA-FIGARCH models. The QMLE are calculated using the normal likelihood function. Robust standard errors are reported in parentheses. LL is the value of the maximized Gaussian log likelihood; and Q(20) and $Q^2(20)$ are the Ljung-Box test statistics with 20 degrees of freedom based on the standardized residuals and squared standardized residuals respectively. $W_{d=1}$ and $W_{\delta=0}$ are both Wald test statistics for testing long memory property in the conditional mean and variance.

Table 20. Estimated ARFIMA(p,d,q)-FIGARCH(P,δ,Q) models for domestic credit

$$(1-\Phi L)(1-L)^d(y_t-\mu)=(1+\theta L)(1+\Theta L^k)\varepsilon_t,$$

$$\varepsilon_t|\Omega_{t-1}\sim N(0,\sigma_t^2),$$

$$(1-\beta L)\sigma_t^2=\omega+[1-\beta L-\varphi L(1-L)^\delta]\varepsilon_t^2$$

Country	Indonesia	South Korea	Malaysia	Philippines	Thailand
(p,d,q)- (P,δ,Q)	(0,d,0)- (1,1,0)	(1,0,0)- (1,1,0)	(0,d,0)- (1,δ,1)	(0,0,1)- (1,0,1)	(1,d,0)- (1,δ,1)
μ	2.0421 (0.0681)	1.5424 (0.1684)	1.3342 (0.1469)	1.7938 (0.3218)	1.4117 (0.1092)
d	-0.2015 (0.0629)	- (-)	0.0588 (0.0403)	- (-)	-0.0580 (0.0790)
Φ	- (-)	-0.0550 (0.0730)	- (-)	- (-)	0.2712 (0.1130)
θ	- (-)	- (-)	- (-)	0.0377 (0.1078)	- (-)
Θ	0.1719* (0.0589)	0.4153** (0.0429)	0.2200** (0.0413)	- (-)	0.3118** (0.0441)
δ	1.0000 (-)	- (-)	0.5236 (0.0923)	- (-)	0.7105 (0.2448)
ω	0.3320 (0.5631)	2.0517 (0.5176)	0.0416 (0.0322)	0.5906 (0.5310)	0.0079 (0.0150)
β	0.8137 (0.0877)	0.1549 (0.1072)	0.8788 (0.0384)	0.8533 (0.0589)	0.9119 (0.0463)
φ	- (-)	- (-)	0.3989 (0.0660)	0.9625 (0.0447)	0.4632 (0.1937)
LL	-917.429	-701.616	-574.283	-324.310	-591.442
$Q(20)$	70.4911	34.1811	18.0219	56.4719	54.1633
$Q^2(20)$	8.2300	24.2733	25.6956	13.4509	16.6580
Skewness	0.324	0.390	0.207	-0.011	0.284
Kurtosis	6.845	4.570	4.751	3.320	3.675
$W_{d=1}$	364.371	-	546.711	-	179.152
$W_{\delta=0}$	-	-	32.182	-	8.426

Key: The table reports the Quasi Maximum Likelihood Estimates (QMLE) for various ARFIMA-FIGARCH models. The QMLE are calculated using the normal likelihood function. Robust standard errors are reported in parentheses. LL is the value of the maximized Gaussian log likelihood; and Q(20) and $Q^2(20)$ are the Ljung-Box test statistics with 20 degrees of freedom based on the standardized residuals and squared standardized residuals respectively. $W_{d=1}$ and $W_{\delta=0}$ are both Wald test statistics for testing long memory property in the conditional mean and variance. $*k = 3$, $**\ k = 12$

Table 21. Estimated ARFIMA(p,d,q)-FIGARCH(P,δ,Q) models for interest rate

$$(1-\Phi L)(1-L)^d(y_t-\mu)=(1+\theta L)\varepsilon_t,$$

$$\varepsilon_t|\Omega_{t-1}\sim N(0,\sigma_t^2),$$

$$(1-\beta L)\sigma_t^2=\omega+[1-\beta L-\varphi L(1-L)^\delta]\varepsilon_t^2$$

Country	Indonesia	South Korea	Malaysia	Philippines	Thailand	US
(p,d,q)-(P,δ,Q)	(1,d,0)-(1,0,1)	(0,d,0)-(1,0,1)	(0,d,0)-(1,1,0)	(0,d,1)-(1,1,0)	(0,0,1)-(1,1,0)	(0,d,0)-(1,1,0)
μ	-0.0057	-0.0301	0.0193	0.0172	0.1314	0.0380
	(0.0150)	(0.0392)	(0.0118)	(0.0157)	(0.0669)	(0.1461)
d	0.7942	-0.1023	-0.2919	-0.2791	-	0.3397
	(0.1778)	(0.0575)	(0.1374)	(0.1078)	(-)	(0.0862)
Φ	0.7870	-	-	-	-	-
	(0.0865)	(-)	(-)	(-)	(-)	(-)
θ	-	-	-	0.4394	0.3021	-
	(-)	(-)	(-)	(0.0899)	(0.0761)	(-)
δ	-	-	1.0000	1.0000	1.0000	1.0000
	(-)	(-)	(-)	(-)	(-)	(-)
ω	0.9676	0.0341	0.0298	0.0131	0.0668	0.0031
	(0.3360)	(0.0228)	(0.168)	(0.0089)	(0.0266)	(0.0021)
β	0.0956	0.9163	0.6295	0.7130	0.6117	0.7453
	(0.1276)	(0.0339)	(0.0698)	(0.0694)	(0.0572)	(0.0591)
φ	0.9809	0.9730	-	-	-	-
	(0.2460)	(0.0283)	(-)	(-)	(-)	(-)
LL	-234.013	-348.631	-396.977	-343.507	-374.630	-132.241
Q(20)	20.4605	17.2562	42.1987	41.3027	18.6411	30.4143
$Q^2(20)$	9.9756	11.2593	22.4051	14.4615	21.7732	19.9586
Skewness	1.135	0.889	1.912	0.783	0.050	0.082
Kurtosis	6.938	6.126	12.036	8.251	3.625	5.054
$W_{d=1}$	101.788	368.118	88.384	140.725	-	58.677
$W_{\delta=0}$	-	-	-	-	-	-

Key: The table reports the Quasi Maximum Likelihood Estimates (QMLE) for various ARFIMA-FIGARCH models. The QMLE are calculated using the normal likelihood function. Robust standard errors are reported in parentheses. LL is the value of the maximized Gaussian log likelihood; and Q(20) and $Q^2(20)$ are the Ljung-Box test statistics with 20 degrees of freedom based on the standardized residuals and squared standardized residuals respectively. $W_{d=1}$ and $W_{\delta=0}$ are both Wald test statistics for testing long memory property in the conditional mean and variance.

Table 22. Estimated ARFIMA(p,d,q)-FIGARCH(P,δ,Q) models for real GDP

$$(1-\Phi L)(1-L)^d(y_t-\mu)=(1+\theta L)(1+\Theta L^k)\varepsilon_t,$$

$$\varepsilon_t|\Omega_{t-1}\sim N(0,\sigma_t^2),$$

$$(1-\beta L)\sigma_t^2=\omega+[1-\beta L-\varphi L(1-L)^\delta]\varepsilon_t^2$$

Country	Indonesia	South Korea	Malaysia	Philippines	Thailand
(p,d,q)- (P,δ,Q)	(0,0,1)- (0,0,0)	(0,d,0)- (1,0,1)	(1,d,0)- (1,δ,1)	(1,d,0)- (0,0,0)	(1,0,0)- (0,0,0)
μ	0.8294 (0.1146)	0.9631 (0.0975)	0.8122 (0.0438)	0.0489 (0.0198)	0.6762 (0.1083)
d	- (-)	-0.1072 (0.0436)	-0.3236 (0.0707)	-0.3361 (0.1303)	- (-)
Φ	- (-)	- (-)	-0.3271 (0.0885)	-0.5115 (0.1093)	0.4242 (0.1816)
θ	0.2556 (0.1477)	- (-)	- (-)	- (-)	- (-)
Θ	- (-)	- (-)	0.3769*** (0.0477)	0.7941**** (0.0584)	- (-)
δ	- (-)	- (-)	0.5282 (0.2278)	- (-)	- (-)
ω	0.2290 (0.0687)	0.8691 (1.1564)	0.1561 (0.3233)	0.2366 (0.0468)	0.1134 (0.0321)
β	- (-)	0.7731 (0.2084)	0.8480 (0.0666)	- (-)	- (-)
φ	- (-)	0.8928 (0.1382)	0.5684 (0.1913)	- (-)	- (-)
LL	-17.731	-778.994	-922.806	-43.990	-8.592
Q(20)	8.0922*	44.6843	26.9356	95.3482	11.2687*
$Q^2(20)$	11.5002**	14.4686	14.0204	29.9340	4.7804**
Skewness	-0.057	-0.795	-0.193	0.256	-0.256
Kurtosis	3.336	7.271	4.948	3.464	3.089
$W_{d=1}$	-	644.669	350.816	105.088	-
$W_{\delta=0}$	-	-	5.374	-	-

Key: The table reports the Quasi Maximum Likelihood Estimates (QMLE) for various ARFIMA-FIGARCH models. The QMLE are calculated using the normal likelihood function. Robust standard errors are reported in parentheses. LL is the value of the maximized Gaussian log likelihood; and Q(20) and $Q^2(20)$ are the Ljung-Box test statistics with 20 degrees of freedom based on the standardized residuals and squared standardized residuals respectively. $W_{d=1}$ and $W_{\delta=0}$ are both Wald test statistics for testing long memory property in the conditional mean and variance. * *Q(10)*, ** Q^2 *(10)*, *** $k = 12$, **** $k = 4$

Figure 4. Correlograms of standardized residuals from ARFIMA (0,*d*,1) model for U.S. CPI inflation

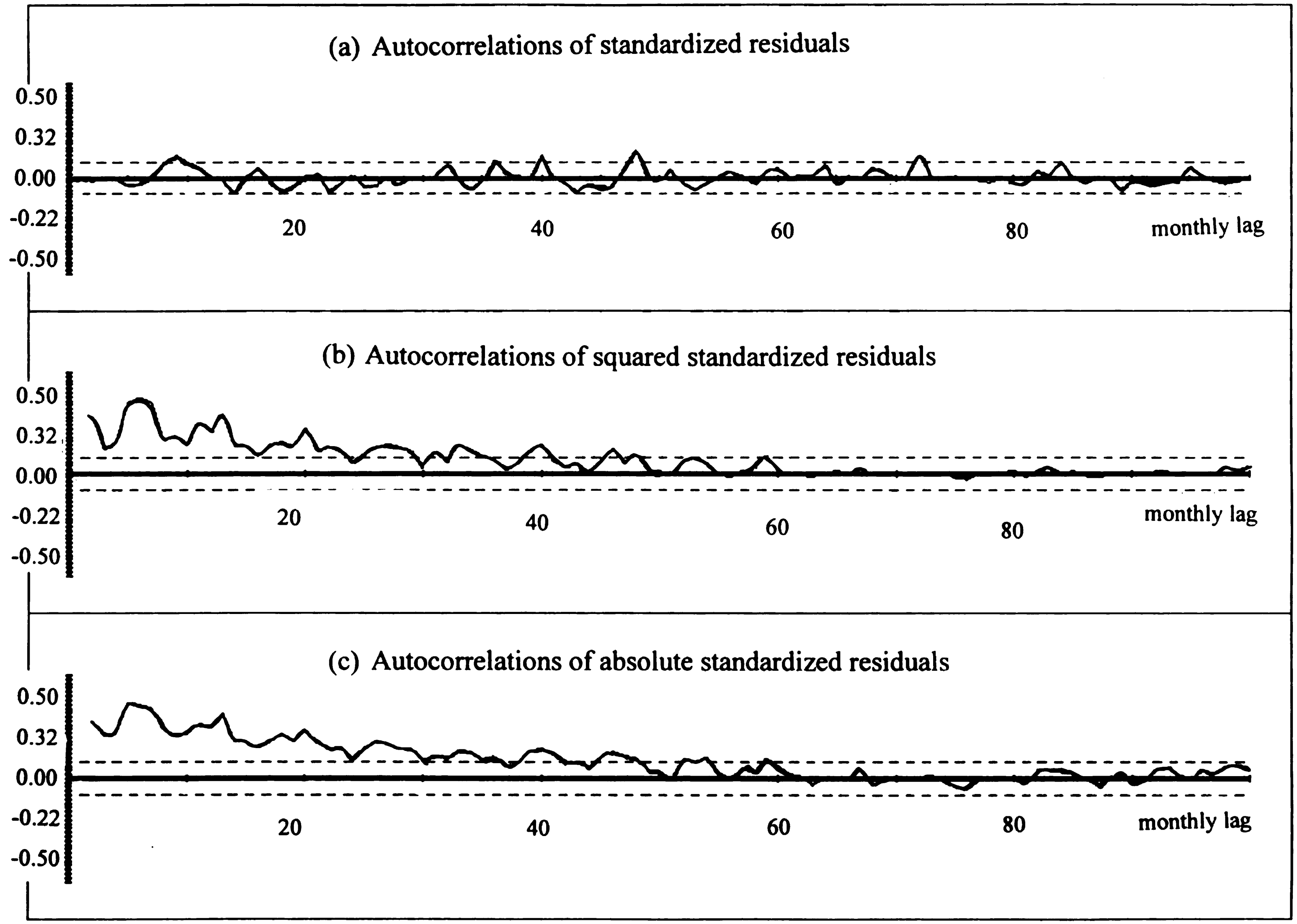
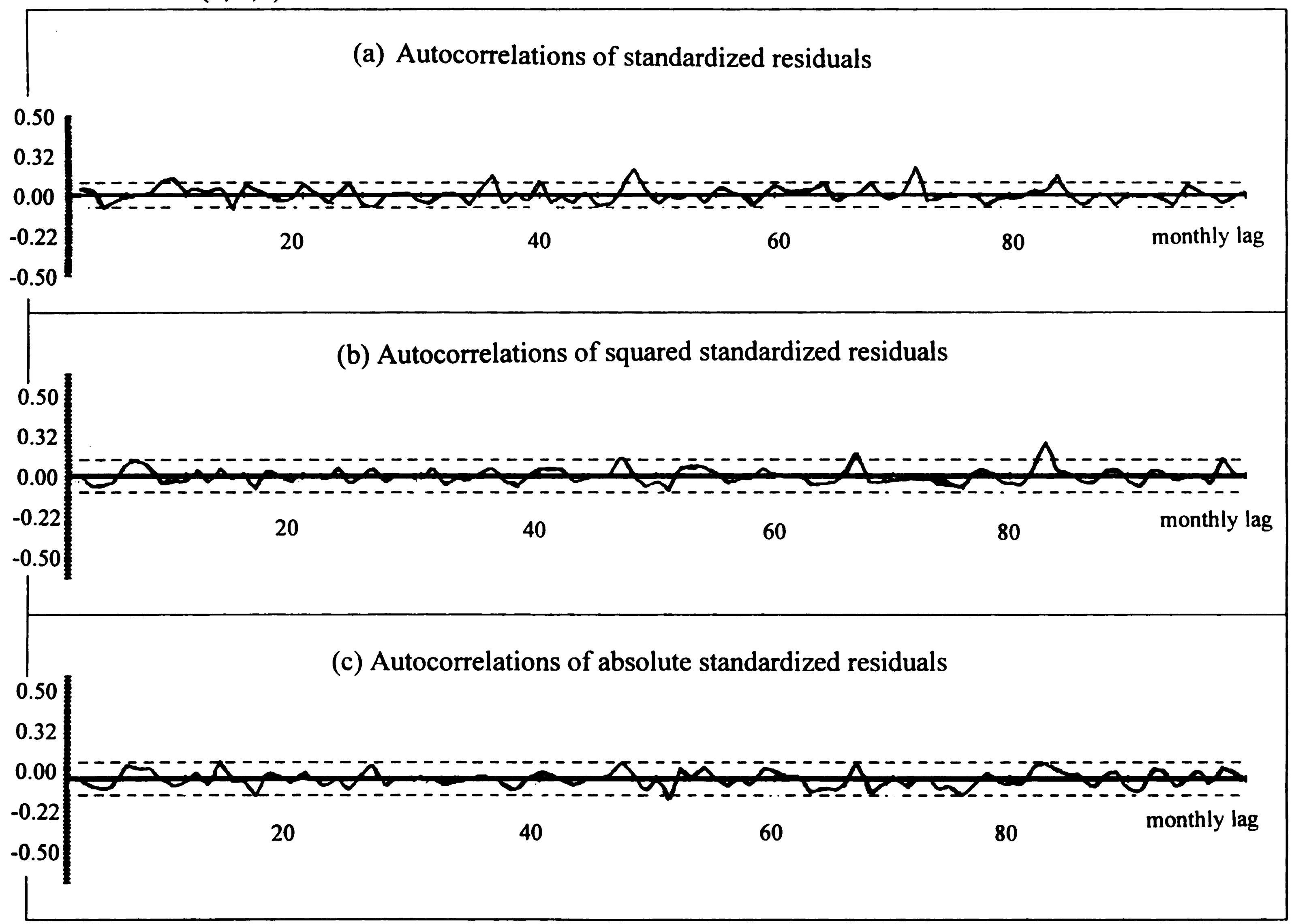


Figure 5. Correlograms of standardized residuals from ARFIMA (0,d,1)-FIGARCH (1,δ,1) model for U.S. CPI inflation



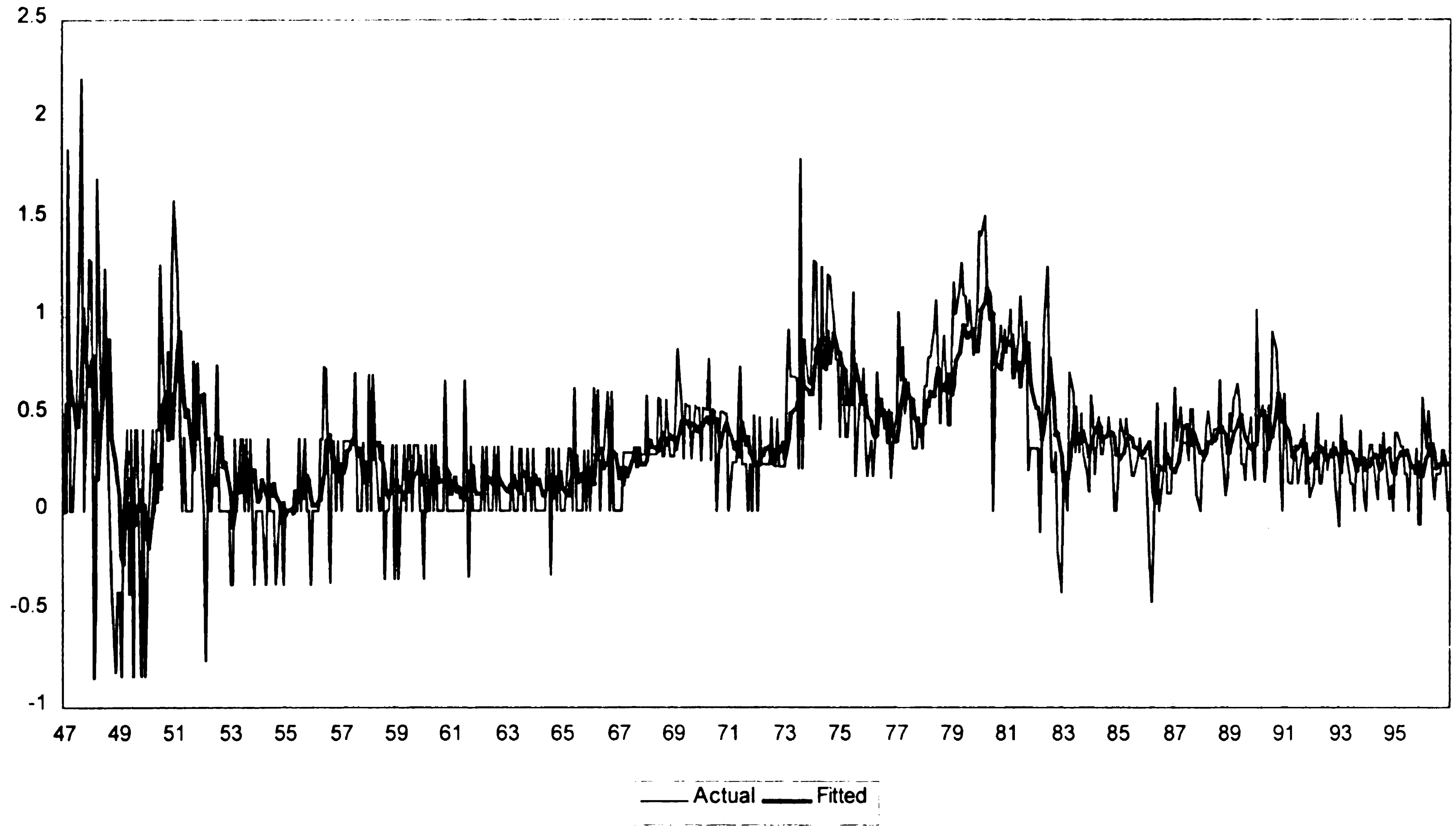


Figure 6. Actual and fitted values of U.S. CPI inflation rate

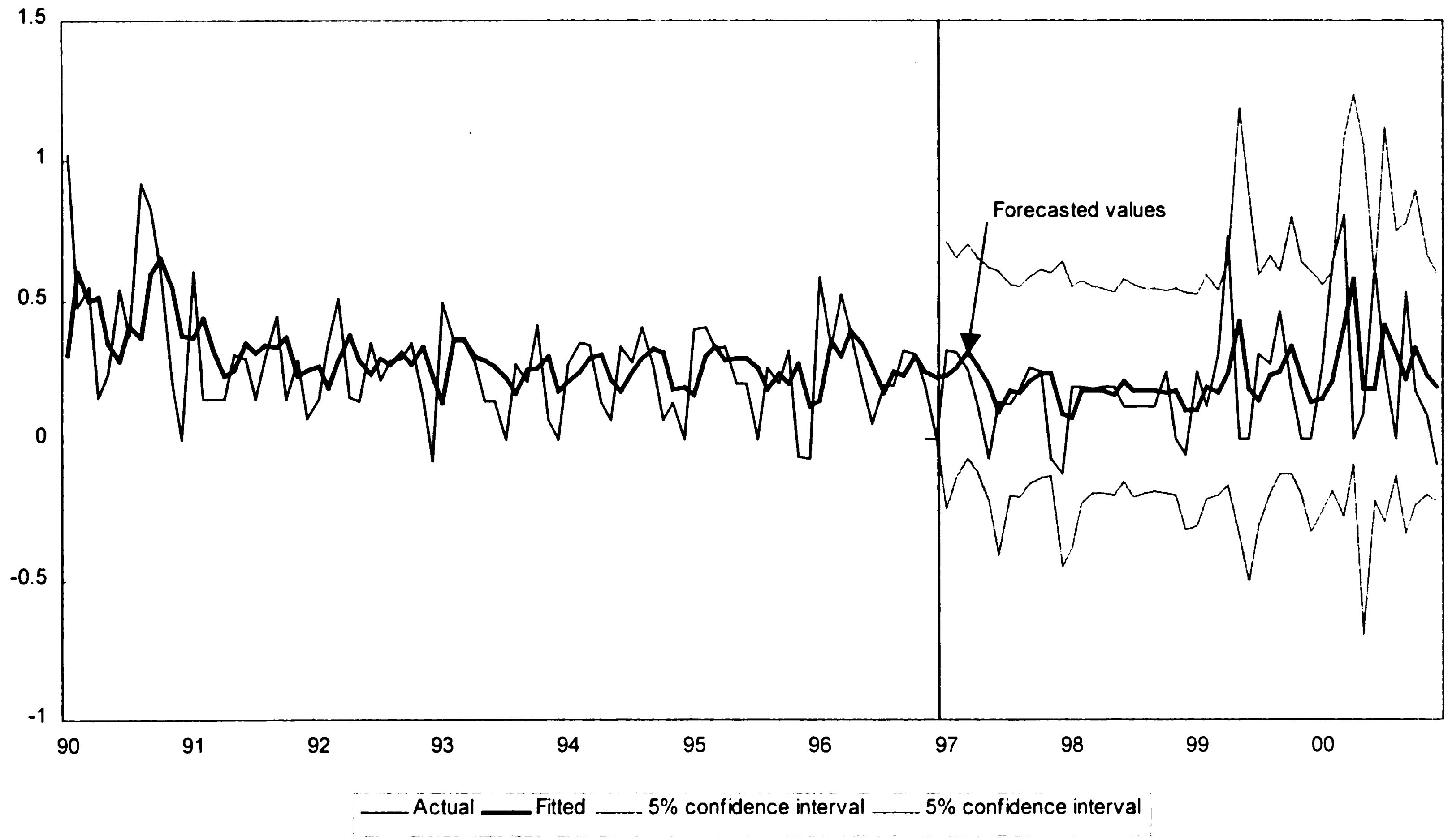


Figure 7. Forecasted values of U.S. CPI inflation rate

Figure 8. Correlograms of standardized residuals for deviations from PPP, Indonesia

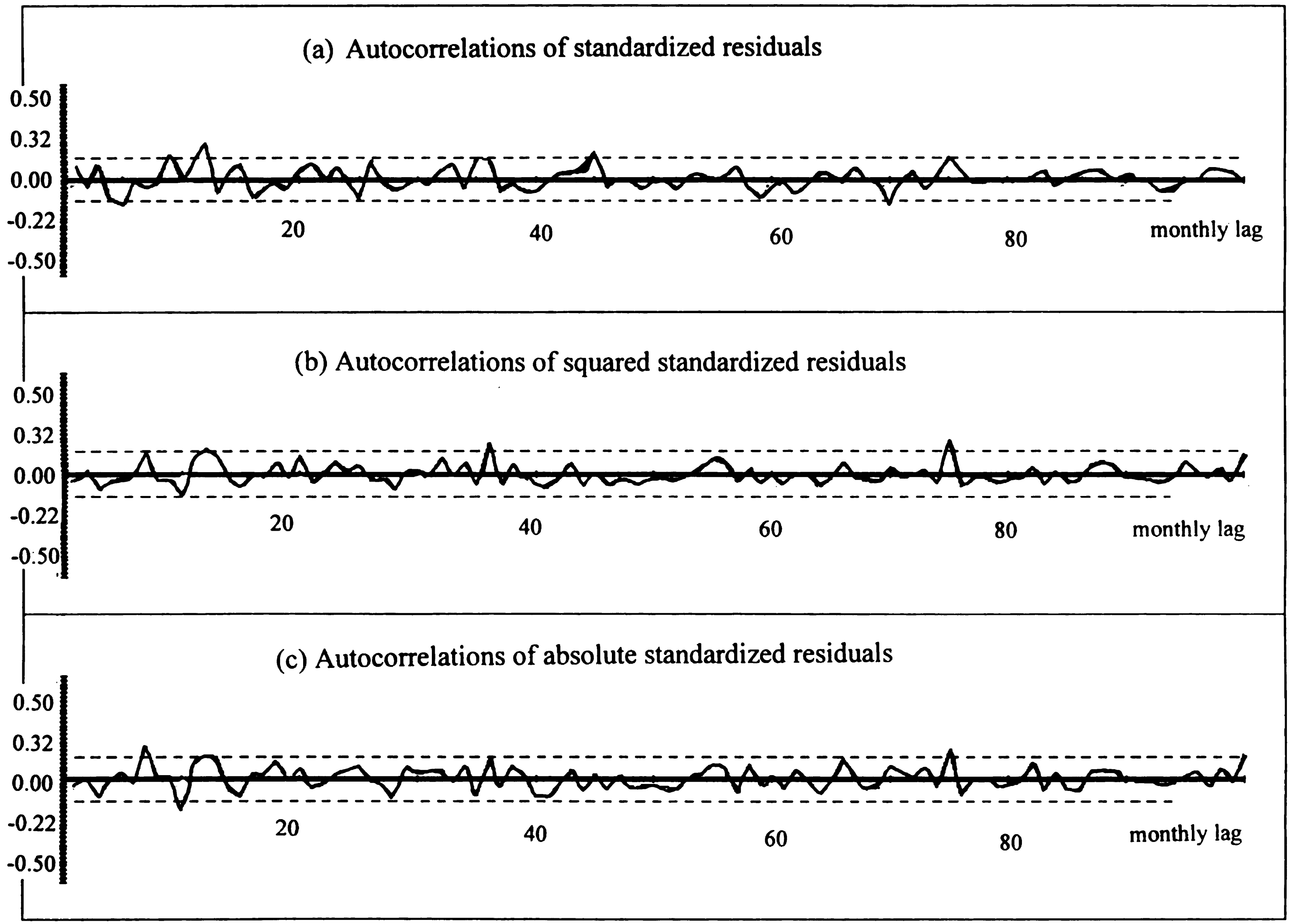


Figure 9. Correlograms of standardized residuals for deviations from PPP, South Korea

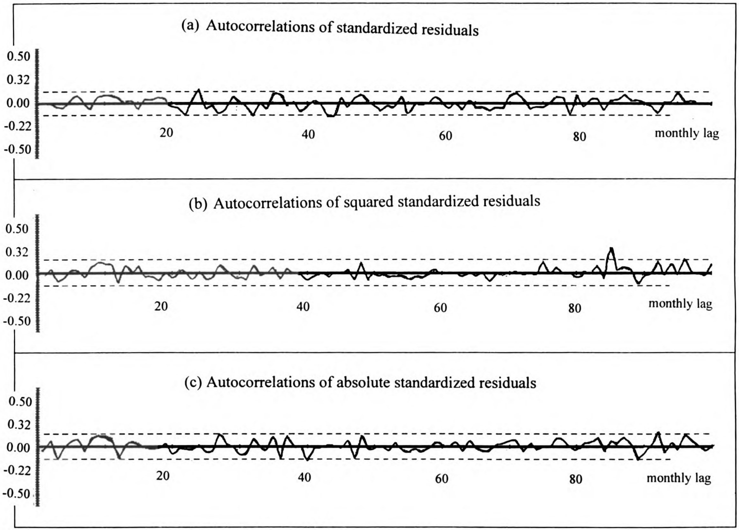


Figure 10. Correlograms of standardized residuals for deviations from PPP, Malaysia

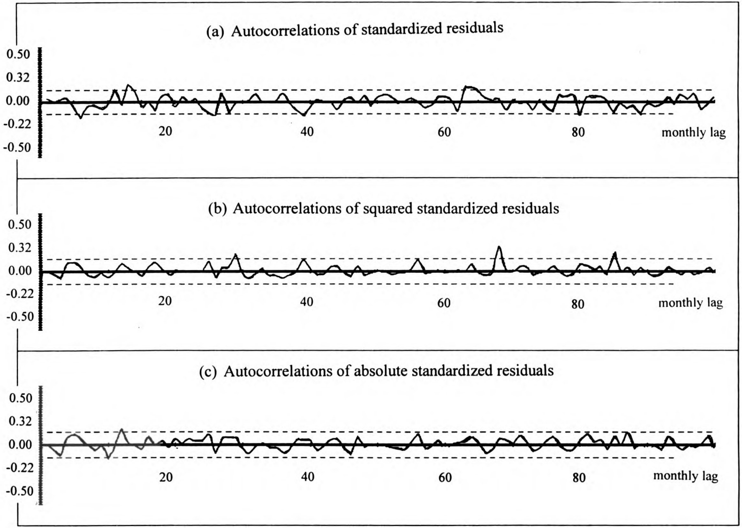


Figure 11. Correlograms of standardized residuals for deviations from PPP, Philippines

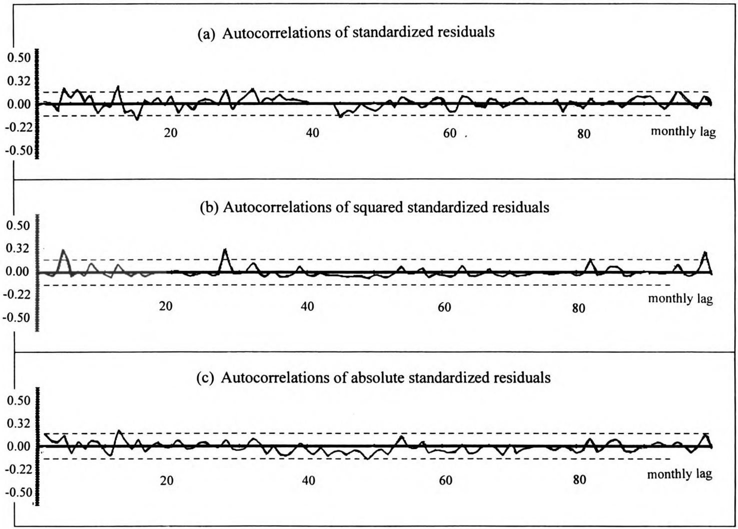


Figure 12. Correlograms of standardized residuals for deviations from PPP, Thailand

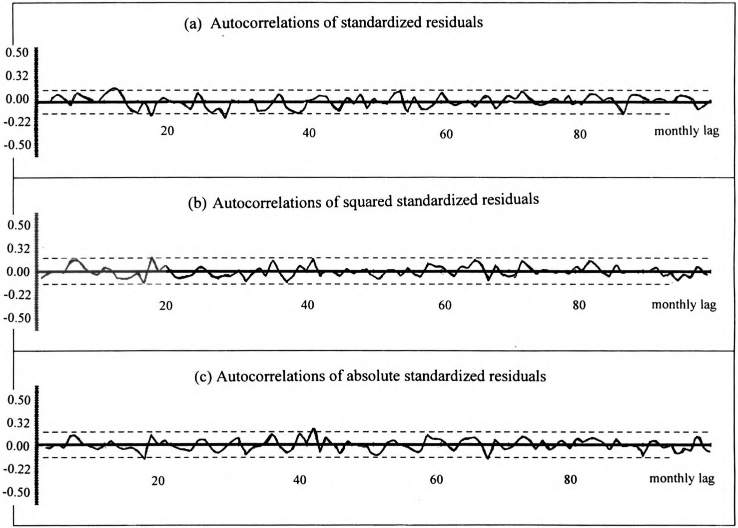


Figure 13. Correlograms of standardized residuals for domestic credit, Indonesia

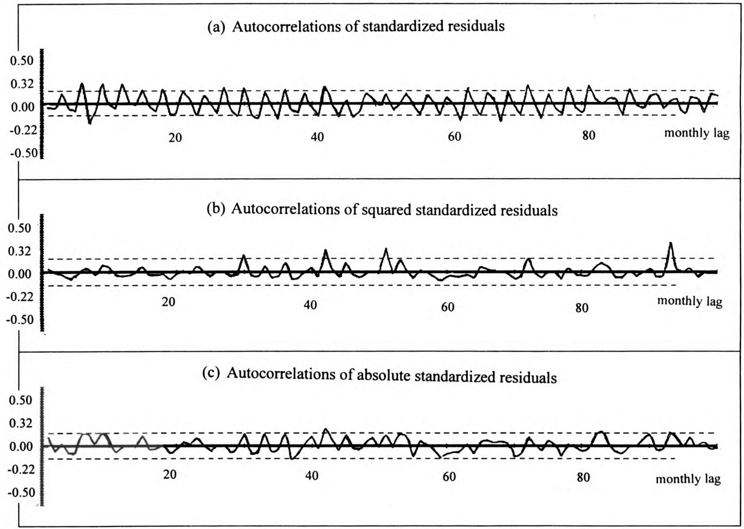


Figure 14. Correlograms of standardized residuals for domestic credit, South Korea

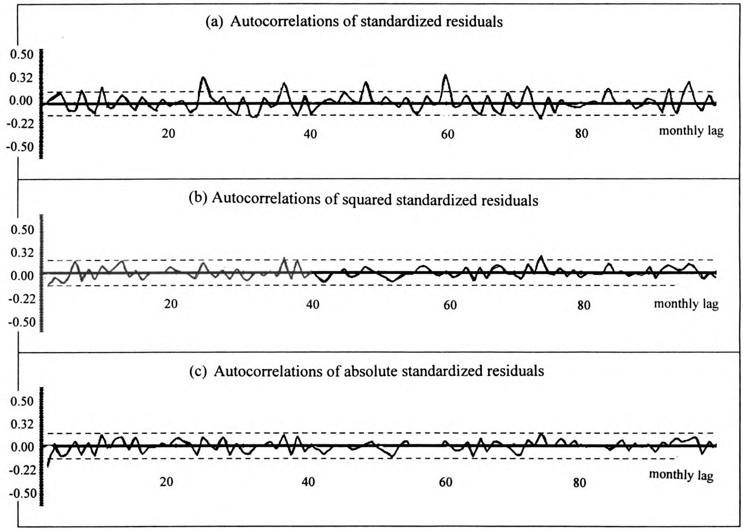


Figure 15. Correlograms of standardized residuals for domestic credit, Malaysia

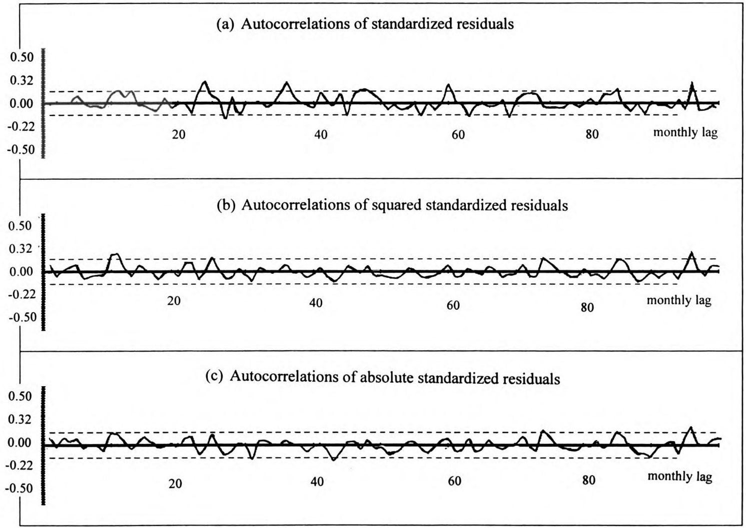


Figure 16. Correlograms of standardized residuals for domestic credit, Philippines

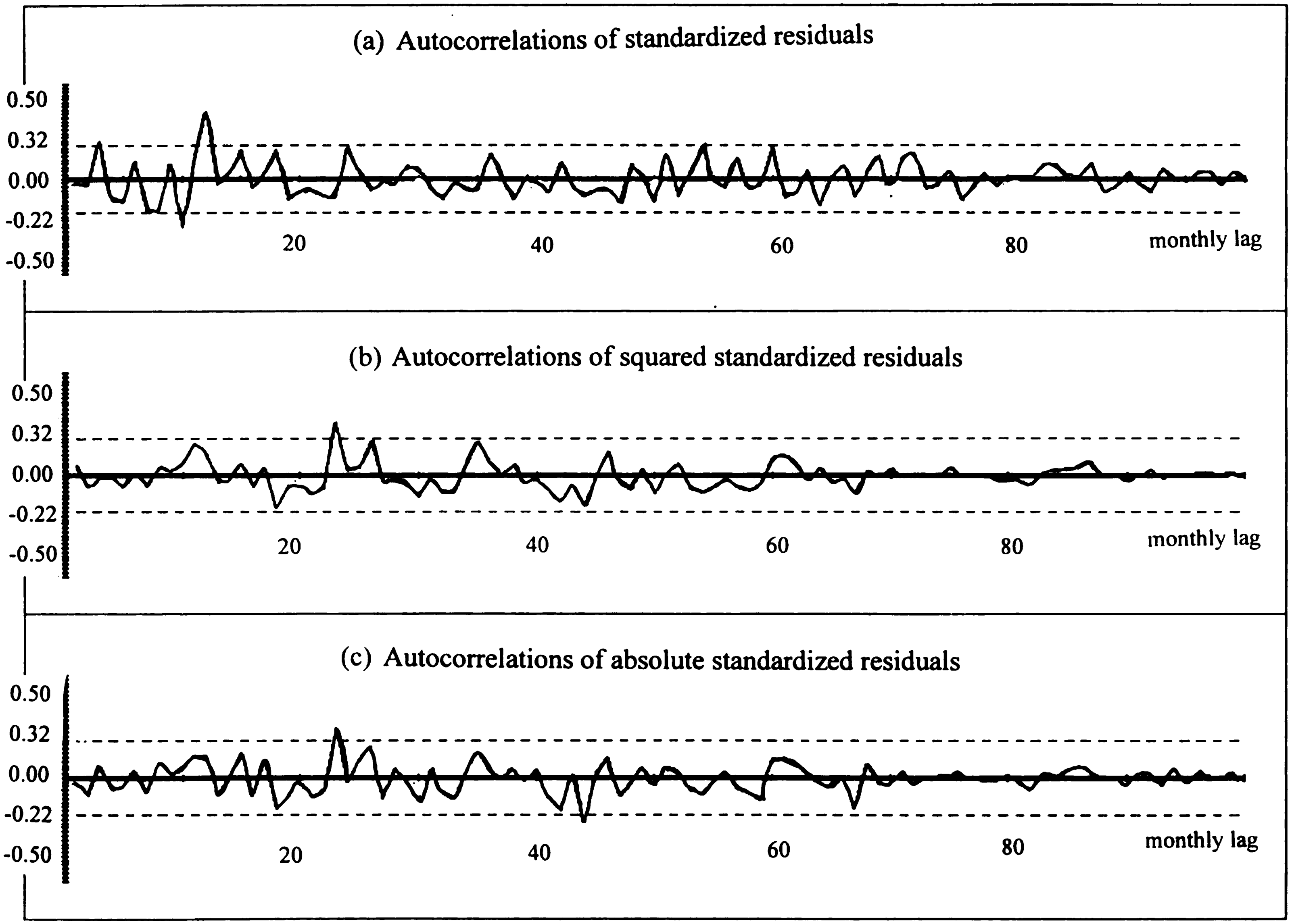


Figure 17. Correlograms of standardized residuals for domestic credit, Thailand

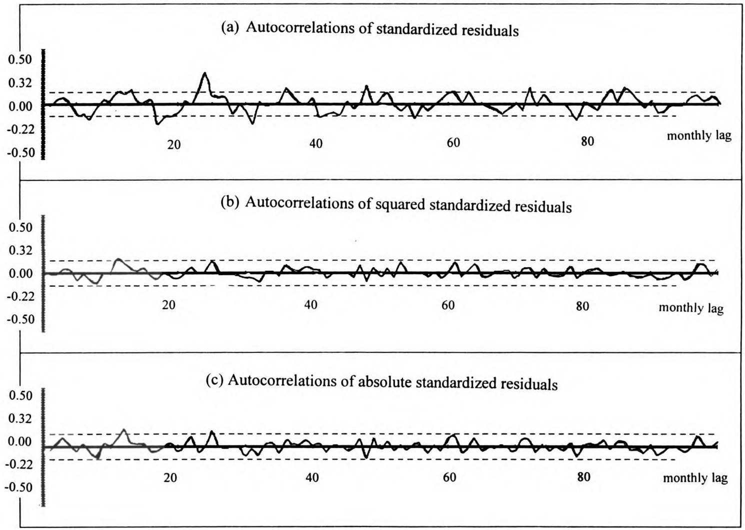


Figure 18. Correlograms of standardized residuals for interest rate, Indonesia

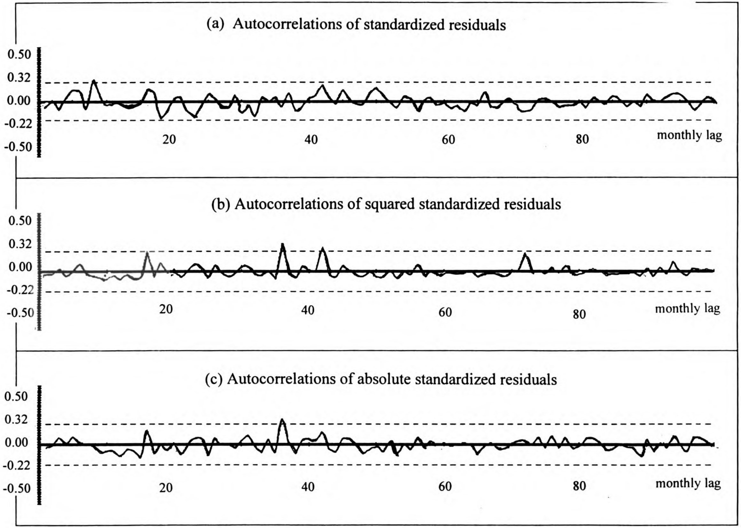


Figure 19. Correlograms of standardized residuals for interest rate, South Korea

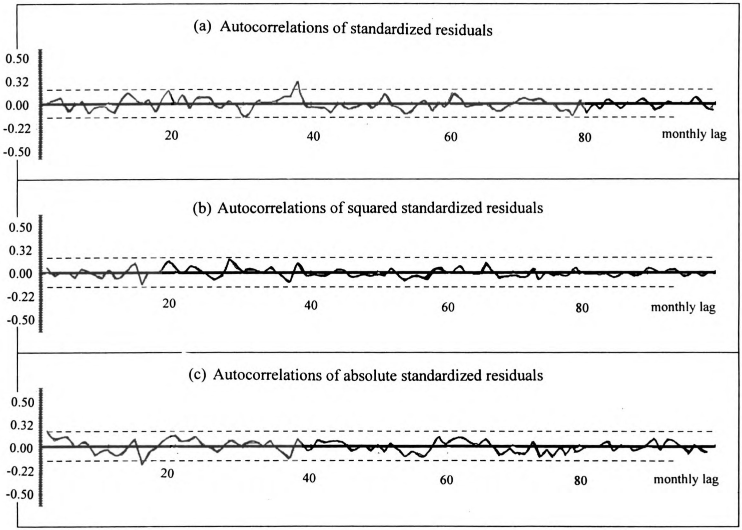


Figure 20. Correlograms of standardized residuals for interest rate, Malaysia

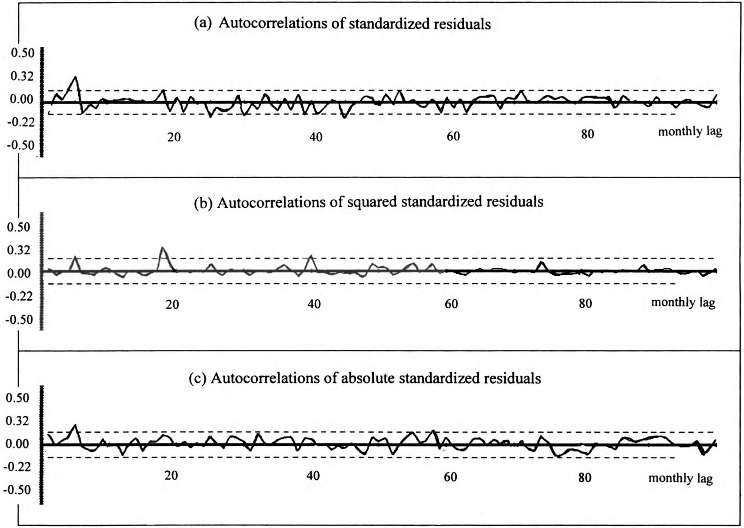


Figure 21. Correlograms of standardized residuals for interest rate, Philippines

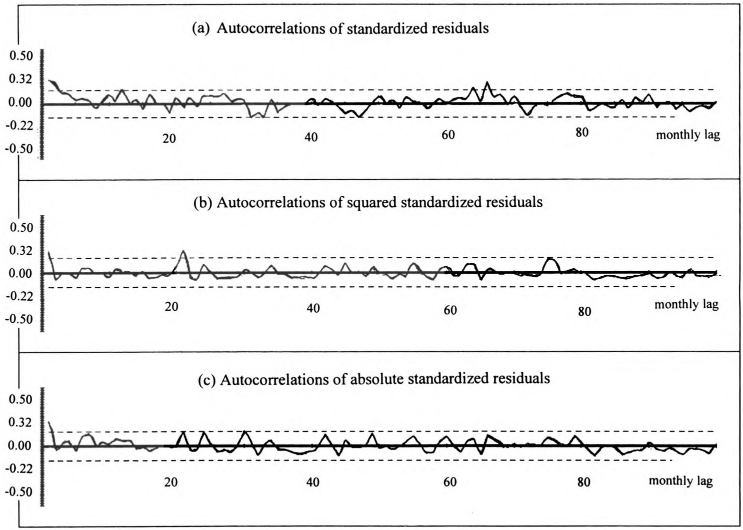


Figure 22. Correlograms of standardized residuals for interest rate, Thailand

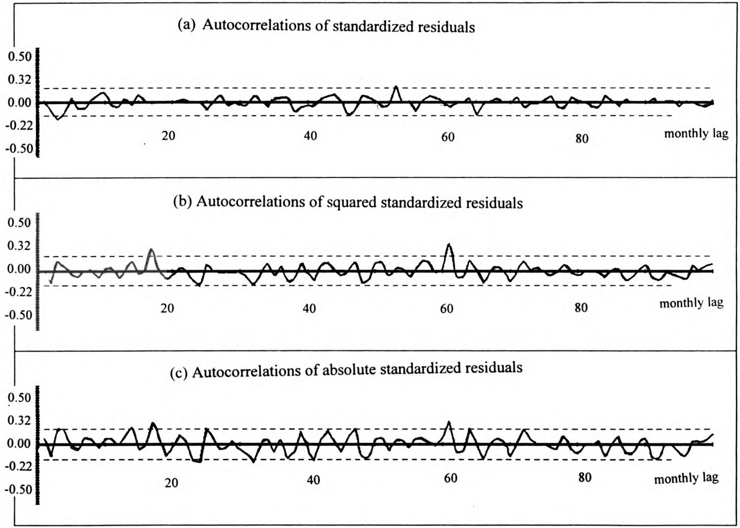


Figure 23. Correlograms of standardized residuals for interest rate, U.S.

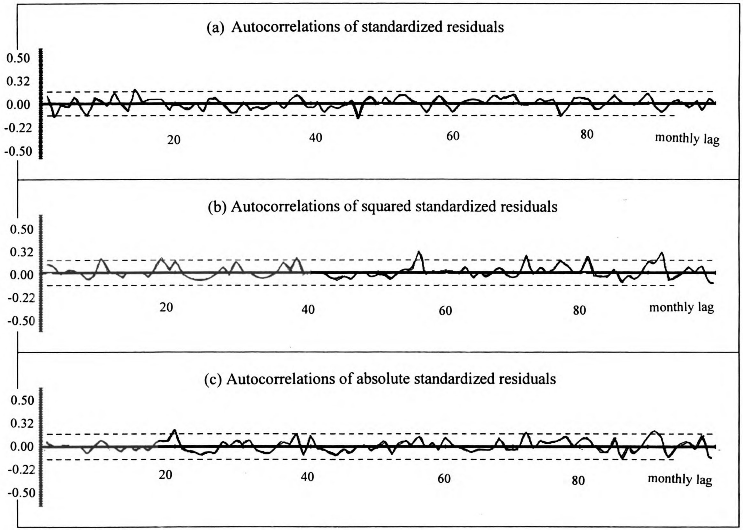


Figure 24. Correlograms of standardized residuals for real GDP, Indonesia

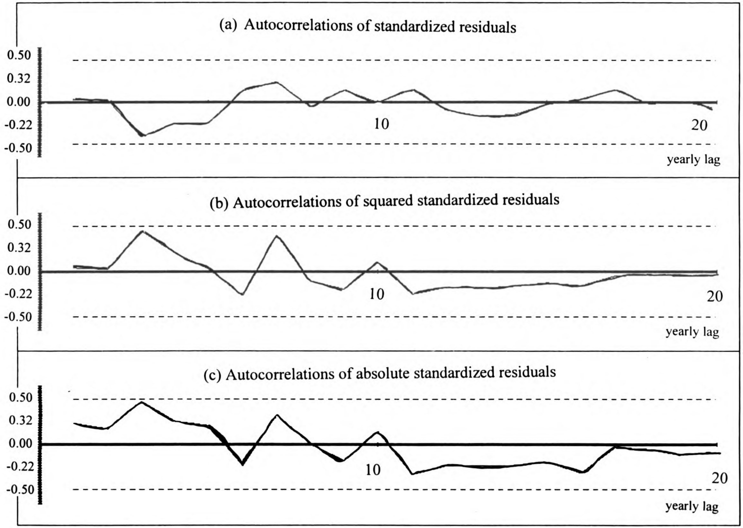


Figure 25. Correlograms of standardized residuals for real GDP, South Korea

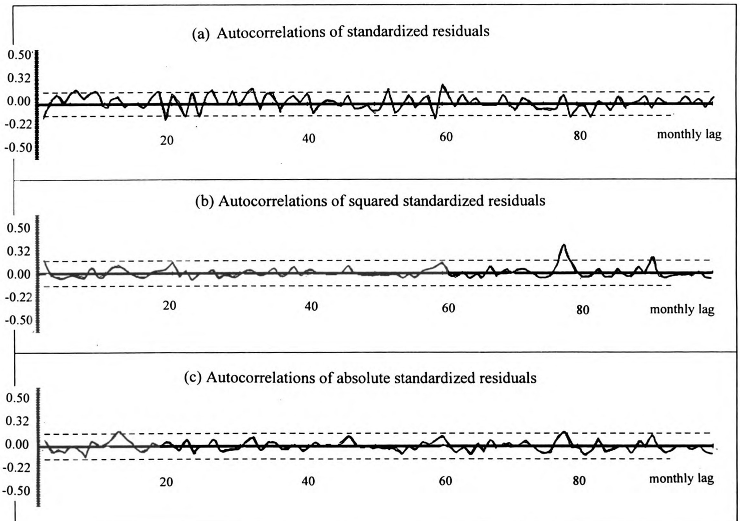


Figure 26. Correlograms of standardized residuals for real GDP, Malaysia

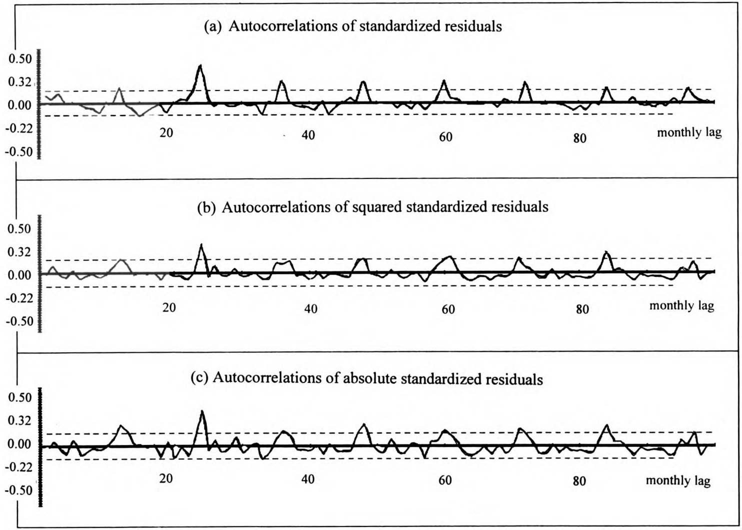


Figure 27. Correlograms of standardized residuals for real GDP, Philippines

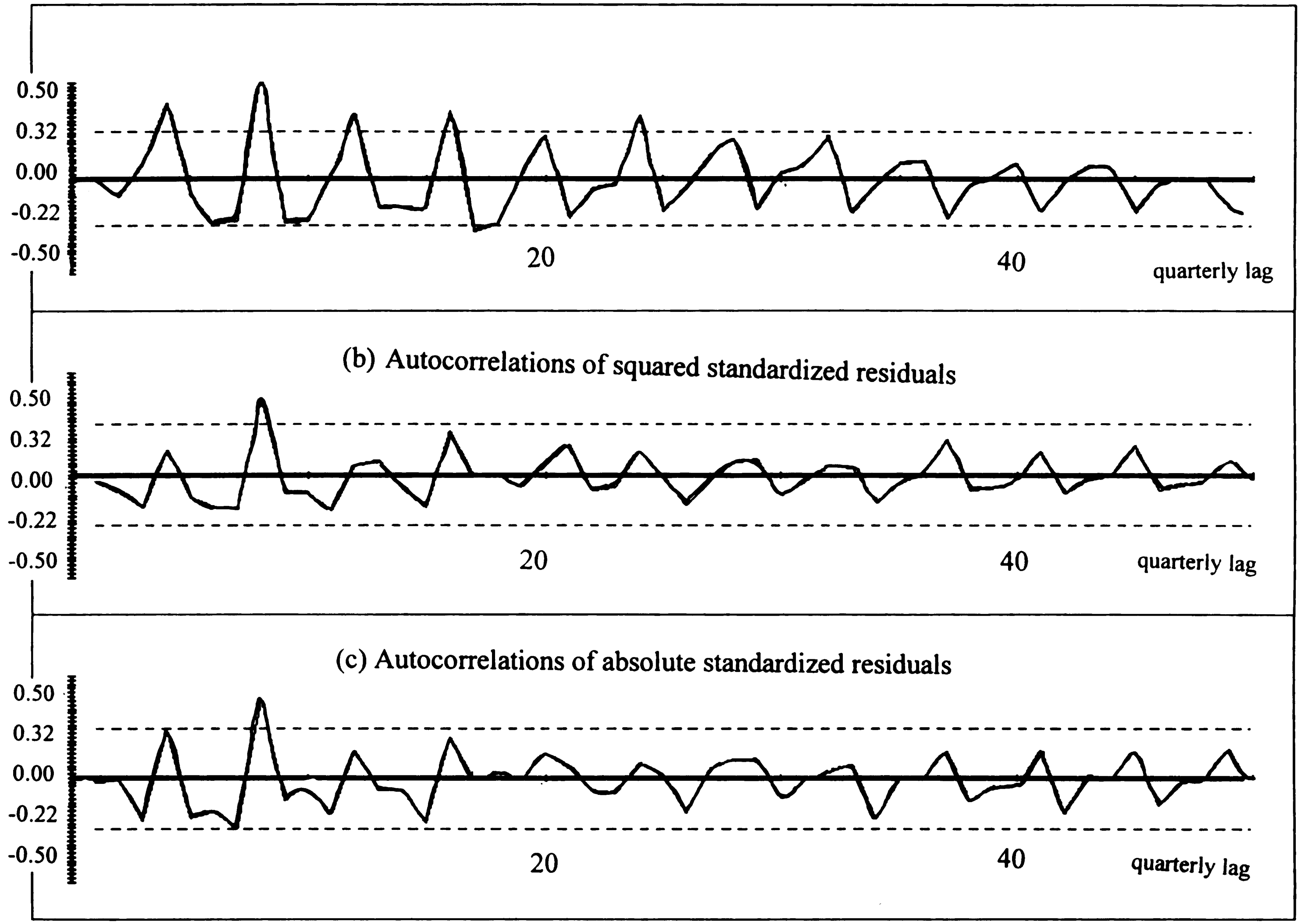
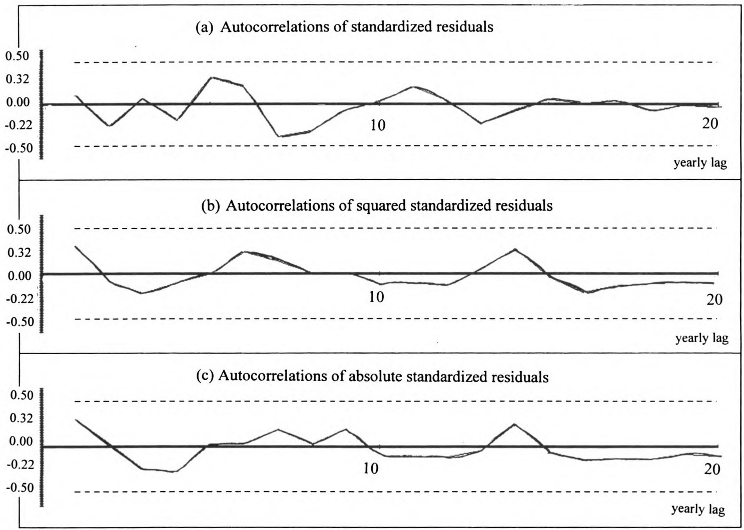


Figure 28. Correlograms of standardized residuals for real GDP, Thailand



Appendix 1. Previous studies of currency crises based on time series data

Author and Year Published	Countries	Time Period	Data Frequency	Comments
Balanco and Garber (1986)	Mexico	1973-1982	quarterly	Main focus is on the one-step-ahead probability of devaluation. The expected next-period exchange rate conditional on the devaluation also is constructed
Cumby and Van Wijnbergen (1989)	Argentina	1979-1980	monthly	One-month-ahead probability of a collapse of the crawling peg is yielded.
Goldberg (1994)	Mexico	1981-1986	monthly	The emphasis is on explaining the forces contributing to speculative attacks on the Mexican Peso.
Otker and Pazarbasioglu (1996)	Mexico	1982-1994	monthly	A main role of deterioration in fundamentals in the collapse is clarified by a Probit model.
Otker and Pazarbasioglu (1997b)	European countries	1979-1995	monthly	This study shows that the weak fundamentals are not the only cause in the crisis.

CHAPTER V

ESTIMATES OF REAL MONEY DEMAND FUNCTIONS

1. Introduction

Well-known structural analyses of currency crisis, Blanco and Garber (1986), Cumby and Van Wijnbergen (1989), and Goldberg (1994), estimate a real money demand function without taking into consideration the non-stationarity of variables. Therefore, their analyses display potentially a spurious regression problem and the conventional *t*-ratio and *F* significance tests can not be applied.

However, cointegration and error correction techniques in modeling of real money demand (see Granger, 1986; Engle and Granger, 1987; Hendry 1986; Johansen and Juselius, 1990) help to avert spurious regression and correct for the non-standard limiting distribution of coefficient. Furthermore, it provides a more robust means of estimating long run relationships and short run dynamics among the set of macroeconomic variables of interest.

This chapter examines the existence of a stable long-run real money demand function in South Korea and Malaysia for which high frequency data are available using the cointegration and error correction model. In addition, a short-run relationship among real money, real income, interest rate and interest rate differential is tested. The results of the study in this chapter suggest that both long and short-run models can be specified in South Korea and in Malaysia.

The regression specification to estimate real money demand includes a constant term, and the coefficients of real income, interest rate and the interest rate differential. The estimated constant term, $\hat{a}_0$, and coefficients, $\hat{a}_1$, $\hat{a}_2$ and $\hat{a}_3$ will be used in Chapter VI to obtain the shadow exchange rate,

$$\widetilde{s}_t = m_t - \hat{a}_0 + \hat{a}_1 i_t^* - \hat{a}_2 y_t - p_t^* - u_t + (\hat{a}_1 + \hat{a}_3)(\mu_t + \rho_t)$$ [19].

The following sections present a theoretical framework and an empirical model of real money demand. Then, univariate unit-root tests are presented to detect the nonstationarity of variables. Finally, an error-correction model is applied to estimate long and short-run dynamics in the real money demand function.

2.Theoretical framework

There is a diverse spectrum of money demand theories emphasizing the transactions, speculative, precautionary or utility considerations. These theories implicitly address a broad range of hypotheses. One significant aspect, however, is that they share common important elements (variables) among almost all of them. In general, they bring forth relationship between the quantity of money demanded and a set of few important economic variables linking money to the real sector of the economy (see Judd and Scadding, 1982).

The general specification begins with the following functional relationship for the long-run demand for money:

---

[19] See Chapter III for the details.

$$\left(\frac{M}{P}\right)^d = f(y, i),\ f_y \rangle 0,\ f_i \langle 0 \qquad (1)$$

where the demand for real money balances, M / P, is a function of a scale variable, such as real income y, and an opportunity cost variable, interest rate i. Except for interest rate, all variables have been natural log transformed in the estimation process. Using the real money balance as the dependent variable means that price homogeneity is explicitly imposed into the model. In addition, there are less severe econometric problems associated with using real rather than nominal balances as the dependent variable (see Boughton, 1891, and Johansen, 1992).

3. Empirical model

In general, the empirical work begins with a typical formulation of a simple theoretical money demand function relating demand for real money balances, m, to a measure of transactions or scale variable, y, and the opportunity cost of holding money, i. Empirical formulations also incorporate lagged dependent variable to capture the short-run dynamics.

3.1 Error-correction models

Error correction models (ECMs) are one of the most successful tools in applied money demand research. This is a dynamic error-correction representation where the long-run equilibrium relationship between money and its determinants is embedded in an equation that captures short-run dynamics. A specific model is

$$\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-1} + \mu_0 + \eta_t \qquad (2)$$

where x_t is a p-dimensional vector of I(1) variables, $\eta_1, \cdots, \eta_T$ are $IIN_p(0, \Lambda)$ and $x_{-k+1} \cdots x_0$ are fixed. The Π matrix conveys the long-run information in the data. The impetus for this type of model came from the findings that appropriate consideration needs to be given not just in the selection of an appropriate theoretical set up and empirical make up, but also in the specification of the proper dynamic structure of the model. Hence, economic theory should be allowed to specify the long-term equilibrium while the short-term dynamics can be defined from the data.

4. Application of ECM to the estimation of real money demand

4.1 Data set

The interest in the demand for money of developing countries has increased in recent years. This interest was triggered primarily by the concern among central banks and researchers about the impact of movement toward a flexible exchange rate regime, globalization of capital markets, and ongoing domestic financial liberalization and innovation. South Korea and Malaysia, both upper-middle-income developing countries, were selected for the estimation of the real money demand function because their foreign exchange rate policy toward a financial liberalization continued from the 1980s to the 1990s. Then, after the abrupt devaluation in 1997, South Korea and Malaysia chose opposite foreign exchange regimes. While South Korea allowed a free foreign exchange rate regime, Malaysia returned to the fixed exchange rate system. Hence, the shadow exchange rates and the probabilities of collapse obtained later with the estimated real money demand function in this chapter will provide an opportunity to compare the economic conditions after the crisis under two very different exchange rate regimes. Furthermore, the two countries' data, also, have a higher frequency than the other countries' that experienced the Asian currency crisis in 1997-98. This will increase the power of statistical tests and, in turn, raise the confidence one can place in the empirical results.

Monthly data from 1970:01 through 1996:12 for South Korea and Malaysia are used for the estimation. All of the variables, $m2$, y, p, i, and i^* are taken from the CD-ROM version of the International Monetary Fund's *International Financial Statistics* (IFS) and defined in Table 17. Real money balance, *rm2 (m2-p)*, and y are seasonally adjusted. In particular, data not seasonally adjusted is preferable for the unit roots tests and cointegration analysis since the seasonal-adjustment filters have adverse effects on the power of the unit root and cointegration tests according to Ghysels (1990) and Davidson and MacKinnon (1993). However, seasonally adjusted data are used in this analysis due to the availability of data[20]. Despite the loss of power from the seasonal adjustment filters, the use of seasonally adjusted data removes the need to test for the order of seasonal integration. In addition, it removes the need to add seasonal dummy variables in the real money balance equation. Figures 29 to 36 present graphical descriptions of the variables used in the analysis.

4.2 Unit-root tests

4.2.1 The Dickey-Fuller unit-root tests

A univariate test for unit roots was first advocated by Fuller (1976) and Dickey and Fuller (1981). The Dickey-Fuller based approach basically involves the running of the following univariate regression

$$x_t = \alpha x_{t-1} + \sum_{j=1}^{k} \alpha_j \Delta x_{t-j} + [set\ of\ fixed\ regressors] + \varepsilon_t \quad (3)$$

where the set of fixed regressors include a constant and a linear trend. We consider two alternatives, with an intercept and with an intercept and a trend. Since the presence of autocorrelation destroys the properties of the test, it is important to make the correct augmentation to be able to interpret test results. If the number of lags is above or below the number necessary to render the error as white noise, it biases the test's power and size. Too many lags reduce the power of the test. Too few lags distort the test size. The number of lags in equation (3) is selected such that the regression yields non-serially correlated errors.

If x_t is non-stationary, α will assume a unit value and x_t likely has a unit root. The null hypothesis that $\alpha = 1$ can be tested by reference to its usual t-statistic based on $\hat{\alpha}$, the OLS estimator. This statistic is referred to as the Augmented Dickey-Fuller (ADF) statistic. However, the distribution of ADF does not follow the usual student's t distribution. Approximate critical values of this statistic were originally given in Fuller (1976). In addition, the normalized bias test statistic, $Tc(\hat{\alpha} - 1)$, where T is the number of

[20] Only seasonally adjusted real income for South Korea can be collected in IFS.

the observations and $c = (1 - \sum_{j=1}^{k} \hat{\alpha}_j)^{-1}$, can be employed to test the non-stationarity of the variable.

The results in Table 23 strongly indicate that there is a unit root in *rm2*, *y*, *i*, and *if* ($i - i^*$) in South Korea and *rm2* and *y* in Malaysia since the null hypothesis cannot be rejected at the 5% level. However, the null hypothesis is rejected at the 5% level by the normalized bias tests $Tc(\hat{\alpha} - 1)$ for *i* and *if* in Malaysia and, at the 10% level, for *if* in South Korea. In addition, the results, as summarized in Table 24, unambiguously reveal that *rm2*, *y*, *i*, and *if* in South Korea and *rm2* in Malaysia contain a unit root. This is because none of the tests can reject the null hypothesis of a unit root in the regression with an intercept and a linear trend. The null is rejected at the 5% level by the normalized bias tests for *y*, *i*, and *if* in Malaysia.

4.2.2 The Phillips-Perron unit-root tests

Phillips and Perron (1988) propose a nonparametric method of controlling for higher-order serial correlation in a series. The test regression for the Phillips-Perron (PP) test is the AR(1) process:

$$x_t = \alpha x_{t-1} + [set\ of\ fixed\ regressors] + \varepsilon_t \quad (4)$$

While the Dickey-Fuller test corrects for higher order serial correlation by adding lagged differenced terms on the right-hand side, the Phillips-Perron test makes a correction to the *t*-statistic of the coefficient from the AR(1) regression to account for the serial correlation in ε. The PP *t*-statistic is computed as

$$Z_t = \frac{\hat{\gamma}_0^{1/2} t_a}{\hat{\lambda}} - \frac{(\hat{\lambda}^2 - \hat{\gamma}_0) T \hat{\sigma}_a}{2\hat{\lambda} s} \quad (5)$$

where $\hat{\lambda}^2$ is a Newey-West estimator, t_a and $\hat{\sigma}_a$ are the t-statistic and estimated standard error of α , and γ_j and s^2 are the jth autocovariance and the estimated variance of ε. The correction is nonparametric since we use an estimate of the spectrum of ε at frequency zero that is robust to heteroskedasticity and autocorrelation of unknown form. The Newey-West heteroskedasticity autocorrelation consistent estimate

$$\hat{\lambda}^2 = \hat{\gamma}_0 + 2\sum_{j=1}^{k} [1 - j/(k+1)]\hat{\gamma}_j \quad (6)$$

where k is the truncation lag. The asymptotic distribution of the PP t-statistic is the same as the Dickey-Fuller t-statistic. With the PP t-statistic, Z_t, the Phillips-Perron ρ statistic, Z_ρ, also can be used for the unit root test. The Z_ρ is calculated as

$$Z_\rho = T(\hat{a} - 1) - \frac{(\hat{\lambda}^2 - \hat{\gamma}_0) T^2 \hat{\sigma}_a^2}{2s} \quad (7)$$

The results of applying the PP test procedure to the variables in the South Korean and Malaysian demand functions for real balances are presented in Table 25. While the results of a regression with an intercept indicate that there is a unit root in *rm2*, *i*, and *if* in South Korea and *rm2* and *y* in Malaysia since the null hypothesis could not be rejected at the 5% level, the null is rejected at the 10% level by Z_t for *y* in South Korea and at the 5% level by the Z_t and Z_ρ for *i* and *if*, respectively, in Malaysia. In addition, the Z_t and Z_ρ tests with an intercept and trend show that *rm2*, *y*, *i*, and *if* in South Korea and *rm2* and if in Malaysia contain a unit root. The null hypothesis of a unit root is rejected at the 5% level by the Z_t and Z_ρ tests for y and i in Malaysia.

4.2.3 The KPSS unit-root tests

Due to the well known low power of the standard unit-root tests from the empirical evidence, Nelson and Plosser (1982), Kwiatkowski, Phillips, Schmidt, and Shin (1992) proposed a test of the null hypothesis that an observable series is stationary around a deterministic trend. They assume that a series can be decomposed into the sum of a deterministic trend, a random walk, and a stationary error:

$$y_t = \xi t + r_t + \varepsilon_t . \quad (8)$$

Here r_t is a random walk:

$$r_t = r_{t-1} + u_t , \qquad (9)$$

where the u_t are iid (0, σ_u^2); the initial value r_0 is fixed as the intercept. In this setting, The stationary hypothesis is simply $\sigma_u^2 = 0$. Then, since ε_t is assumed to be stationary, under the null hypothesis y_t is trend-stationary. They also considered the special case of the model (8) where $\xi = 0$, in which case under the null hypothesis y_t is stationary around a level rather than around a trend.

The KPSS LM(and LBI) statistic is defined as:

$$\hat{\eta}_\mu = T^{-2} \sum_{t=1}^{T} S_t^2 / s^2(l) , \quad (10)$$

where $S_t = \sum_{i=1}^{t} e_i^2, t = 1,2,\ldots,T$ and $s^2(l) = T^{-1}\sum_{t=1}^{T} e_t^2 + 2T^{-1}\sum_{s=1}^{l} w(s,l) \sum_{t=s+1}^{T} e_t e_{t-s}$. Here $w(s, l)$ is the Bartlett window which is the same as the Newey-West heteroskedasticity autocorrelation consistent estimate in the PP t-statistic.

In the study we consider two null hypotheses: one is the level stationary hypothesis, the other is a trend stationarity. The resulting test statistics are denoted $\hat{\eta}_\mu$, and $\hat{\eta}_\tau$ respectively. For each test, we consider values of the lag truncation parameter, l, from 2 to 16. Since the data series are highly dependent over time and the residuals from the regressions are serially correlated, it is not realistic to assume iid errors under the null and use $l = 0$, no correction for autocorrelation, in estimation of the long-run variance. The choices of 15 and 16 for l follow the values of l as a function of T : $l = integer$ $[12(T/100)^{1/4}]$. This is the rule suggested by Schwert (1989) to sufficiently correct for the autocorrelation problems in the residuals.

The test results are provided in Tables 26 and 27. First we consider the null hypothesis of stationarity around a level. The null hypothesis of level stationarity is rejected at a 5% level for all series except *if* for South Korea. This is the case regardless of the value of l chosen. This is shown in Table 26 for both South Korea and Malaysia. The rejection of the null is not surprising for the real M2 and GDP, since obvious deterministic trends are present (see Figure 29, 30, 33, and 34). For the *if* in South Korea, the test result cannot reject the null hypothesis at a 10% level when the value of 15 is chosen for l. Then, when the trend stationary hypothesis is tested, as reported in Table 27, for all of the series apart from *rm2* in South Korea, the null is rejected at a 5% or a 10%

level for all values of l. The outcome for the *rm2* in South Korea depends on the lag truncation parameter, l.

4.3 Residual based cointegration tests

Engle and Granger (1987) suggest that the residuals from an OLS estimation of the cointegrating regression can be examined for the presence of a unit root in the autoregressive representation. If there is no cointegration, there should be a unit root in the residuals.

Let the observed data X_t be a px1 dimensional time series, partitioned as $X_t = (x_{1t}, x'_{2t})$, where x_{1t} is a scalar and x_{2t} is an m-vector, $p = m+1$, and each element of X_t is known to be I(1). By regressing one of the variables, say x_{1t}, on the others with ordinary least squares, we obtain the cointegrating regression:

$$x_{1t} = \hat{\alpha}' X_{2t} + \hat{u}_t \quad (11)$$

where X_{2t} may also contain a constant or time trend, other than x_{2t}; $\hat{u}_t$ are the residuals. Our null hypothesis of no cointegration then corresponds to the null hypothesis that $\hat{u}_t$ is I(1) where

$$\hat{u}_t = \hat{\rho}\hat{u}_{t-1} + \hat{\varepsilon}_t . \quad (12)$$

While a simple Dickey-Fuller test can be used in this model, instead we consider Z_ρ and Z_t tests, since these test statistics have the advantage that they correct for both potential serial correlation and heteroskedasticity in the cointegrating errors.

Following Hansen (1992), we consider two procedures. The first is to run an unrestricted OLS regression with a time trend included to test for the existence of cointegration in a series with a drift. This procedure is equivalent to detrending the series first:

$$x_{1t} = \hat{\mu} + \hat{\alpha}' x_{2t} + \hat{\beta} t + \hat{u}_t \qquad (13)$$

The inclusion of time trends in the regression has the advantage of rendering estimates of the cointegrating vector invariant to the presence of trends in the regressors. This also simplifies the asymptotic theory as shown in Phillips and Hansen (1990) and Hansen (1992). The second approach is to estimate an OLS regression without a time trend.

$$x_{1t} = \tilde{\mu} + \tilde{\alpha}' x_{2t} + \tilde{u}_t \qquad (14)$$

As Hansen (1992) notes, cointegration tests without a time trend are generally more powerful than cointegration tests with a time trend.

The results of unit root tests identiry *rm2*, *y*, *i*, and *if* in South Korea and *rm2* in Malaysia as I(1) processes. However, the unit root tests do not indicate any non-stationarity of *y*, *i*, and *if* in Malaysia. Therefore, while the results of the Z_ρ and Z_t tests on the residuals of the cointegrating regression in South Korea are valid for detecting cointegration, their justification in Malaysia is suspect. In spite of this, the estimated coefficients in the OLS estimation implemented that were used for the residual-based tests will be applied for obtaining the shadow exchange rates and the probabilities of collapse using the same framework defined by Blanco and Garber (1986), Cumby and Van Wijnbergen (1989), and Goldberg (1994).

Since residual-based cointegration tests are developed from single-equation regression models, they depend on an arbitrary normalization of the cointegrating regression. As far as the demand for money function is concerned, the long-run money demand relation with no structural breaks may be written as

$$rm2_t^d = a_0 + a_1 y_t + a_2 i_t + a_3 if_t + u_t \quad (15)$$

where $rm2_t^d = m_t^d - p_t$ and $if_t = i_t - i_t^*$. We consider both cases (13), which includes time trend, and (14), which does not. The results are as follows.

In Table 28, there are the results of both OLS estimation and residual based cointegration tests. The null hypothesis of no cointegration can be rejected by Z_ρ and Z_t tests in South Korea regardless of the inclusion of a deterministic trend. In particular, the tests show that there is a long-run cointegrating relationship between the variables in the demand for real M2 equation. From the OLS estimates, we notice that the sign of income, *y*, is positive as we expected, but the signs of *if* in South Korea and *i* in Malaysia are the opposite of what we anticipated. However, since the money market interest rate is used for *i*, we cannot argue that the sign of the domestic interest rate and the gap between the domestic and foreign interest rate needs to be negative. In particular, the money market rate is a representative interest rate reflecting various interest rates in the financial market. Thus, it is probable that the negative effect on real money demand from an alternative asset interest rate, such as the treasury bill rate, is not strongly reflected on the sign.

Due to the uncertainty about the nonstationarity of the variables, the results of the Z_ρ and Z_t tests in Malaysia are not reported in Table 28. Despite the uncertainty, however, the t-statistics and adjusted R-squares are exceptionally high which are common symptoms of a spurious regression.

4.4 Johansen's full information maximum likelihood estimation

Even though the residual-based cointegration tests indicate a cointegration between the variables, we cannot determine from those tests whether there are other linearly independent cointegrating vectors in the system. Approaches other than residual-based tests for cointegration are available such as the likelihood ratio tests of cointegration rank of Johansen (1988, 1991), Johansen and Juselius (1990), and a common stochastic trends test proposed by Stock and Watson (1988). These tests are developed from system methods designed to help researchers avoid invalid restriction from arbitrary normalization in cointegrating regression of residual-based tests. As an alternative, Johansen's (1988, 1991) maximum likelihood methods for the analysis of cointegration can simultaneously detect the number of the cointegration vectors in the system, estimate and test for linear hypothesis about the cointegrating vectors and their adjustment coefficients. Based on these advantages, we will apply this technique to continue our study.

To begin with, the following model without a time trend is fitted to the demand for real *M2* data.

$$H_2 : \Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-1} + u_0 + \eta_t \qquad (16)$$

In addition, a subsequent model with a linear trend in the cointegrating relations is estimated to check the significance of linear trend in the estimation. That is, under the null we estimate

$$H_2^*: \Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \alpha(\beta', \beta_1)(x'_{t-1}, t)' + u_0 + \eta_t. \quad (17)$$

The lag length k is chosen to be the minimum length for which there is no significant autocorrelation in the estimated VECM residuals using the Ljung-Box Q statistics (1979). The misspecification tests for the normal iid assumption for the residuals in the model are reported. The normality assumption is tested by the Jarque and Bera statistic (Jarque and Bera, 1980).

4.4.1 Misspecification tests

The misspecification tests for the model are provided in Table 29 and 30. In Table 29, while the p-value of the Q statistics shows no autocorrelations in the residuals, the excess skewness and kurtosis in the residuals of y, i, and if cause the Jarque-Bera test statistic to become significant. However, the deviations from normality are not a serious problem. So long as the cumulative sums of errors converge to a Brownian motion, the asymptotic analysis is the same as that given under the assumption of normality (Johansen, 1991, Johansen and Juselius, 1992, and Gonzalo, 1994).

The results of misspecification tests in Table 30 also do not specify any evident autocorrelation except for y. Nevertheless, the Jarque-Bera test statistics are significant for y, i, and if.

4.4.2 Testing for reduced rank and normalized cointegrating vector

The cointegration tests results are provided in Tables 31 and 32. In Table 31, the trace and λ_{max} tests fail to reject $r = 0$ at either the 1% or 5% level when the model is regressed with a drift. However, both the trace and λ_{max} tests in the model with a deterministic time trend can reject $r = 0$ at the 5% level. This indicates that whether there is a time trend or not is important to establish the conclusion of cointegration with stable coefficients in South Korea. Alternatively, in Malaysia, the trace test rejects $r = 0$ at the 5% level both in the models with and without the time trend. Hence, the existence of a time trend in Malaysia is relatively less important compared to South Korea.

The normalized cointegrating vector and the error correction coefficients reported in Table 33 and 34 offer a long run relationship between *rm2*, *y*, *i*, and *if*. However, the parameters α and β are not identified. This is because, given any choice of the matrix $\varsigma(r \times r)$, $\alpha\varsigma$ and $\beta(\varsigma')^{-1}$ also produces the same matrix Π. The data only identify the space spanned by the columns in β, and the space spanned by α.

As shown in Table 33, the signs of coefficients on income and the interest rate differential in the model without a deterministic trend are not consistent with the prediction of the theories. They also are not consistent with the result of the residual-based cointegration test in the model with a drift. However, since no cointegration relationship is detected in the model for South Korea, no significance can be attached to the directions of coefficients on both variables. As such, the signs of the coefficients on income and the interest rate in the model with a deterministic trend do agree with the expectations from the theories. Nevertheless, the sign of the interest rate differential, *if*, representing an impetus of currency substitution caused by an expected devaluation and a risk premium, does against theoretical expectations. However, since the money market interest rate is used for *i*, we cannot presume a negative relationship. While the cointegration tests detect a cointegrating relationship only in the model with time trend for South Korea, they indicate a cointegrating relationship both in the models with or without a time trend for Malaysia. Therefore, we need to consider both models with and without the deterministic time trend in Table 34. As for the Table 34, the signs of coefficients on income and the interest rate differential coincide with what we expected in both models with or without a time trend, but the positive signs of domestic interest rates in both models indicate the traits of money market rate representing various interest rates.

In general, the α matrix should contain the weights used to enter the cointegrating vectors into the system. Each nonzero column of the α matrix also measures the speed of the short-run response to disequilibrium in the equations of endogenous variables. For example, the coefficients of α in Table 33 measure the feedback effects of the lagged disequilibrium in the cointegrating vector onto the variables in the vector autoregression (VAR). In particular, the absolute value of the first term in α represents the speed at which *rm2*, the dependent variable in the first equation of the VAR, moves toward restoring the long-run equilibrium. We can see that equilibrium errors cause *i* and *if* to adjust more rapidly than *rm2* and *y* in the Table 33 and Table 34. This suggests that the adjustments of the domestic interest rate and the difference between the domestic and foreign interest rate are crucial to the cointegrating relation.

4.4.4 Weak exogeneity tests

The weak exogeneity tests permit one to draw inferences from the cointegration relationship that is to examine whether the short-run demand for money could be modeled in a simpler setting. Let observed data X_t be a px1 dimensional time series, partitioned as $X_t = (x'_{1t}, x'_{2t})$, where x_{1t} and x_{2t} are m- and n-vectors, respectively; p = m+n. The variable x_{2t} is said to be weakly exogenous for α and β, if the conditional distribution of Δx_{1t}, given Δx_{2t} as well as the lagged values of X_t and ΔX_t, contains the parameters α and β, whereas the marginal distribution of Δx_{2t}, given the lagged values of X_t and ΔX_t, does not contain the parameters α and β. In particular, the parameters in the conditional and marginal distribution must be variation-free or, in other words, they cannot have any joint restrictions. These conditions are taken from Johansen and Juselius (1990, 1992) and Johansen (1991).

Since one cointegrating relationship has been identified in the cointegration test with the time trend in South Korea, and with or without the time trend in Malaysia, the weak exogeneity tests are evaluated under the assumption of rank (r) = 1. The test statistics will be asymptotically distributed as $\chi^2(1)$ if a given variable for the cointegrating vector is weakly exogeneous. Here, the null hypothesis is the existence of weak exogeneity. This is usually examined by the restriction of a particular α to zero. When the null hypothesis is not rejected, disequilibrium in the cointegrating relationship does not have a feedback on the variable of interest. However, any disequilibrium of a given variable will still impact the cointegrating relationship.

Table 35 shows that weak exogeneity is rejected for *rm2* and *y* at the 5% significance level and for *rm2* and *i* at the same level respectively in the models with a drift and with a time trend. Therefore, a short-run model can be designed with a system of two equations, one with *rm2* and another with y or i by considering other variables as weakly exogenous. However, since the cointegration relationship represents the demand for money, an alternative single equation framework can be estimated for the short-run model with $\Delta rm2$ as the endogenous variable in spite of a loss of efficiency. Besides, the results in Table 36 indicate that the weak exogeneity is rejected for only y at 5% significance level in the model with drift. Nonetheless, it is rejected for *rm2* and y at 10% and 5% respectively in the model with a time trend. Given the results in Table 36, the short-run model involving $\Delta rm2$ as the endogenous variable in the single equation framework can be used in the model with a time trend as well.

4.4.5 Stability of long-run parameters

So as to ensure the robustness of estimation parameters, they are evaluated for their stability throughout the sample period. To accomplish this task, a VECM with drift is estimated using the recursive estimation method beginning in early 1985. From 1990 onward, a series of deregulatory measures was put in place to meet the increasing need for liberalization to improve the efficiency of domestic financial markets and to respond effectively to the rapid changes in international financial markets in South Korea. Similarly, Malaysia continued its policy to liberalize interest rates first set out in 1978 and continued during the years of the sample. Because of the time frames for policy changes, the initial point of early 1985 in the recursive estimation should still leave enough data points from which to examine whether the demand for the real *M2* has remained stable over time.

Figures 37 to 39 provide evidence about the stability of parameters of real GDP, domestic interest rate, and the interest rate differential in South Korea. As expected from the cointegration tests, the graphical evidence indicates instability in the long-run parameters during this period. While the parameters are particularly unstable in South Korea, Figures 40 to 42 present weaker evidence of parameter instability in Malaysia. The elasticity of real GDP is fairly stable and close to unity throughout the period for Malaysia. Also, other parameters, such as the coefficients on the interest rate and the interest rate differential, exhibit notable constancy. This is an assuring result and it is expected given that the cointegration test found a cointegration relationship at the 5% critical level.

4.4.6 Short-run model

The short-run model provides information about how the adjustments take place among various variables to restore the long-run equilibrium in response to short-term disturbances in the demand for money. Essentially it is an ECM with an error-correction (EC) term to control for the existence of a long-run relationship. In general, short-run models have the I(0) representation of the variables both on the left-hand side and the right-hand side of the equation. Since the variables are assumed to be either I(0) or I(1), the right-hand side will consist of the first differences of the relevant variables with the exception being the inclusion of level variables in the EC term.

Based on the weak exogeneity tests, a single equation reduced form model such as equation (18) is sufficient to analyze the short-run dynamics for $\Delta rm2$.

$$\Delta rm2_t = u_0 + \alpha(L)\Delta i_{t-1} + \beta(L)\Delta y_{t-1} + \gamma(L)\Delta if_{t-1} + \alpha_1(\beta',\beta_1)(x'_{t-1},t)' + \eta_t \quad (18)$$

The right-hand side of the equation (18) includes an EC term, which is $\alpha_1(\beta',\beta_1)(x'_{t-1},t)'$, calculated as *rm2* minus the estimated *rm2* in time *t-1*. In economics, it represents excess money in the previous period. Since all variables are I(0), the above model can be estimated by OLS. The results of the estimation are shown in Table 37 and 38. The error-correction term, α_1, is negative in both South Korea and Malaysia. This validates the significance of the cointegration relationship. A significant negative EC term conveys two pieces of information: first, agents have corrected in the current period a proportion of the previous disequilibrium in money balances, Rose (1985); second, it assures us that the cointegration relationship established previously is valid by Granger's Representation Theorem, Engle and Granger (1987). The negative EC sign implies that a fall in excess money in the last period will increase the level of desired money holdings in the current period. In other words, any particular disequilibrium will be reduced over time. The results of diagnostic tests concerning autocorrelation and normality are already presented in Table 29 and 30.

5. Conclusion

This chapter has examined the empirical relationships between money and other macroeconomic variables in South Korea and Malaysia, using the residual-based cointegration tests based upon the results of the unit-root tests such as ADF-*t*, ADF normalized bias test, the PP Z_t and Z_ρ tests of a unit root against trend stationarity and the KPSS tests of trend stationarity and applying Johansen's procedure along with the VECM approach.

Whereas the residual-based cointegration tests indicate a stable cointegration relationship among the variables in the model regardless of the inclusion of a deterministic trend in South Korea, the Johansen's likelihood ratio tests of cointegration rank do not show any cointegration relationship in the model without a deterministic time trend. The result of the Johansen's likelihood tests imply that whether there is a time trend or not in South Korea is important for the conclusion of cointegration with stable coefficients. However, in Malaysia, Johansen's test detects a cointegration relationship in the model with or without a deterministic time trend. The graphical evidence of stability of parameters in the real money demand equation confirms that there is a stable cointegration relationship among the variables in Malaysia. However, the parameters are not stable in the model without a deterministic time trend in South Korea.

Based on the findings from the tests for weak exogeneity, a single equation reduced form model is formulated to analyze the short-run dynamics for $\Delta rm2$. Although a considerable number of variables turn out to be insignificant from the *t*-statistics in the estimation, the error-correction term's sign is negative for both South Korea and Malaysia, validating the significance of the cointegration relationship.

In conclusion, the results suggest that both long and short-run models can be specified for both South Korea and Malaysia. However, due to the results of the cointegration tests and the tests for the stability of the parameters in the model, we cannot determine whether the model without deterministic time trend can be used to explain real money demand in South Korea. The use of monthly data, previously not implemented in earlier work on South Korea, helps to discover an unstable long-run relationship in the real money demand (See Arize, 1994). Finally, the estimates of a constant term, and coefficients of real income, interest rate and interest rate differential were made in the procedure of testing the stable long and short-run relationships in this chapter. These estimated constant term, $\hat{a}_0$, and coefficients, $\hat{a}_1$, $\hat{a}_2$ and $\hat{a}_3$ will be used in Chapter VI to obtain the shadow exchange rate and the probability of collapse.

Table 23. Dickey-Fuller tests for unit roots: Model I

Model I: Regression with an intercept (South Korea)

Series	*Lags* (k)	$\hat{\alpha}$	*t-test*	$Tc(\hat{\alpha}-1)$	Q *p-value*
rm2	2	0.999	-0.920	-0.229	0.634
y	6	0.996	-2.445	-0.886	0.117
i	0	0.963	-2.163	-8.973	0.424
if	0	0.950	-2.594	-12.185*	0.242
Critical values					
5%			-2.88	-14.0	
10%			-2.57	-11.2	

Notes. An * indicates significance at 10% level.
(a) *rm2* is a real money demand.
(b) *rm2* is from 1970:01 through 1996:12; a total of T=324; *y* is from 1970:01 through 1996:12; a total of T=324; *i* is from 1976:08 through 1996:12; a total of T=245; *if* is from 1976:08 through 1996:12; a total of T=245.

Model I: Regression with an intercept (Malaysia)

Series	*Lags* (k)	$\hat{\alpha}$	*t*-test	$Tc(\hat{\alpha}-1)$	Q *p-value*
rm2	0	1.000	0.110	0.035	0.520
y	4	1.000	0.374	0.200	0.001
i	3	0.924	-2.440	-47.003**	0.053
if	3	0.947	-2.181	-27.545**	0.162
Critical values					
5%			-2.88	-14.0	
10%			-2.57	-11.2	

Notes. An ** indicates significance at 5% level.
(a) *rm2* is from 1970:01 through 1996:12; a total of T=324; *y* is from 1971:01 through 1996:12; a total of T=312; *i* is from 1970:01 through 1996:12; a total of T=324; *if* is from 1970:01 through 1996:12; a total of T=324.

Table 24. Dickey-Fuller tests for unit roots: Model II

Model II: Regression with an intercept and a linear trend (South Korea)

Series	*Lags* (k)	$\hat{a}$	*t-test*	$Tc(\hat{a}-1)$	Q *p-value*
rm2	2	0.958	-3.110	-10.191	0.449
y	6	0.985	-1.304	-7.328	0.116
i	0	0.952	-2.422	-11.750	0.439
if	0	0.951	-2.546	-12.039	0.237
Critical values					
5%			-3.43	-21.4	
10%			-3.13	-18.0	

Notes. An * indicates significance at 10% level.
(a) *rm2* is from 1970:01 through 1996:12; a total of T=324; *y* is from 1970:01 through 1996:12; a total of T=324; *i* is from 1976:08 through 1996:12; a total of T=245; *if* is from 1976:08 through 1996:12; a total of T=245.

Model II: Regression with an intercept and a linear trend (Malaysia)

Series	*Lags* (k)	$\hat{a}$	*t-test*	$Tc(\hat{a}-1)$	Q *p-value*
rm2	0	0.989	-1.251	-3.558	0.494
y	2	0.902	-2.811	-52.570**	0.001
i	3	0.896	-2.861	-62.618**	0.056
if	3	0.927	-2.626	-37.021**	0.162
Critical values					
5%			-3.43	-21.4	
10%			-3.13	-18.0	

Notes. An ** indicates significance at 5% level.
(a) *rm2* is from 1970:01 through 1996:12; a total of T=324; *y* is from 1971:01 through 1996:12; a total of T=312; *i* is from 1970:01 through 1996:12; a total of T=324; *if* is from 1970:01 through 1996:12; a total of T=324.

Table 25. Phillips-Perron tests for unit roots

Model I: Regression with an intercept

Series	$\hat{\alpha}$		Z_t		Z_ρ	
	S. Korea	Malaysia	S. Korea	Malaysia	S. Korea	Malaysia
rm2	0.999	1.000	-0.823	0.112	-0.316	0.035
y	0.996	0.996	-2.827*	-0.564	-1.052	-0.443
i	0.963	0.820	-2.058	-5.132**	-8.077	-47.259**
if	0.950	0.902	-2.639	-3.526**	-12.607	-23.693**
Critical values						
5%			-2.88		-14.0	
10%			-2.57		-11.2	

Notes. An ** (*) indicates significance at 5%(10%) level.

(a) For South Korea, *rm2* is from 1970:01 through 1996:12; a total of *T*=324; *y* is from 1970:01 through 1996:12; a total of *T*=324; *i* is from 1976:08 through 1996:12; a total of *T*=245; *if* is from 1976:08 through 1996:12; a total of *T*=245.

(b) For Malaysia, *rm2* is from 1970:01 through 1996:12; a total of *T*=324; *y* is from 1971:01 through 1996:12; a total of *T*=312; *i* is from 1970:01 through 1996:12; a total of *T*=324; *if* is from 1970:01 through 1996:12; a total of *T*=324.

Model II: Regression with an intercept and a linear trend

Series	$\hat{\alpha}$		Z_t		Z_ρ	
	S. Korea	Malaysia	S. Korea	Malaysia	S. Korea	Malaysia
rm2	0.967	0.989	-2.891	-1.268	-15.612	-3.641
y	0.975	0.752	-1.775	-6.520**	-4.537	-73.239**
i	0.952	0.768	-2.330	-6.249**	-10.869	-68.931**
if	0.951	0.973	-2.582	-2.289	-12.408	-10.524
Critical values						
5%			-3.43		-21.4	
10%			-3.13		-18.0	

Notes. An ** indicates significance at 5% level.

(a) For South Korea and Malaysia, *rm2*, *y*, *i*, and *if* have the same sample size as the above model I.

Table 26. KPSS tests for stationarity : Model I

Model I: Regression with an intercept (South Korea)

Series	Lag truncation parameter (l) 2	4	5	6	8	15	16
	$\hat{\eta}_\mu$: 5% critical value is 0.463; 10% critical value is 0.347						
rm2	10.802**	6.527**	5.458**	4.695**	3.677**	2.119**	2.000**
y	10.616**	6.416**	5.365**	4.614**	3.613**	2.081**	1.965**
i	3.051**	1.875**	1.580**	1.369**	1.087**	0.650**	0.612**
if	1.351**	0.842**	0.715**	0.625**	0.503**	0.311	0.296

Notes. An ** (*) indicates significance at 5%(10%) level.
(a) *rm2* is from 1970:01 through 1996:12; a total of T=324; y is from 1970:01 through 1996:12; a total of T=324; i is from 1976:08 through 1996:12; a total of T=245; *if* is from 1976:08 through 1996:12; a total of T=245.

Model I: Regression with an intercept (Malaysia)

Series	Lag truncation parameter (l) 2	4	5	6	8	15	16
	$\hat{\eta}_\mu$: 5% critical value is 0.463; 10% critical value is 0.347						
rm2	10.652**	6.440**	5.387**	4.634**	3.631**	2.096**	1.978**
y	10.316**	6.239**	5.219**	4.490**	3.518**	2.032**	1.919**
i	3.355**	2.112**	1.790**	1.564**	1.263**	0.795**	0.759**
if	3.202**	1.998**	1.691**	1.473**	1.182**	0.732**	0.698**

Notes. An ** (*) indicates significance at 5%(10%) level.
(a) *rm2* is from 1970:01 through 1996:12; a total of T=324; y is from 1971:01 through 1996:12; a total of T=312; i is from 1970:01 through 1996:12; a total of T=324; *if* is from 1970:01 through 1996:12; a total of T=324.

Table 27. KPSS tests for stationarity: Model II

Model II: Regression with an intercept and a linear trend (South Korea)

Series	Lag truncation parameter (l) 2	4	5	6	8	15	16
	$\hat{\eta}_\tau$: 5% critical value is 0.146 10% critical value is 0.119						
rm2	0.277**	0.171**	0.146**	0.127*	0.103	0.069	0.066
y	2.005**	1.219**	1.022**	0.881**	0.693**	0.407**	0.386**
i	1.136**	0.704**	0.596**	0.519**	0.415**	0.253**	0.240**
if	1.357**	0.846**	0.718**	0.627**	0.505**	0.312**	0.297**

Notes. An ** (*) indicates significance at 5%(10%) level.
(a) *rm2* is from 1970:01 through 1996:12; a total of T=324; y is from 1970:01 through 1996:12; a total of T=324; i is from 1976:08 through 1996:12; a total of T=245; *if* is from 1976:08 through 1996:12; a total of T=245.

Model II: Regression with an intercept and a linear trend (Malaysia)

Series	Lag truncation parameter (l) 2	4	5	6	8	15	16
	$\hat{\eta}_\tau$: 5% critical value is 0.146 10% critical value is 0.119						
rm2	1.507**	0.914**	0.766**	0.660**	0.519**	0.304**	0.289**
y	1.497**	0.951**	0.810**	0.707**	0.570**	0.357**	0.341**
i	0.575**	0.367**	0.313**	0.275**	0.225**	0.148**	0.142*
if	0.651**	0.410**	0.348**	0.305**	0.247**	0.157**	0.151**

Notes. An ** (*) indicates significance at 5%(10%) level.
(a) *rm2* is from 1970:01 through 1996:12; a total of T=324; y is from 1971:01 through 1996:12; a total of T=312; i is from 1970:01 through 1996:12; a total of T=324; *if* is from 1970:01 through 1996:12; a total of T=324.

Table 28. Testing for no cointegration in demand for real M2

Model: $rm2_t^d = a_0 + a_1 y_t + a_2 i_t + a_3 if_t + [\beta trend] + u_t$								
	OLS estimates						Cointegration tests	
	a_0	y_t	i_t	if_t	$trend$	$\bar{R}^2$	Z_t	Z_ρ
South Korea	2.384 (0.047)	1.042 (0.009)	-0.007 (0.002)	0.012 (0.002)		0.991	-4.523**	-37.146**
	3.804 (0.077)	0.368 (0.035)	-0.007 (0.001)	0.010 (0.001)	0.006 (0.000)	0.997	-4.569**	-36.371*
Malaysia	1.924 (0.042)	1.157 (0.011)	0.041 (0.003)	-0.023 (0.002)		0.980	_	_
	4.601 (0.138)	0.021 (0.058)	0.024 (0.002)	-0.017 (0.002)	0.008 (0.000)	0.991	_	_
Critical Values								
5%							-4.16 (-4.49)	-32.2 (-37.7)
10%							-3.84 (-4.20)	-27.8 (-33.2)

Notes. An ** indicates significance at 5% level.
( ) means the critical value when the regression includes a deterministic trend.

Table 29. Residual misspecification tests (South Korea)

Model: $\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-1} + u_0 + \varepsilon_t$

Eq.	S.E.E[a]	Sk[b]	Ek	Q *P-value*	τ[c]
$\Delta rm2$	0.014	-0.117	3.563	0.848	3.704
Δy	0.029	-0.588	7.703	0.078	233.979**
Δi	1.068	0.328	4.988	0.756	43.634**
Δif	1.204	0.081	5.129	0.202	45.387**

Model: $\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \alpha(\beta', \beta_1)(x'_{t-1}, t)' + u_0 + \varepsilon_t$

Eq.	S.E.E[a]	Sk[b]	Ek	Q *P-value*	τ[c]
$\Delta rm2$	0.014	-0.015	3.069	0.909	0.056
Δy	0.029	-0.481	7.809	0.289	234.487**
Δi	1.048	0.304	4.193	0.905	17.498**
Δif	1.157	0.265	4.123	0.972	15.034**

Notes. An ** indicates significance at 5% level. k=5 for the first model and k=10 for the second model.

a. S.E.E denotes the standard error of regression estimate.

b. Sk and Ek are the skewness and kurtosis statistics.

c. The Jarque and Bera test for normality (Jarque and Bera, 1980),

$$\tau = \frac{T-m}{6}\left(Sk^2 + \frac{Ek^2}{4}\right) \sim \chi(2)$$ where m is the number of regressors.

Table 30. Residual misspecification tests (Malaysia)

Model: $\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-1} + u_0 + \varepsilon_t$

Eq.	S.E.E[a]	Sk[b]	Ek	Q *P-value*	τ[c]
$\Delta rm2$	0.014	-0.018	2.882	0.899	0.193
Δy	0.042	0.309	3.707	0.003	11.143**
Δi	1.186	0.289	8.416	0.999	374.529**
Δif	1.352	0.342	7.090	0.989	217.114**

Model: $\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \alpha(\beta', \beta_1)(x'_{t-1}, t)' + u_0 + \varepsilon_t$

Eq.	S.E.E[a]	Sk[b]	Ek	Q *P-value*	τ[c]
$\Delta rm2$	0.014	-0.014	2.870	0.898	0.222
Δy	0.042	0.204	3.496	0.003	5.196*
Δi	1.187	0.228	8.255	0.997	351.240**
Δif	1.352	0.354	7.032	0.990	211.566**

Notes. An *(**) indicates significance at 10% (5%) level. k=8 for all the models.

a. S.E.E denotes the standard error of regression estimate.

b. Sk and Ek are the skewness and kurtosis statistics.

c. The Jarque and Bera test for normality (Jarque and Bera, 1980),

$$\tau = \frac{T-m}{6}(Sk^2 + \frac{Ek^2}{4}) \sim \chi(2)$$ where m is the number of regressors.

Table 31. Test of the cointegration rank (South Korea)

Model: $\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-1} + u_0 + \varepsilon_t$

H_2	eigenvalue	trace	$\lambda_{\max}$
$r = 0$	0.097	43.641	24.470
$r \leq 1$	0.060	19.171	14.750
$r \leq 2$	0.017	4.421	4.062
$r \leq 3$	0.002	0.359	0.359

Model: $\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \alpha(\beta', \beta_1)(x'_{t-1}, t)' + u_0 + \varepsilon_t$

H_2	Eigenvalue	trace	$\lambda_{\max}$
$r = 0$	0.127	67.095**	31.775**
$r \leq 1$	0.084	35.320	20.572
$r \leq 2$	0.044	14.748	10.585
$r \leq 3$	0.018	4.163	4.163

Notes. An* (**) indicates significance at 10%(5%) level.
k=5 for the first model and k=10 for the second model.

Table 32. Test of the cointegration rank (Malaysia)

Model: $\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-1} + u_0 + \varepsilon_t$

H_2	eigenvalue	trace	λ_{max}
$r = 0$	0.083	49.591**	26.304*
$r \leq 1$	0.054	23.287	16.931
$r \leq 2$	0.018	6.357	5.491
$r \leq 3$	0.003	0.866	0.866

Model: $\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \alpha(\beta', \beta_1)(x'_{t-1}, t)' + u_0 + \varepsilon_t$

H_2	Eigenvalue	trace	λ_{max}
$r = 0$	0.091	64.826**	29.022
$r \leq 1$	0.068	35.803	21.349
$r \leq 2$	0.037	14.454	11.553
$r \leq 3$	0.010	2.902	2.902

Notes. An *(**) indicates significance at 10%(5%) level.
k=8 for all the models.

Table 33. Normalized cointegrating vectors ($\hat{\beta}$) and error correction coefficient ($\hat{\alpha}$) (South Korea)

Model: $\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-1} + u_0 + \varepsilon_t$

	$rm2_{t-1}$	y_{t-1}	i_{t-1}	if_{t-1}	Constant
$\hat{\beta}$	1.000	0.447 (3.863)	0.374 (0.951)	-0.297 (0.736)	-11.272
Eq.	$\Delta rm2$	Δy	Δi	Δif	-
$\hat{\alpha}$	-0.003 (0.001)	-0.008 (0.003)	-0.042 (0.103)	0.145 (0.116)	-

Model: $\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \alpha(\beta', \beta_1)(x'_{t-1}, t)' + u_0 + \varepsilon_t$

	$rm2_{t-1}$	y_{t-1}	i_{t-1}	if_{t-1}	Constant	Trend
$\hat{\beta}$	1.000	-0.037 (0.112)	0.003 (0.004)	-0.005 (0.003)	-4.474	-0.008 (0.001)
	$\Delta rm2$	Δy	Δi	Δif	-	-
$\hat{\alpha}$	-0.104 (0.029)	-0.067 (0.061)	6.872 (2.224)	3.245 (2.459)	-	-

Notes. k=5 for the first model and k=10 for the second model.

Table 34. Normalized cointegrating vectors ($\hat{\beta}$) and error correction coefficient ($\hat{\alpha}$) (Malaysia)

Model: $\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-1} + u_0 + \varepsilon_t$

	$rm2_{t-1}$	y_{t-1}	i_{t-1}	if_{t-1}	Constant
$\hat{\beta}$	1.000	-1.133 (0.039)	-0.089 (0.013)	0.036 (0.009)	-1.732
Eq.	$\Delta rm2$	Δy	Δi	Δif	-
$\hat{\alpha}$	-0.013 (0.008)	0.068 (0.025)	0.377 (0.700)	-0.720 (0.798)	- (-)

Model: $\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \alpha(\beta', \beta_1)(x'_{t-1}, t)' + u_0 + \varepsilon_t$

	$rm2_{t-1}$	y_{t-1}	i_{t-1}	if_{t-1}	Constant	Trend
$\hat{\beta}$	1.000	-4.090 (0.889)	-0.149 (0.033)	0.072 (0.020)	5.396	0.022 (0.007)
Eq.	$\Delta rm2$	Δy	Δi	Δif	-	-
$\hat{\alpha}$	-0.007 (0.004)	0.043 (0.011)	-0.054 (0.300)	-0.299 (0.342)	-	-

Notes. k=8 for all the models.

Table 35. Weak exogeneity test (South Korea)

Model: $\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-1} + u_0 + \varepsilon_t$

	$\alpha_1 = 0$	$\alpha_2 = 0$	$\alpha_3 = 0$	$\alpha_4 = 0$
LR	4.892**	4.173**	0.147	1.016
(P-value)	(0.027)	(0.041)	(0.701)	(0.313)

Model: $\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \alpha(\beta', \beta_1)(x'_{t-1}, t)' + u_0 + \varepsilon_t$

	$\alpha_1 = 0$	$\alpha_2 = 0$	$\alpha_3 = 0$	$\alpha_4 = 0$
LR	7.475**	0.795	7.540**	0.122
(P-value)	(0.006)	(0.373)	(0.006)	(0.290)

Notes. An *(**) indicates significance at 10% (5%) level.

Table 36. Weak exogeneity test (Malaysia)

Model: $\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-1} + u_0 + \varepsilon_t$

	$\alpha_1 = 0$	$\alpha_2 = 0$	$\alpha_3 = 0$	$\alpha_4 = 0$
LR	2.093	4.494**	0.141	0.488
(P-value)	(0.148)	(0.034)	(0.707)	(0.485)

Model: $\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \alpha(\beta', \beta_1)(x'_{t-1}, t)' + u_0 + \varepsilon_t$

	$\alpha_1 = 0$	$\alpha_2 = 0$	$\alpha_3 = 0$	$\alpha_4 = 0$
LR	2.832*	6.577**	0.017	0.542
(P-value)	(0.092)	(0.010)	(0.897)	(0.462)

Notes. An *(**) indicates significance at 10% (5%) level.

Table 37. Estimated coefficients of short-run model (South Korea)

Model: $\Delta rm2_t = u_0 + \alpha(L)\Delta i_{t-1} + \beta(L)\Delta y_{t-1} + \gamma(L)\Delta if_{t-1} + \alpha_1(\beta', \beta_1)(x'_{t-1}, t)' + \eta_t$

	coefficient	Std. Error	t-value
Δrm_{t-1}	0.002	0.072	0.025
Δrm_{t-2}	0.117	0.072	1.626
Δrm_{t-3}	0.043	0.072	0.594
Δrm_{t-4}	-0.037	0.073	-0.507
Δrm_{t-5}	0.033	0.073	0.454
Δrm_{t-6}	0.073	0.073	0.996
Δrm_{t-7}	0.026	0.073	0.356
Δrm_{t-8}	0.002	0.073	0.031
Δrm_{t-9}	0.056	0.074	0.754
Δrm_{t-10}	0.125	0.074	1.692
Δy_{t-1}	0.000	0.035	0.004
Δy_{t-2}	0.004	0.038	0.117
Δy_{t-3}	-0.049	0.039	-1.252
Δy_{t-4}	0.010	0.039	0.249
Δy_{t-5}	0.059	0.039	1.532
Δy_{t-6}	0.067	0.039	1.746
Δy_{t-7}	0.083	0.039	2.137
Δy_{t-8}	0.062	0.038	1.632

Δy_{t-9}	0.053	0.037	1.437
Δy_{t-10}	0.040	0.034	1.171
Δi_{t-1}	0.001	0.002	0.320
Δi_{t-2}	0.002	0.002	1.028
Δi_{t-3}	0.001	0.002	0.534
Δi_{t-4}	-0.002	0.002	-1.233
Δi_{t-5}	0.003	0.002	1.442
Δi_{t-6}	-0.002	0.002	-1.033
Δi_{t-7}	0.000	0.002	0.113
Δi_{t-8}	-0.004	0.002	-2.054
Δi_{t-9}	0.001	0.002	0.583
Δi_{t-10}	-0.003	0.002	-1.419
Δif_{t-1}	0.001	0.002	0.599
Δif_{t-2}	-0.000	0.002	-0.164
Δif_{t-3}	-0.001	0.002	-0.722
Δif_{t-4}	0.003	0.002	1.560
Δif_{t-5}	-0.003	0.002	-1.790
Δif_{t-6}	0.003	0.002	1.761
Δif_{t-7}	0.000	0.002	0.095
Δif_{t-8}	0.005	0.002	2.559
Δif_{t-9}	-0.001	0.002	-0.544
Δif_{t-10}	0.005	0.002	3.193
u_0	0.002	0.002	0.884
α_1	-0.104	0.029	-3.558

Table 38. Estimated coefficients of short-run model (Malaysia)

Model: $\Delta rm2_t = u_0 + \alpha(L)\Delta i_{t-1} + \beta(L)\Delta y_{t-1} + \gamma(L)\Delta if_{t-1} + \alpha_1(\beta', \beta_1)(x'_{t-1}, t)' + \eta_t$

	coefficient	Std. Error	t-value
Δrm_{t-1}	-0.023	0.063	-0.360
Δrm_{t-2}	-0.061	0.062	-0.977
Δrm_{t-3}	0.115	0.063	1.824
Δrm_{t-4}	0.014	0.063	0.220
Δrm_{t-5}	0.046	0.064	0.717
Δrm_{t-6}	0.065	0.064	1.012
Δrm_{t-7}	-0.063	0.064	-0.988
Δrm_{t-8}	-0.022	0.064	-0.343
Δy_{t-1}	-0.045	0.022	-1.993
Δy_{t-2}	-0.021	0.025	-0.825
Δy_{t-3}	-0.022	0.025	-0.857
Δy_{t-4}	-0.040	0.025	-1.577
Δy_{t-5}	-0.018	0.026	-0.723
Δy_{t-6}	0.013	0.025	0.512
Δy_{t-7}	0.034	0.024	1.410
Δy_{t-8}	0.027	0.019	1.405
Δi_{t-1}	0.001	0.001	0.766
Δi_{t-2}	-0.004	0.002	-1.988
Δi_{t-3}	-0.000	0.002	-0.181

Δi_{t-4}	-0.000	0.002	-0.162
Δi_{t-5}	-0.001	0.002	-0.293
Δi_{t-6}	-0.000	0.002	-0.186
Δi_{t-7}	-0.001	0.002	-0.450
Δi_{t-8}	-0.001	0.002	-0.499
Δif_{t-1}	-0.001	0.002	-0.855
Δif_{t-2}	0.003	0.002	1.623
Δif_{t-3}	-0.001	0.002	-0.322
Δif_{t-4}	-0.000	0.002	-0.109
Δif_{t-5}	-0.001	0.002	-0.317
Δif_{t-6}	0.002	0.002	1.032
Δif_{t-7}	0.001	0.002	0.551
Δif_{t-8}	0.001	0.002	0.480
u_0	0.009	0.002	4.629
α_1	-0.006	0.004	-1.817

Table 39. Summary of studies of the demand for real money balances involving Cointegration/Error-Correction modeling in South Korea and Malaysia

Country/ Author	Sample period/ Frequency	Monetary Aggregates	Determinants		Unit-Root Test(s)	Cointegration Test(s)	Findings
			Scale Variable	Opportunity Cost Variable			
South Korea							
Arize (1994)	1973:1-1990:1 /Quarterly	ln(M1/CPI) ln(M2/CPI)	ln(real GDP)	Yield on CB: interest rate on loans and TDR on NCB; interest rate differential; expected rate of inflation	ADF (1981); Hylleberg and others (1990); PP(1988)	EY (1987); J (1988)	Two to three cointegrating vectors among real money (both M1 and M2), real income, interest rate, and foreign exchange rate risk and return. Well-specified ECM.
Kim (2002)	1970:1-1996:12 /Monthly	ln(M2/CPI)	ln(IIP)	Money market rate; interest rate differential;	ADF (1981); PP (1988); KPSS (1992)	PO (1990); JJ (1990)	No cointegrating vector among real money, real income, interest rate, and interest rate differential for the model with a drift. One cointegrating vector among the above variables for the model with a time trend. Well-specified ECM.

Malaysia							
Tan (1997)	1981:4 - 1991:4 /Quarterly	ln(M0/CPI) ln(M1/CPI) ln(M2/CPI)	ln(real GDP)	TDR on CBS; TBR; expected exchange rate	ADF (1981); Hylleberg and others (1990);	JJ (1990)	One cointegrating vector among real money (all M0, M1, and M2), real income, interest rate, and expected exchange rate. Well-specified ECM.
Sriram (1999a)	1973:8-1995:12 /Monthly	ln(M2/CPI)	ln(IIP)	CBTDR3M; discount rate on TB; foreign interest rate; expected rate of inflation; nominal exchange rate	DF (1976); ADF (1981);	JJ (1990)	One cointegrating vector among real money, real income, interest rate, expected rate of inflation, and exchange rate depreciation for the model of closed and open economy. Well-specified ECM.
Kim (2002)	1970:1-1996:12 /Monthly	ln(M2/CPI)	ln(IIP)	Money market rate; interest rate differential;	ADF (1981); PP(1988); KPSS (1992)	PO (1990); JJ (1990)	One cointegrating vector among real money, real income, interest rate, and interest rate differential for the models with a drift and with a time trend. Well-specified ECM.

Notes. The following abbreviations are used:
Monetary aggregates: M0 = currency-in-circulation
Scale variables: IIP = index of industrial production
Interest rate: TDR = time deposit rate; TBR = treasury bill rate; CBTDR3M = Three-month deposit rates at commercial banks
Unit-Root tests: ADF = augmented Dickey-Fuller; PP = Phillips and Perron; KPSS = Kwiatkowski, Phillips, Schmidt, and Shin; DF = Dickey-Fuller
Cointegration tests: EY = Engle and Yoo; J = Johansen; PO = Phillips and Ouliaris; JJ = Johansen and Juselius
General: CB = corporate bonds; CBS = commercial banks; NCB = nationwide commercial banks

Figure 29. The log of real M2 (South Korea)

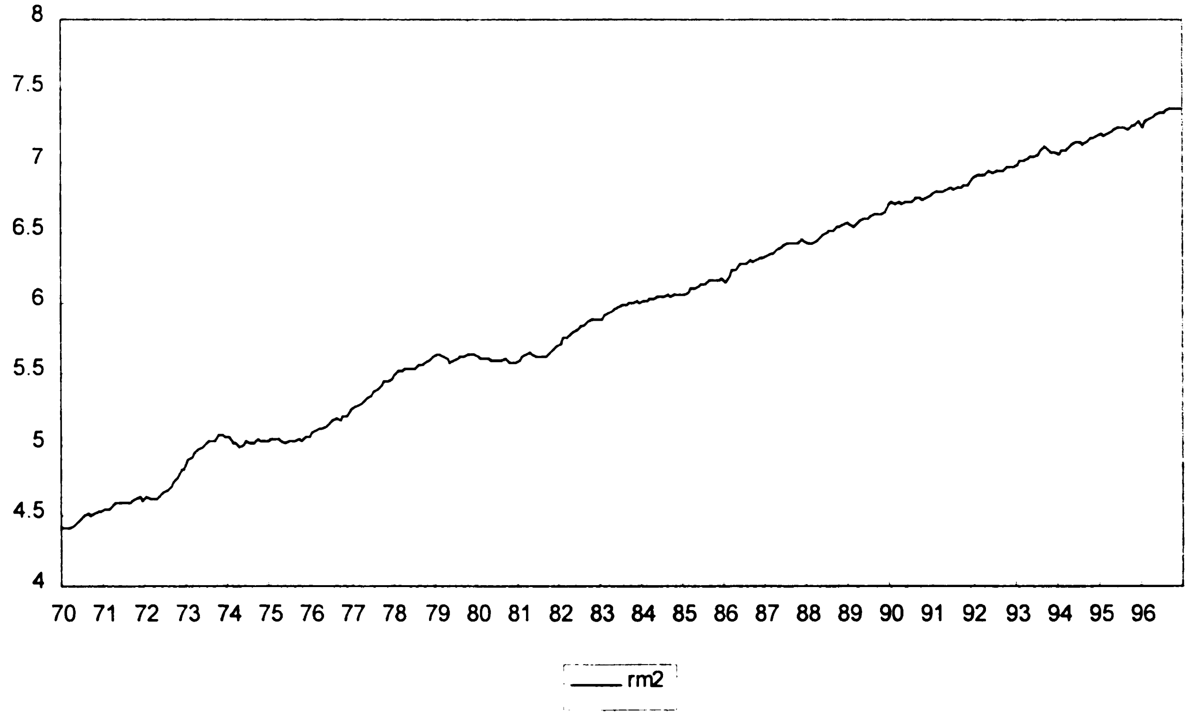


Figure 30. The log of real GDP (South Korea)

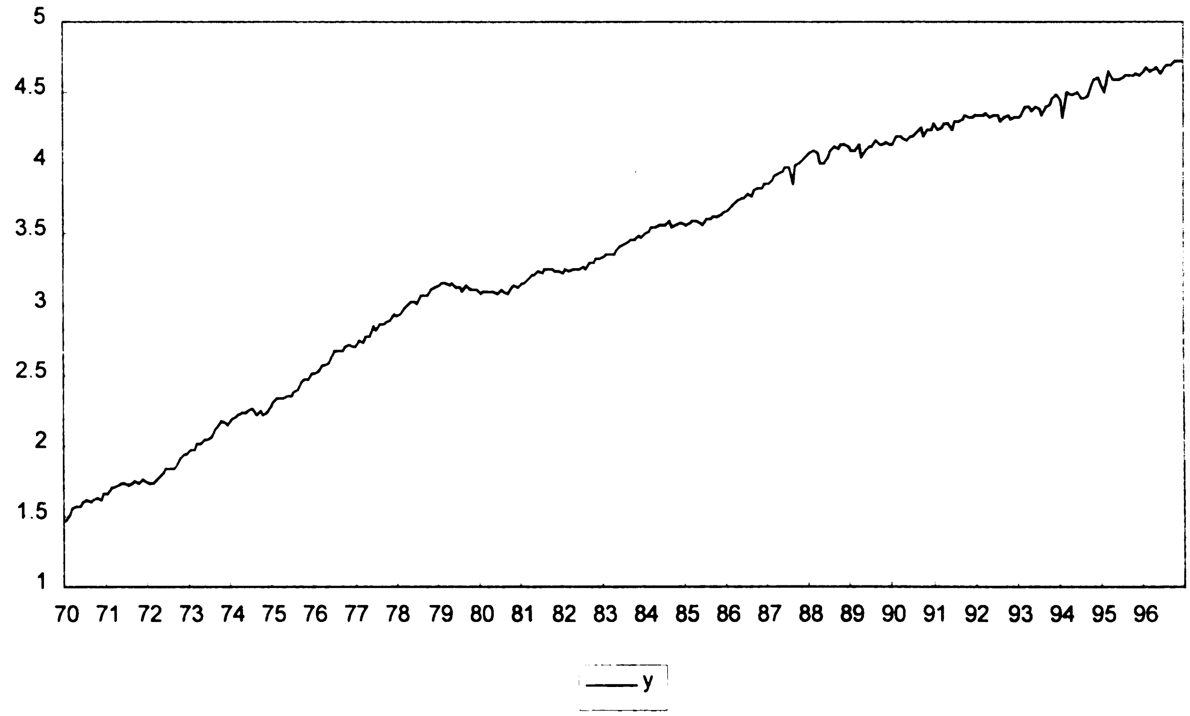


Figure 31. Interest rate (South Korea)

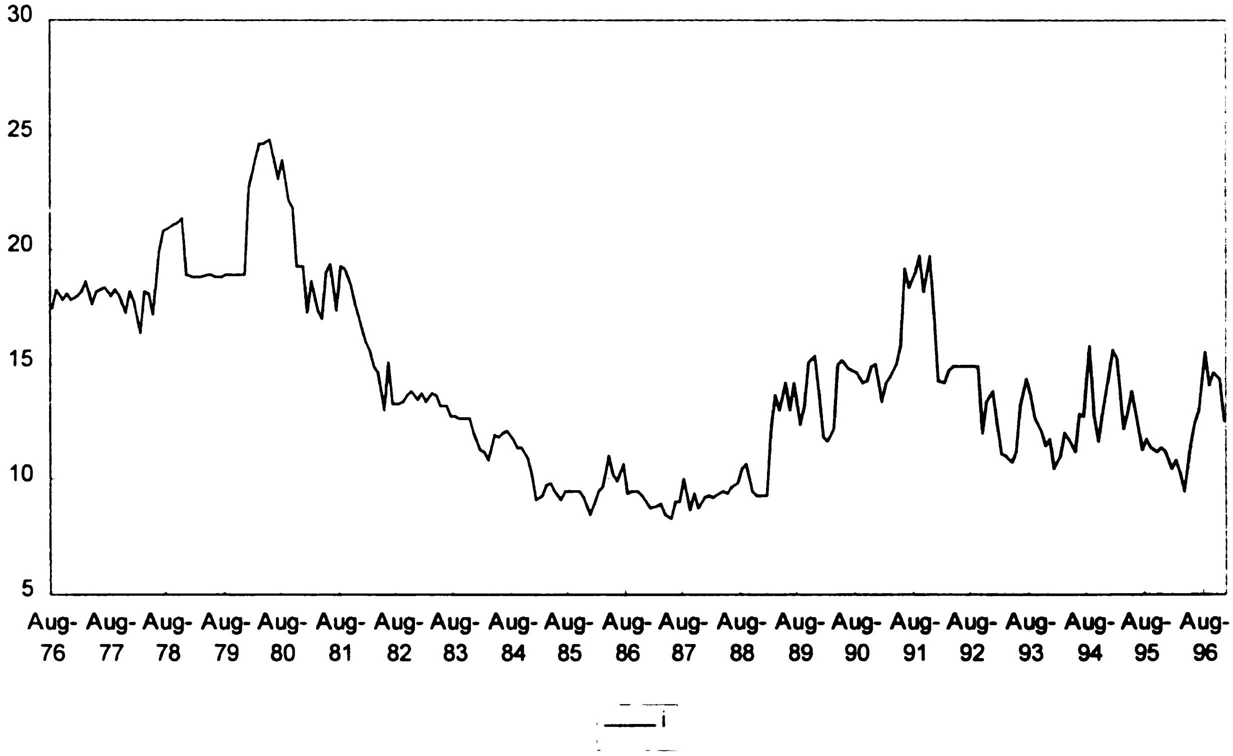


Figure 32. Difference between domestic and foreign interest rate (South Korea)

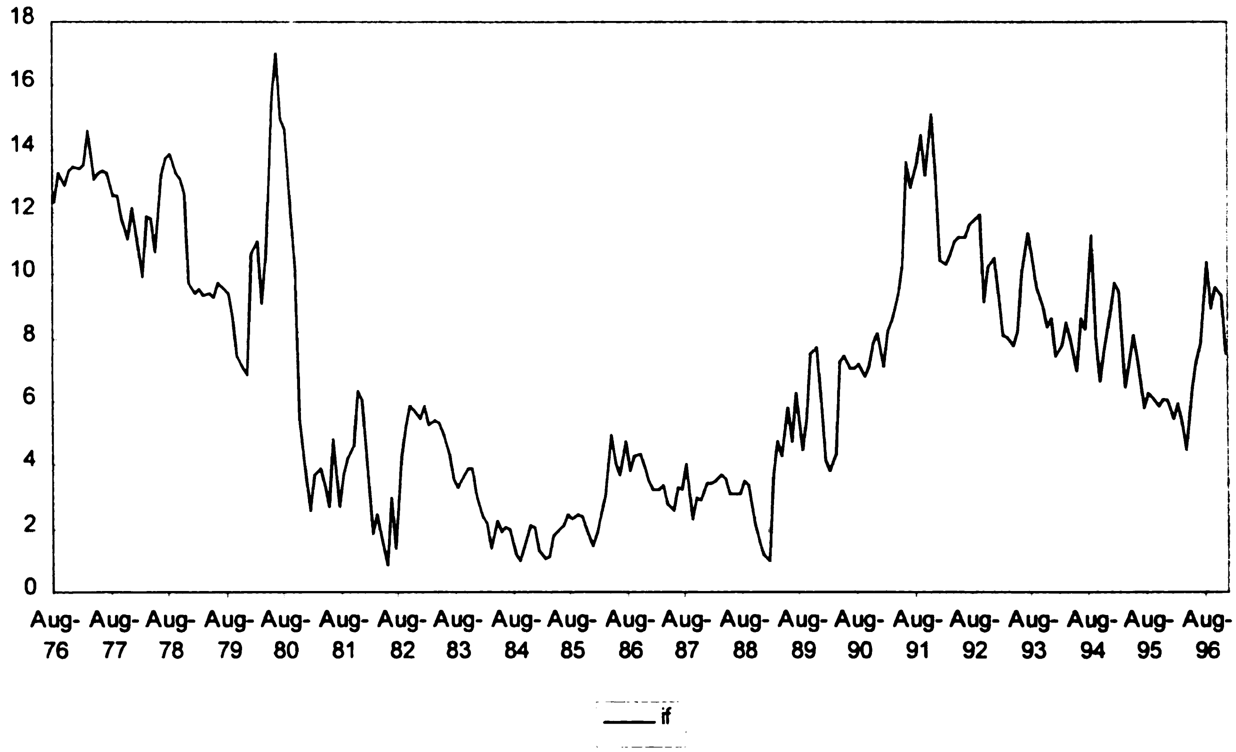


Figure 33. The log of real M2 (Malaysia)

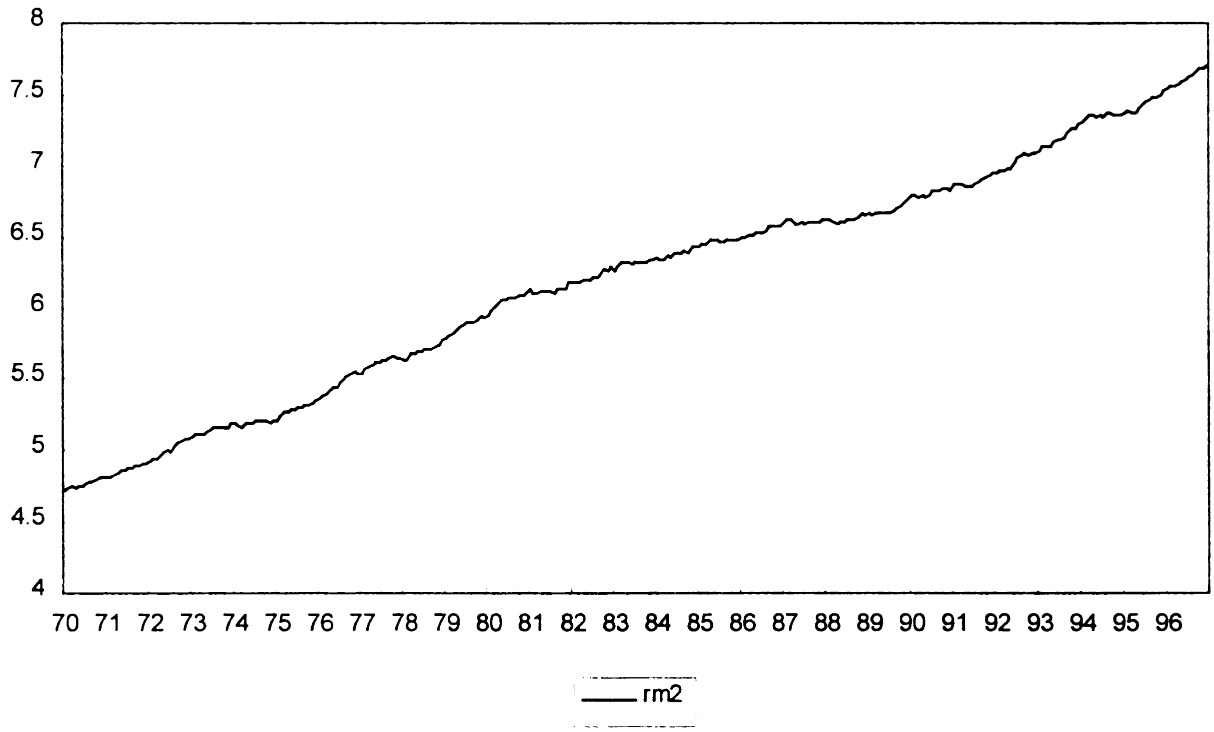


Figure 34. The log of real GDP (Malaysia)

Figure 35. Interest rate (Malaysia)

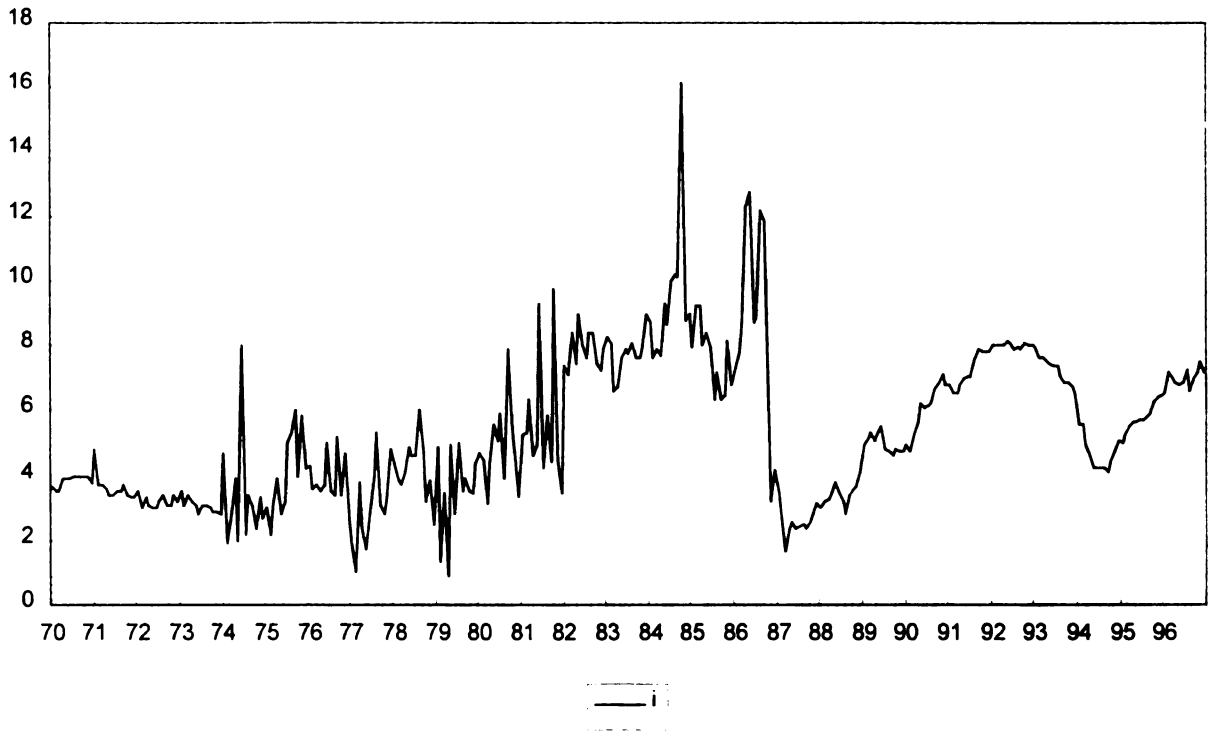


Figure 36. Difference between domestic and foreign interest rate (Malaysia)

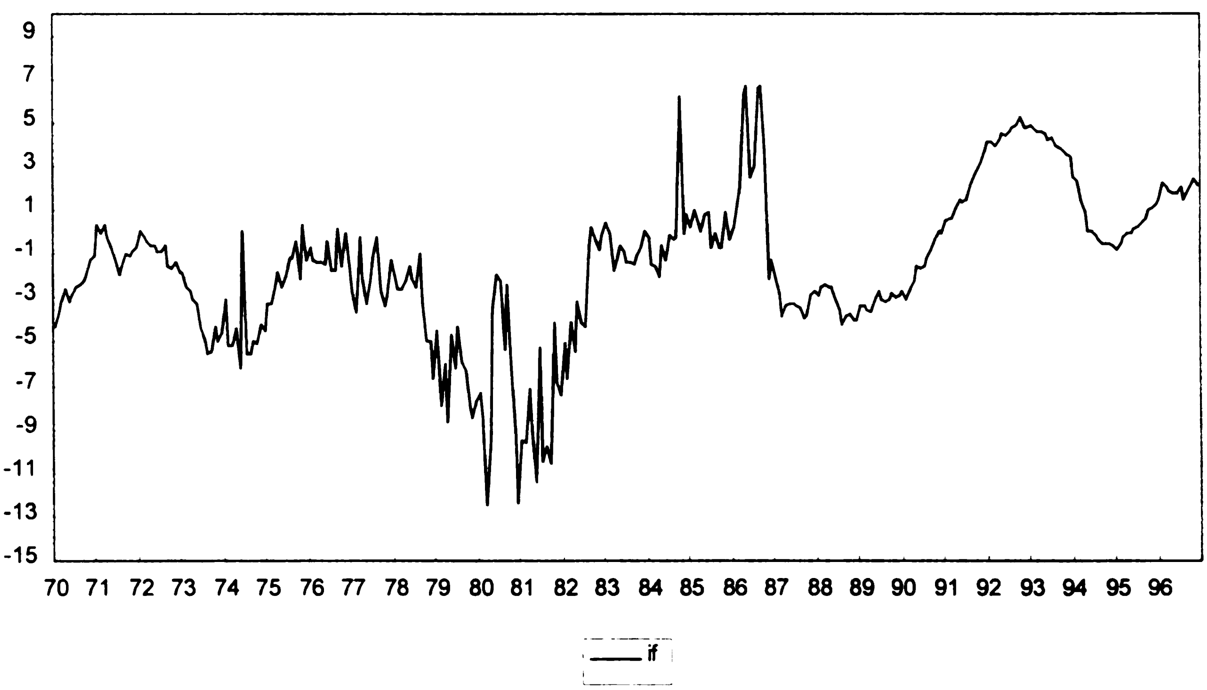


Figure 37. Recursive estimates of the long-run parameter of real GDP (South Korea)

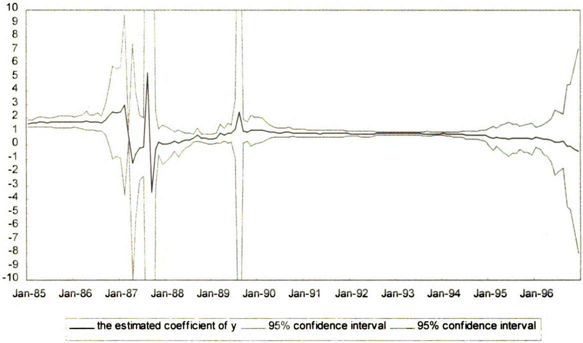


Figure 38. Recursive estimates of the long-run parameter of interest rate (South Korea)

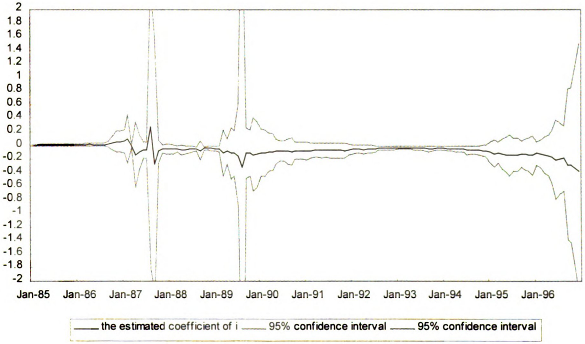


Figure 39. Recursive estimates of the long-run parameter of the difference of domestic and foreign interest rate (South Korea)

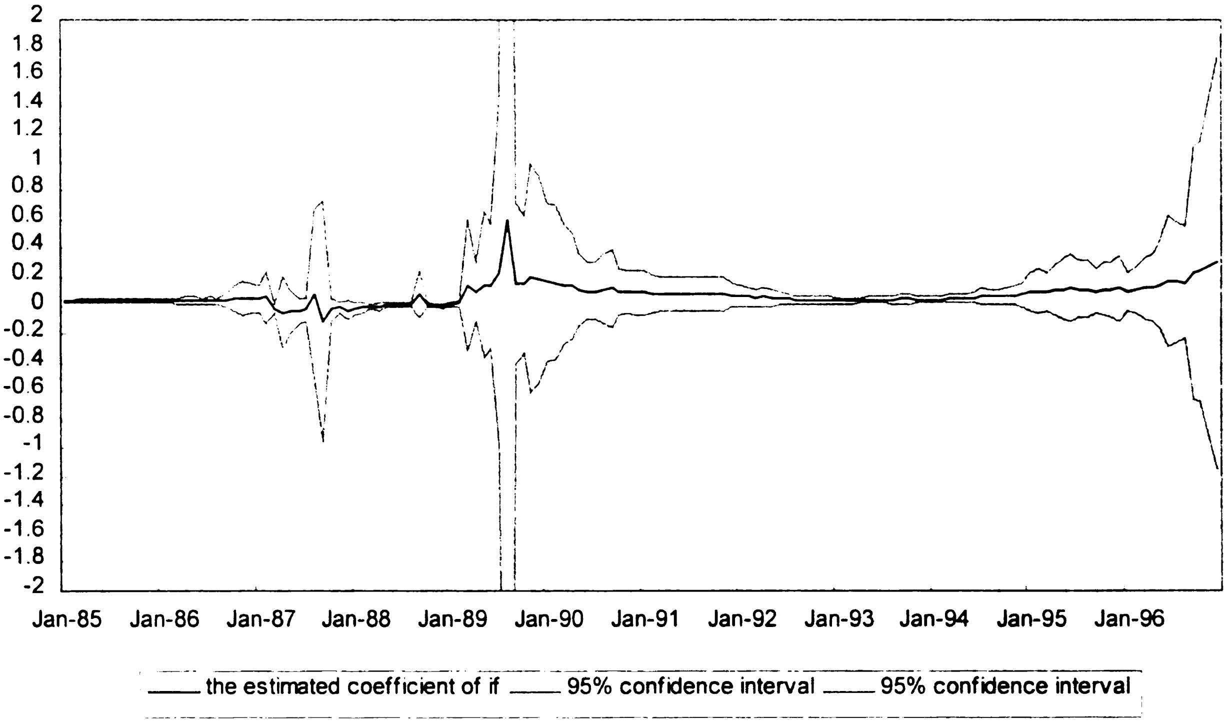


Figure 40. Recursive estimates of the long-run parameter of real GDP (Malaysia)

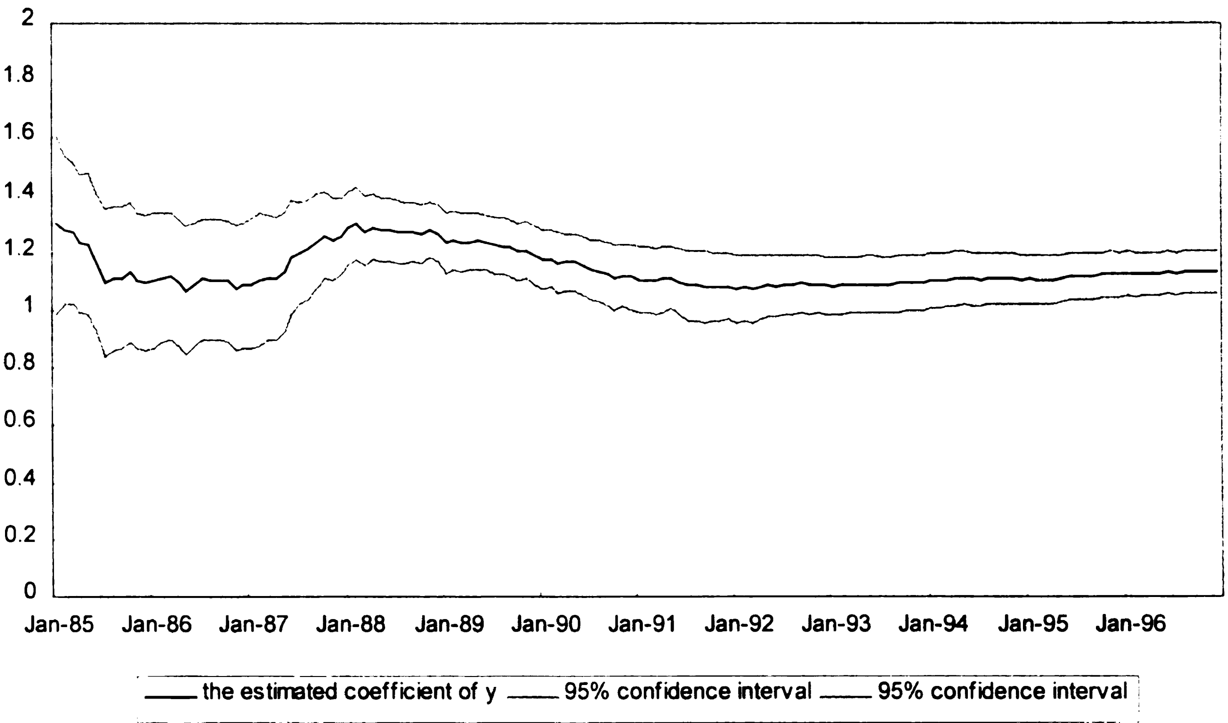


Figure 41. Recursive estimates of the long-run parameter of interest rate (Malaysia)

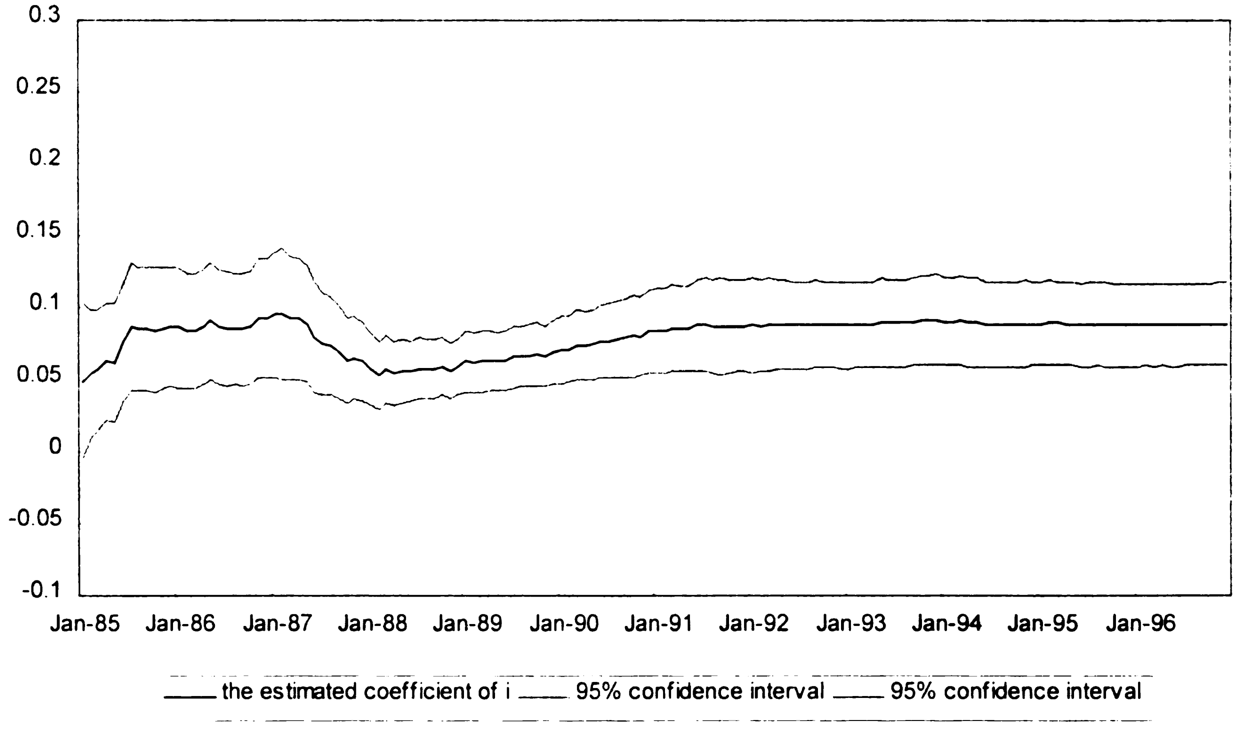
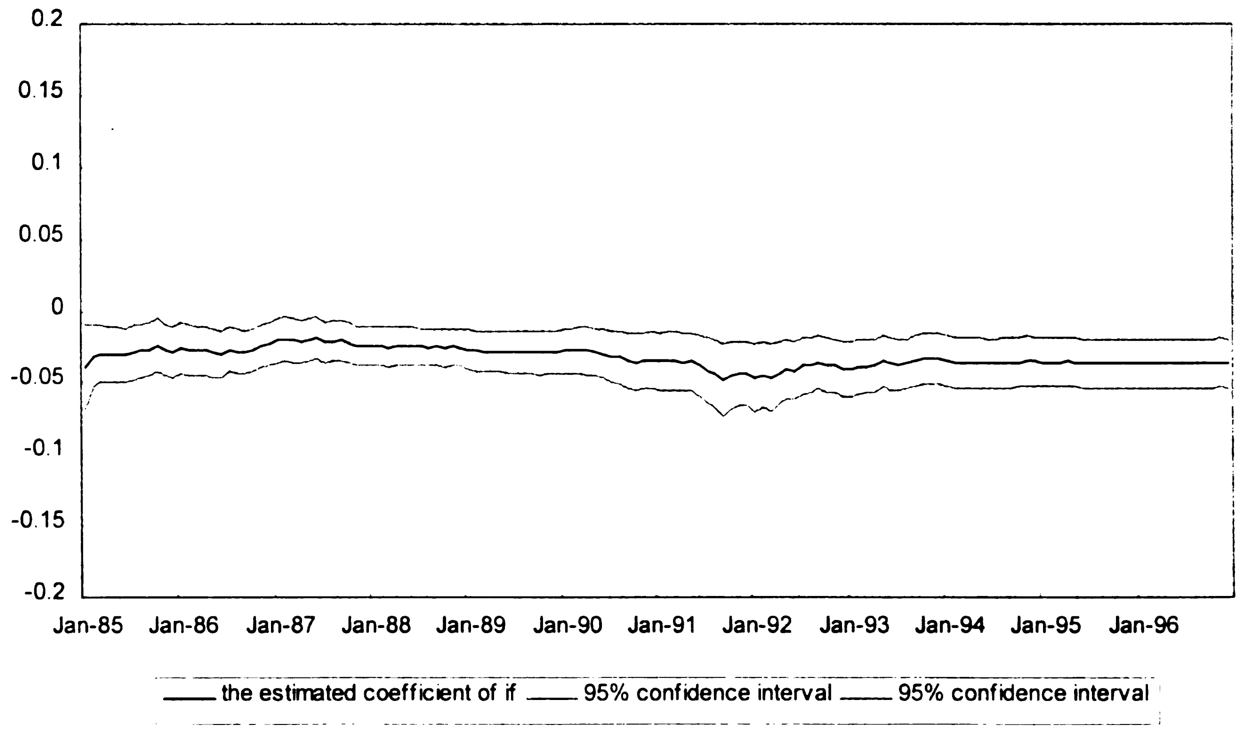


Figure 42. Recursive estimates of the long-run parameter of the difference of domestic and foreign interest rate (Malaysia)



CHAPTER VI

FORECAST OF SHADOW EXCHANGE RATE AND PROBABILITY OF COLLAPSE

1. Introduction

Currency crises are thought to have a significant predictable component. The first generation models of currency crises identify fundamentals judged to be useful in the prediction of currency crises as typified in the influential paper, Krugman (1979). A fiscal deficit financed by domestic credit creation is considered to be the root cause of a speculative attack. Since the monetary authority monetizes the fiscal deficit, the oversupply of money causes a gradual decline in international reserves. Accordingly, investors attack the fixed exchange rate with the effect being the depletion of the government's reserve holdings needed to defend its currency.

The currency crises in Europe and Mexico during the early 1990s, however, do not lend support to these traditional factors playing a major role in their crises. Otker and Pazarbasioglu (1996), when thy computed the probability of an exchange rate regime change using Blanco and Garber's model, found the Mexican financial crisis in 1994 was not the result of fiscal imbalances, which had previously played a major role in Mexico's balance of payments crises. Instead, it was the rise in private sector indebtedness and a corresponding increase in the amount of credit owed to the banking system that built up the pressure in Mexico's exchange market in mid 1994. Moreover, the experiences of several European countries in the context of the European Monetary System (Otker and

Pazarbasioglu, 1997b) show evidence additional triggering determinants of crises - e.g. pure speculation - that cannot be explained by the fundamentals. These studies focus on the uniqueness of each country's currency crisis during a specific time period.

However, while the wide range of analyses of European and Mexican currency crises have found a consensus about the causes of the crises, the cause of the Asian currency crisis is still under discussion among researchers. Therefore, the empirical study in this chapter will apply a basic and an extended model, first introduced in Chapter III, to the crises in South Korea and Malaysia. Two countries experienced severe devaluations of their currencies during the Asian currency crisis.

The objective of this chapter is to obtain the probability of an exchange rate regime change as a function of economic fundamentals by using the implications of the speculative attack literature to identify the contribution of weak economic fundamentals in both South Korea's and Malaysia's currency crisis. If the probability of collapse was high enough to cause the currency crisis in South Korea and Malaysia during the Asian currency crisis, then support can be found that sound macroeconomic policy may have been able to prevent the currency crisis. Hence, the arguments made by weak fundamentalists are justified.

The remainder of this chapter is organized as follows. Section 2 outlines the estimation procedure. Section 3 presents the empirical results and section 4 offers concluding remarks.

2. Estimation procedure

To obtain the probability of collapse, it is necessary to estimate the shadow exchange rates. As shown in Chapter III, the shadow exchange rate derived from the basic model is

$$\tilde{s}_{t+1} = m_{t+1} - a_0 + a_1 i^*_{t+1} - a_2 y_{t+1} - p^*_{t+1} - u_t + (a_1 + a_3)(E_t \mu^d_{t+1} + \rho_{t+1}) \qquad (1)$$

where $m_{t+1} = d_{t+1} + r_c$ and $E_t \mu^d_{t+1} + \rho_{t+1} = i_{t+1} - i^*_{t+1}$.[21] Then, the probability of collapse, π_t, is obtained. The probability of collapse is the probability that the shadow exchange rate, $\tilde{s}_{t+1}$, will exceed $\bar{s}_t$[22] in period *(t+1)* or $\pi_t = \Pr[\tilde{s}_{t+1} - \bar{s}_t \rangle 0]$.

However, the basic model was extended to resolve the problems caused by the non-stationarity of variables. The shadow exchange rate derived from the extended model is

$$\tilde{s}_{t+1} = m_{t+1} - \mu_0 - \alpha(L)\Delta i_{t-1} - \beta(L)\Delta y_{t-1} - \gamma(L)\Delta if_{t-1} - \Pi_1 x_t - rm_t - p^*_{t+1} - u_{t+1} \qquad (2)$$

where $m_{t+1} = d_{t+1} + r_c$ and $if_{t-1} = i_{t-1} - i^*_{t-1}$.[23] The probability of collapse in the extended model is defined as same as the one in the basic model.

The estimation of the shadow exchange rate, $\tilde{s}_{t+1}$, requires two additional procedures. The first procedure is to make forecasts of the variables, p^*, *y*, *i*, i^*, *d*, and *u*, which are assumed to evolve according to a period-by-period systematic stationary

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[21] *m*, *d*, *r*, *p*, p^*, and *y* are the logarithms of the money stock, domestic credit extended by the domestic banks, central bank foreign reserves, domestic price level, foreign price level, and real output, respectively. *i* is the domestic interest rate, i^* is the foreign nominal interest rate, ρ is the risk premium on domestic assets, *u* is the logarithms of the deviation from PPP and $E_t \mu^d_{t+1}$ is a period-by-period systematic stationary component of *d*.

[22] This is the time *t* value of the fixed rate.

[23] Π_1 is the first row of the Π and $x'_{t-1} = (rm_{t-1}, a_0, t, i_{t-1}, y_{t-1}, if_{t-1})$.

component, $E_{t-1}\mu_t$, and a stochastic element, ε_t. A forecast of these variables is made in Chapter IV using the ARFIMA(p,d,q)-FIGARCH(P,δ,Q) model. These forecasted values are substituted into equation (1) and (2). The next procedure is an estimation of the money demand parameters a_0, a_1, a_2 and a_3 from equation (1) and $\alpha(L)$, $\beta(L)$, $\gamma(L)$ and Π_1 from equation (2). Under the possibility of a spurious regression, the coefficients estimated in the OLS regression[24], $\hat{a}_0, \hat{a}_1, \hat{a}_2$ and $\hat{a}_3$ are initially used to derive $\tilde{s}_{t+1}$ in equation (1). After testing for the number of cointegration relations and estimating the cointegrating vectors, the estimated coefficients for the extended model, $\hat{\alpha}(L)$, $\hat{\beta}(L)$, $\hat{\gamma}(L)$ and $\hat{\Pi}_1$, are applied to the estimation of $\tilde{s}_{t+1}$ in equation (2).

Assuming that each stochastic part of variable, p^*, y, i, i^*, d, and u, is uncorrelated with each other and their linear combination is normally distributed, the probability of collapse, π_t, in the basic and extended model can be estimated as

$$\pi_t = \Pr[\varepsilon_{t+1} > k_t] = \int_{k_t}^{\infty} \frac{1}{E_t\sigma_{t+1}\sqrt{2\pi}} e^{-\frac{\varepsilon_{t+1}^2}{2E_t\sigma_{t+1}^2}} d\varepsilon \tag{3}$$

where $k_t = \bar{s}_t - d_t - r_c + a_0 - (a_1 + a_3)i_t + a_3 i_t^* + a_2 y_t + p_t^* + u_t - E_t\mu_{t+1}$. $E_t\mu_{t+1}$ and ε_{t+1} are a linear combination of the systematic stationary components and stochastic parts of p^*, y, i, i^*, d, respectively, and u and σ_{t+1}^2 is a conditional variance of ε_{t+1}.

[24] Normalized cointegrating vector which has long run equilibrium coefficients, $\hat{a}_0, \hat{a}_1, \hat{a}_2$ and $\hat{a}_3$, is also used for the comparison with the result derived by OLS estimates.

3. Empirical results

3.1 Behavior of variables in the structural model

Before the results of the estimation of the shadow exchange rate and the probability of collapse are presented, a graphical inspection of the traits of each variable in the structural model needs to be discussed. Figures 43 and 48 show the real *M2*[25] of South Korea and Malaysia, respectively. The real value of *M2* is persistently increasing for both countries up until the collapse. Then, for both countries, the real M2 turned downward right after the crisis. After a few months, though, the real *M2* of South Korea began to grow again at a steeper rate than the rate before the currency crisis. By contrast, the real *M2* kept growing slowly in Malaysia.

Figures 44 and 49 display the domestic credits[26] of South Korea and Malaysia. They show that domestic credit increases at a faster rate in South Korea than Malaysia after the collapse. This indicates that a currency depreciation pressure induced by growing domestic credit began to increase again in South Korea. In addition, the decline in domestic credit following the crisis in both countries shows that the second-generation models' view that speculators would expect an immediate increase in domestic credit after the crisis is not supported by the data.

Figure 45 and 50 show that real GDP[27] for both South Korea and Malaysia declined for a while after the currency collapse. However, it took longer for real GDP to

[25] Logarithm value.
[26] Logarithm value.
[27] Logarithm value.

recover than it took for the real *M2* and domestic credit to begin to rise again. In addition, even after it began to rise again, both countries' real GDPs show an unstable tendency.

The abrupt increases in the interest rates in both countries at the point of collapse in Figure 46 and 51 are not surprising when one considers the bottleneck that the currency crisis caused in the financial market. Nevertheless, interest rate movements after the collapse show stability during the recovery period from the collapse. The deviation from PPP, the negative value of the real exchange rate, mimics, in the opposite direction, the movement of the interest rates at the point of collapse in both countries. Whereas the currency did depreciate in real terms by 9.5 percent in South Korea during 1996[28], there was a substantial appreciation to the real exchange rate relative to the 1980's real exchange rate during the 1990s prior to the collapse as shown in Figure 47. This encourages us to anticipate a higher probability of collapse in the 1990s than in the 1980s. In addition, the deviation from PPP in Malaysia shows that the real exchange rate depreciation that continued during the 1980s turned to appreciation in early 1992 in Figure 52. Therefore, the real exchange rate's graphical appearance for both countries may have signaled the upcoming Asian currency crisis.

3.2 Estimated shadow exchange rate

The shadow exchange rate, $\tilde{s}_{t+1}$, at time $t+1$, is the floating rate crisis[29], that would clear the foreign exchange market if the central bank stops defending its fixed

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[28] Refer to Table 8.
[29] Refer to Chapter III.

parity. As discussed in Chapter III, in the first generation models of currency the shadow exchange rate is an index used by speculators to decide when to attack. This is because the condition for profitable attack is when the postcollapse exchange rate, $\tilde{s}_{t+1}$, is larger than the prevailing fixed rate, $\bar{s}_t$. Profits of speculators are equal to the exchange rate differential multiplied by the reserve stock used to defend the fixed rate regime.

The second and third columns of Table 40 report the quarterly data of actual and shadow exchange rates[30] of South Korea in 1990s. The difference between the actual and shadow exchange rates was not noticeable until March 1994. However, after March 1994, the difference became larger and the actual and shadow exchange rates were 6.82 and 7.04 in September 1997, 3 months prior to South Korea's currency crisis. Therefore, the shadow exchange rate, commonly interpreted to reflect weak fundamentals, gave a warning signal to the policy makers before South Korea's currency crisis.

Figure 55 shows the actual and shadow exchange rates in South Korea from 1977:08 to 2000:11. Three methods are used to derive of shadow exchange rates; the first uses the coefficients from the OLS estimation; the second uses the normalized long-run cointegrating vector; and the third uses the short-run model of real money demand function. Although different coefficients are used to estimate the shadow exchange rates, the different estimated shadow exchange rates exhibit no substantial differences during the period. The graphical trend of the difference in the actual and shadow exchange rates shows an increase in the difference before the currency crisis. This implies that the currency crisis in 1997 may have been predictable.

Similarly the third and fourth columns of Table 40 present a comparison of the actual and shadow exchange rates of Malaysia during the 1990s. The difference between the actual and shadow exchange rate was increasing since March 1992. The actual and shadow exchange rates became 0.93 and 1.06 just before the currency crisis in Malaysia. One can, thus, interpret the divergence between shadow and actual exchange rates in Malaysia as an indication that there was going to be a currency crisis.

Figure 56 provides a graphical inspection of the actual and shadow exchange rates from 1971:02 to 2000:11 for Malaysia. It shows similar features in the1970s as in South Korea. However, while the Won[31] continued to stay its highly depreciated level throughout the 1980s, the Malaysian Ringgit converged to the shadow exchange rate during the 1980s. With Malaysia's capital transaction liberalization still incomplete during the 1980s, its trade balance surplus helped to resist the devaluation of the Ringgit. Then, as Malaysia's financial system became fairly well-developed with the 1970s and 1980s' liberalization and innovation, the capital account transaction surplus matched their growing current account deficit[32] during the 1990s. This happened just as it did in South Korea. Therefore, despite Malaysia's ongoing current account balance deficit in 1990s before the crisis, the Ringgit did not depreciate enough to end the deficit. The divergence between the shadow and actual exchange rates similarly implies weak fundamentals in the Malaysian economy that provided the opportunity for speculators to attack the foreign

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[30] Logarithms

[31] The South Korean currency

[32] See Table 1.

exchange market. Hence, the trend of shadow exchange rate displayed in Figure 50 also provides enough reason for the outbreak of the currency crisis in Malaysia, 1997.

3.3 Estimated probability of collapse

The second and third columns of Table 41 present the probabilities of a collapse in South Korea based on the coefficients estimated by OLS and ECM, respectively. As anticipated before the estimation, the estimated probability hovered around 1.0 for the three years before the currency crisis in South Korea. The probabilities also show that the devaluation of the currency decreased the probability of a currency crisis to a low level when South Korea accepted a free foreign exchange rate regime in December 1997.

Figures 57 to 59 also show the probabilities of collapse in South Korea. Probability I is based on the coefficients from the OLS estimation and probabilities II and III are base on the coefficients from the normalized long-run cointegrating vector and the short-run model of real money demand function estimated by an ECM. While the techniques used to estimate the coefficients are quite different, the results show a similar trend among the three probability series. The probability series in Figures 57 to 59 indicate another warning signal following the crisis. A depreciation pressure caused by weak fundamentals began to grow again after the crisis in 1997 and, around July 1999, reached the same level as it was at right after the Asian currency crisis. Therefore, the current ongoing devaluation of currency in South Korea may be explained by our model as due to continuing weak fundamentals.

Similarly, the probabilities of collapse in Malaysia based on same techniques used for South Korea are presented in the fourth and fifth columns of Table 41. The probability of collapse in Malaysia was rather high for the years, 1996-97, prior to the crisis. Therefore, the depreciation pressure from weak fundamentals may have been cumulated before the currency crisis as an indication of the likely onset of currency crisis. Figures 60 to 62 present the probabilities of collapse in Malaysia. However, unlike the probability series estimated for South Korea, one of the probability series in Figure 61 estimated with a the normalized long-run cointegrating vector does not appear to predict the 1997 crisis as well as the other probability series. If we consider that the coefficients are normalized and the forecasted real money demand using the normalized cointegrating vector does not fit well generally, one would expect the estimated probability not to be completely reliable. For about two years, following South Korea, the probability series settled at the highest level before the regime change in Malaysia over the years 1996-97. Then, the probability series declined to a new low level directly following the steep currency devaluation in July 1997. This upward trend of probabilities also indicates that Malaysia's economic condition may have been unstable before the crisis in 1997.

While South Korea chose to move to a free exchange rate regime after the crisis, Malaysia returned to a fixed exchange rate system with a strict capital transaction regulation. However, unlike South Korea, there was no devaluation pressure under the fixed exchange rate regime after the collapse in 1997 as shown by the probability series in Figures 60 to 62. The lack of devaluation pressure implies that the Malaysian government's economic policies designed to stabilize their economy have been strongly effective.

4. Conclusion

Using the speculative attack model previously applied by Blanco and Garber (1986), Goldberg (1994), and Otker and Pazarbasioglu (1996, 1997b), we estimated the shadow exchange rates and probabilities of collapse in South Korea and Malaysia respectively in this chapter. However, the model employed in this chapter is modified to overcome the spurious regression problem by utilizing the error correction model (ECM) from Engel and Granger (1987).

A simple graphical investigation of important macroeconomic variables in the model initially presents evidence that the Asian currency crisis may have been predictable. The domestic credit of each country, which is one of the indispensable variables causing currency crisis, showed a steady increasing tendency before the Asian currency crisis. This implies that the depreciation pressure from the oversupply of domestic money was cumulating gradually before the crisis. The deviation from PPP for both countries also signaled the possibility of an currency crisis by showing that the domestic currency had excessively appreciated in the 1990s.

The estimated shadow exchange rate and probability of collapse for each country reflects the disequilibrium between real money supply and demand. This disequilibrium provided a signal of upcoming severe currency depreciation. In particular, the shadow exchange rate of each country was so far above the regulated exchange rate for about two to three years before the crises that this might attract speculators to the potential profit attainable following the outbreak of crisis. The probability of collapse, strongly positively correlated with the difference between the shadow and actual exchange rates, indicates that both countries, South Korea and Malaysia, as of the early 1990s fell under severe depreciation pressure that then lead to the Asian currency crisis likewise.

In conclusion, based on the evidence presented by the data, the movement over time of the shadow exchange rates and the probability of collapses, it is confirmed that fundamentals were weak prior to the Asian currency crisis in 1997. However, even though weak fundamentals are strong indicators of upcoming currency crisis, the specific point of the outbreak of the Asian currency crisis is not predicted by the evidence given in this chapter. Therefore, unexpected events such as bank failure, corporate failure and political uncertainty under weak fundamentals are additional factors for the ignition of the currency crisis.

Table 40. Actual and shadow exchange rates

	South Korea		Malaysia	
	Actual exchange rate	Shadow exchange rate	Actual exchange rate	Shadow exchange rate
1990- March	6.55	6.45	1.00	0.97
June	6.57	6.48	1.00	0.97
September	6.57	6.50	0.99	0.98
December	6.57	6.60	0.99	0.99
1991- March	6.59	6.59	1.02	1.00
June	6.58	6.61	1.02	1.00
September	6.61	6.62	1.01	1.01
December	6.63	6.69	1.00	1.01
1992- March	6.65	6.66	0.95	1.00
June	6.67	6.68	0.92	1.00
September	6.67	6.67	0.92	1.00
December	6.67	6.68	0.96	1.00
1993- March	6.68	6.67	0.95	1.00
June	6.69	6.68	0.95	1.01
September	6.70	6.73	0.94	1.01
December	6.69	6.72	0.99	1.03
1994- March	6.69	6.72	0.98	1.03
June	6.69	6.76	0.96	1.03
September	6.68	6.77	0.94	1.04
December	6.67	6.82	0.94	1.04
1995- March	6.65	6.80	0.93	1.05
June	6.63	6.85	0.89	1.05
September	6.64	6.85	0.92	1.06
December	6.65	6.90	0.93	1.06
1996- March	6.66	6.89	0.93	1.06
June	6.70	6.92	0.91	1.06
September	6.71	6.96	0.92	1.06
December	6.74	6.99	0.93	1.06
1997- March	6.80	7.02	0.91	1.05
June	6.79	7.02	0.93	1.06
September	6.82	7.04	1.16	1.08
December	7.44	7.24	1.36	1.11

Table 41. Probabilities of collapse

	South Korea		Malaysia	
	Probability of collapse (OLS)	Probability of collapse (ECM)	Probability of collapse (OLS)	Probability of collapse (ECM)
1990- March	0.01	0.00	0.33	0.31
June	0.01	0.00	0.39	0.38
September	0.02	0.02	0.43	0.42
December	0.47	0.87	0.48	0.48
1991- March	0.34	0.55	0.40	0.38
June	0.47	0.84	0.41	0.41
September	0.23	0.70	0.50	0.50
December	0.60	0.99	0.54	0.52
1992- March	0.37	0.61	0.75	0.74
June	0.29	0.57	0.86	0.85
September	0.36	0.59	0.89	0.87
December	0.42	0.71	0.75	0.71
1993- March	0.20	0.33	0.79	0.75
June	0.30	0.41	0.84	0.81
September	0.71	0.88	0.89	0.85
December	0.54	0.71	0.75	0.67
1994- March	0.42	0.81	0.81	0.73
June	0.87	0.99	0.89	0.83
September	0.93	0.98	0.93	0.90
December	1.00	1.00	0.94	0.91
1995- March	0.98	1.00	0.95	0.93
June	1.00	1.00	0.99	0.98
September	1.00	1.00	0.98	0.97
December	1.00	1.00	0.97	0.96
1996- March	1.00	1.00	0.97	0.95
June	1.00	1.00	0.98	0.97
September	1.00	1.00	0.98	0.96
December	1.00	1.00	0.98	0.96
1997- March	1.00	1.00	0.99	0.97
June	1.00	1.00	0.98	0.96
September	1.00	1.00	0.27	0.16
December	0.02	0.02	0.00	0.00

Figure 43. The log of real M2 (South Korea)

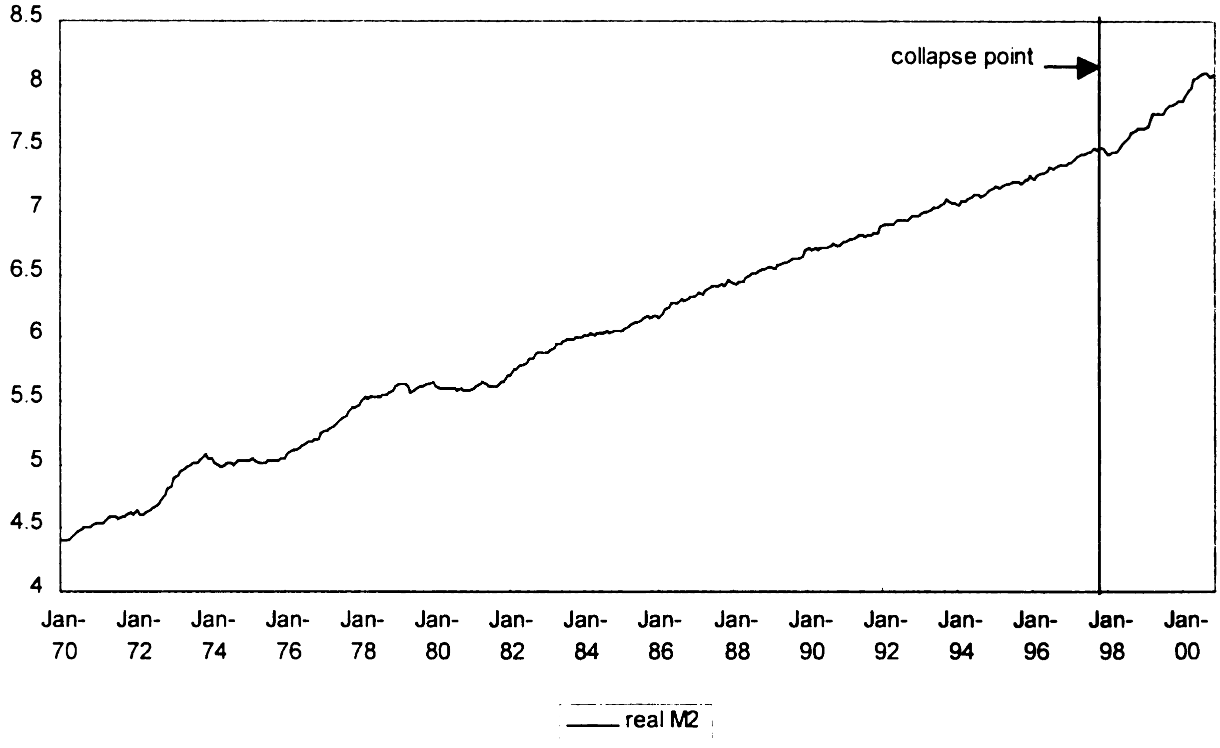


Figure 44. The log of domestic credit (South Korea)

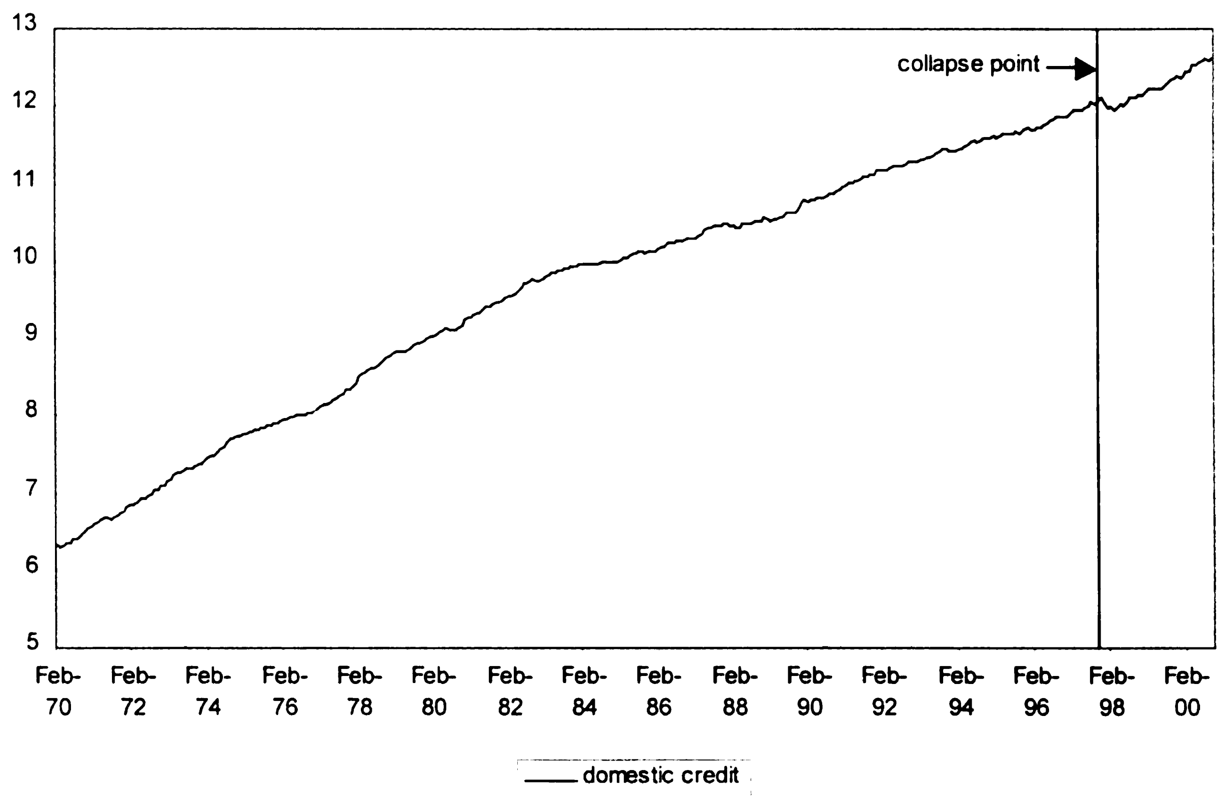


Figure 45. The log of real GDP (South Korea)

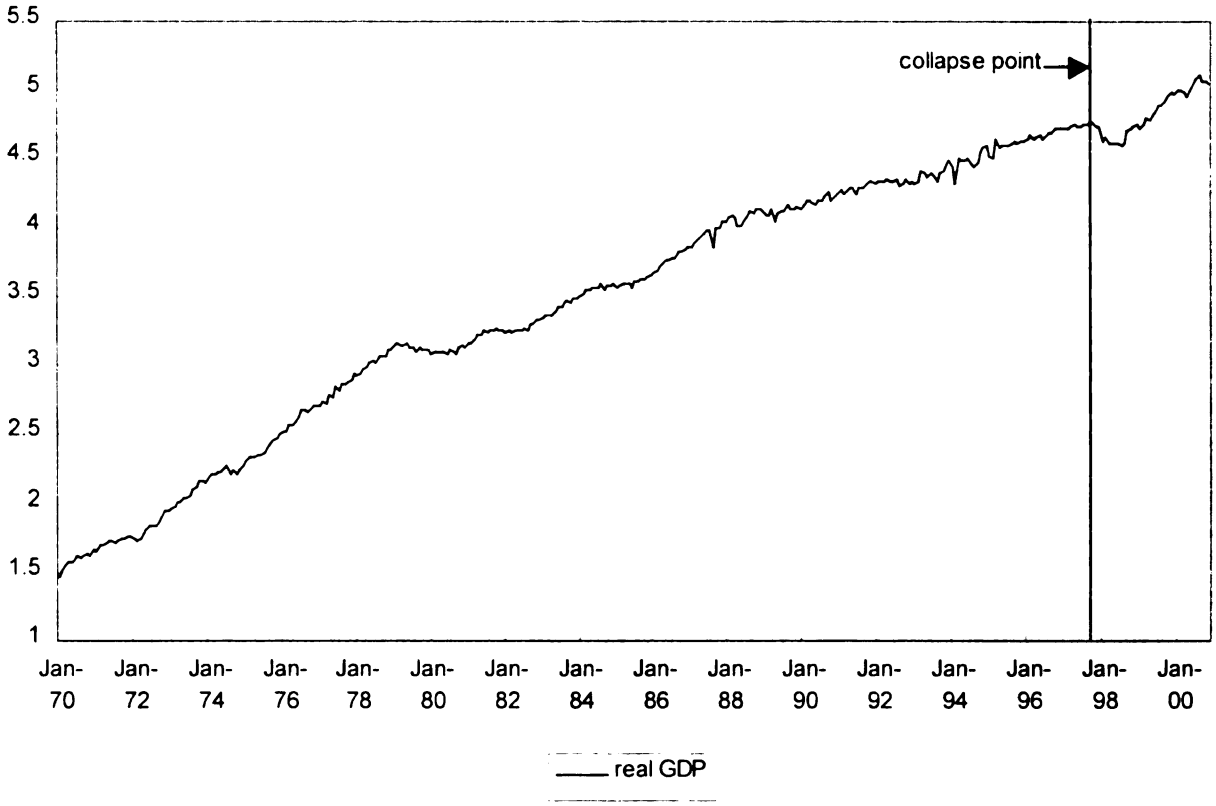


Figure 46. Interest rate (South Korea)

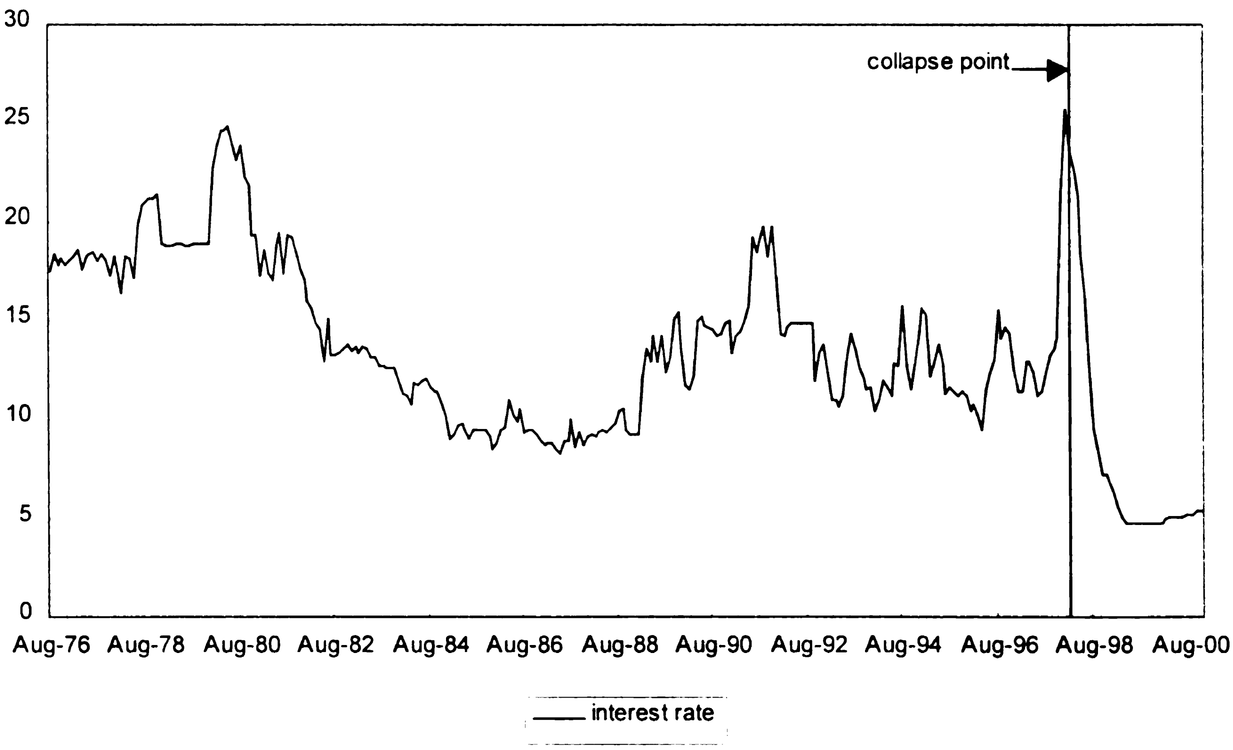


Figure 47. Deviation from PPP (South Korea)

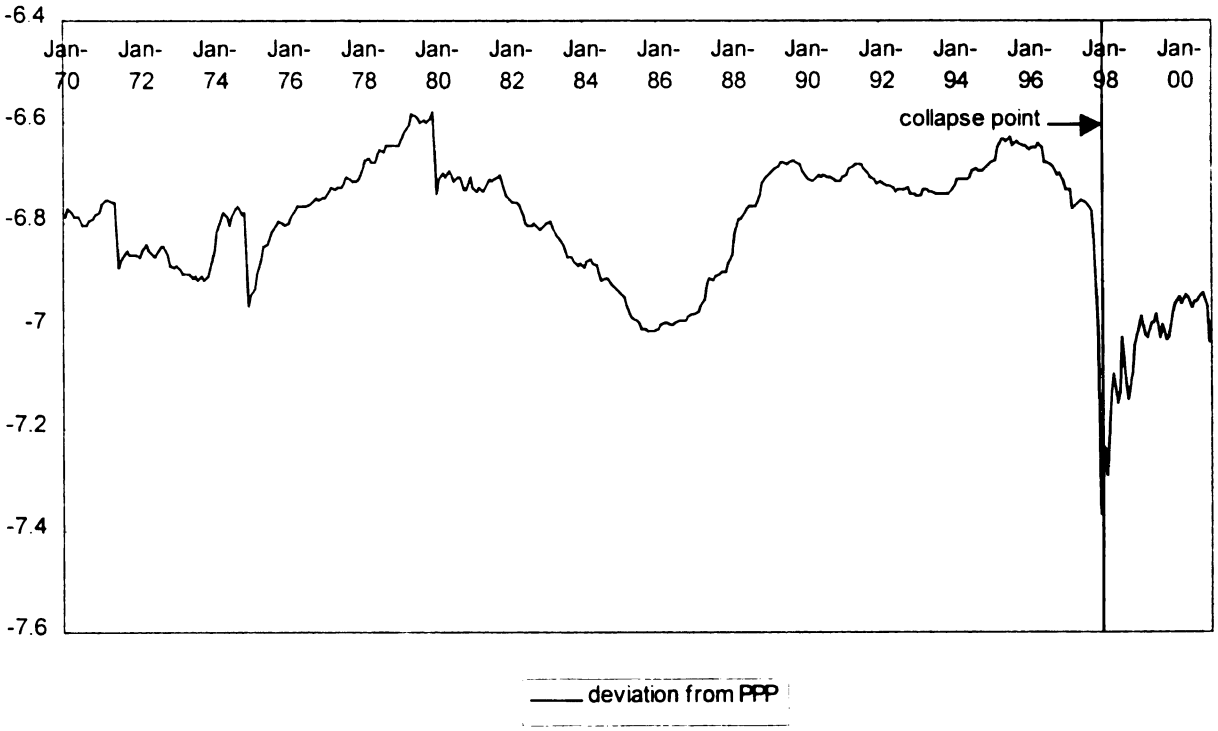


Figure 48. The log of real M2 (Malaysia)

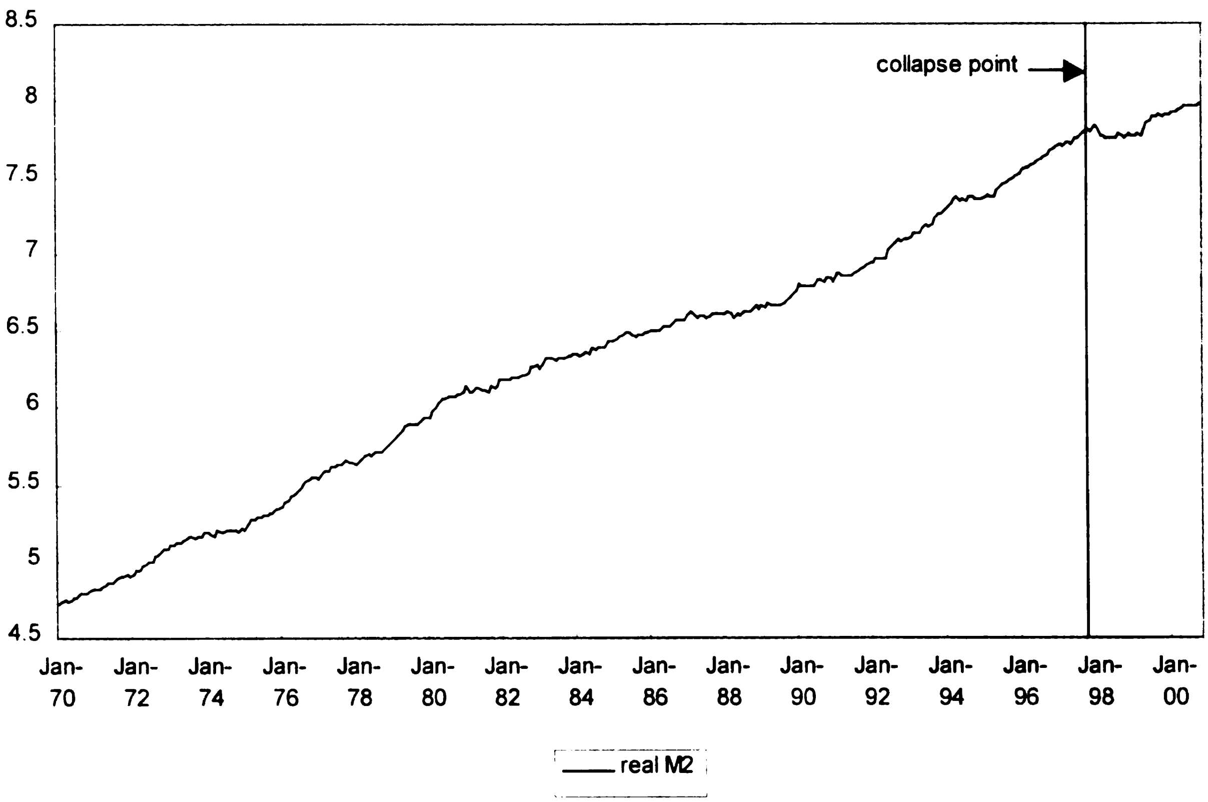


Figure 49. The log of domestic credit (Malaysia)

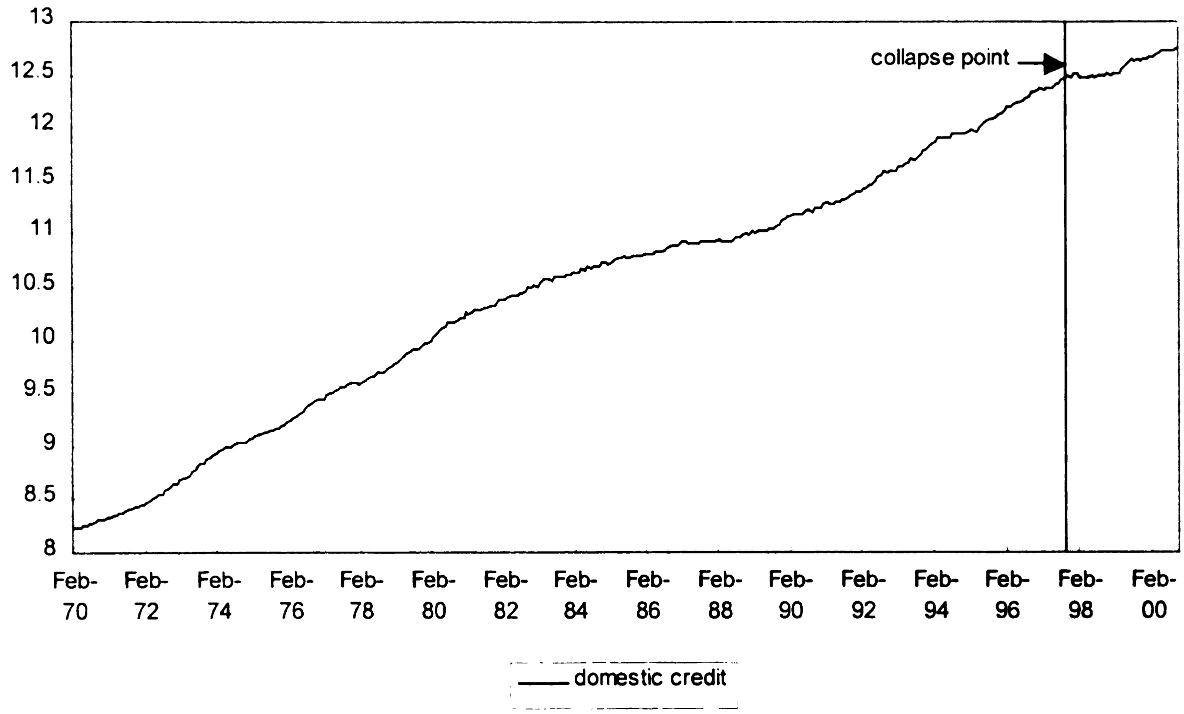


Figure 50. The log of real GDP (Malaysia)

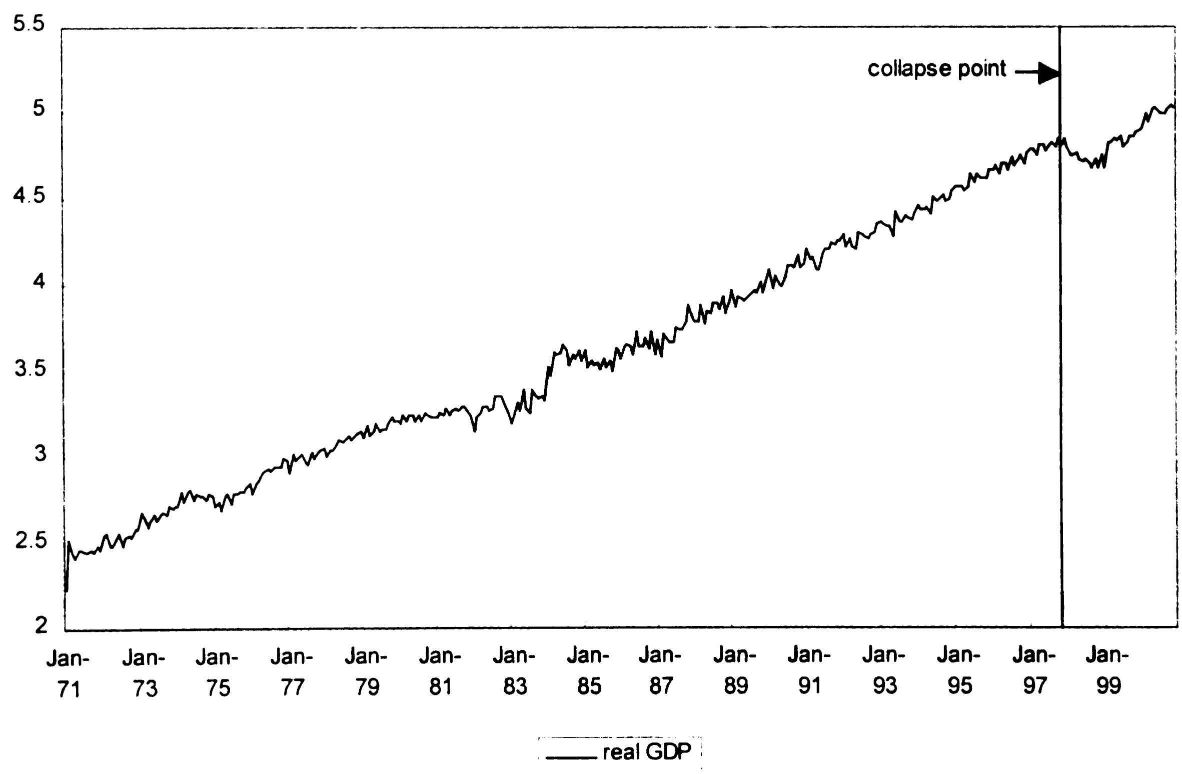


Figure 51. Interest rate (Malaysia)

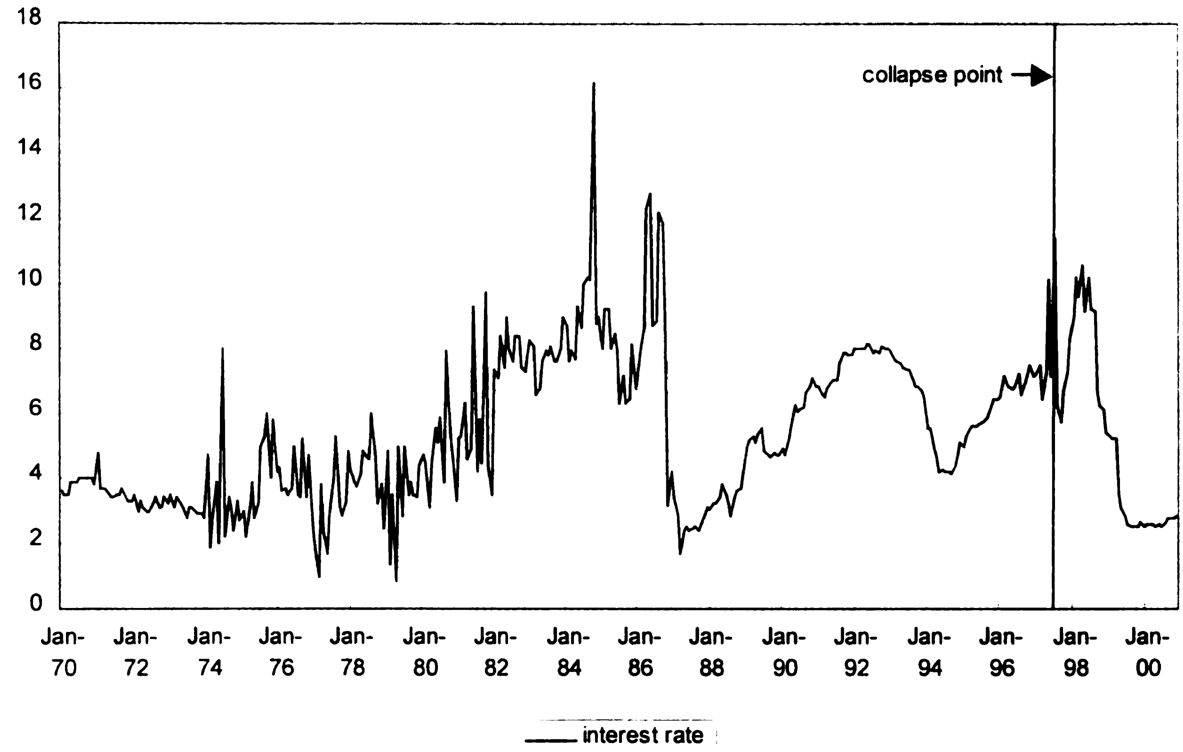


Figure 52. Deviation from PPP (Malaysia)

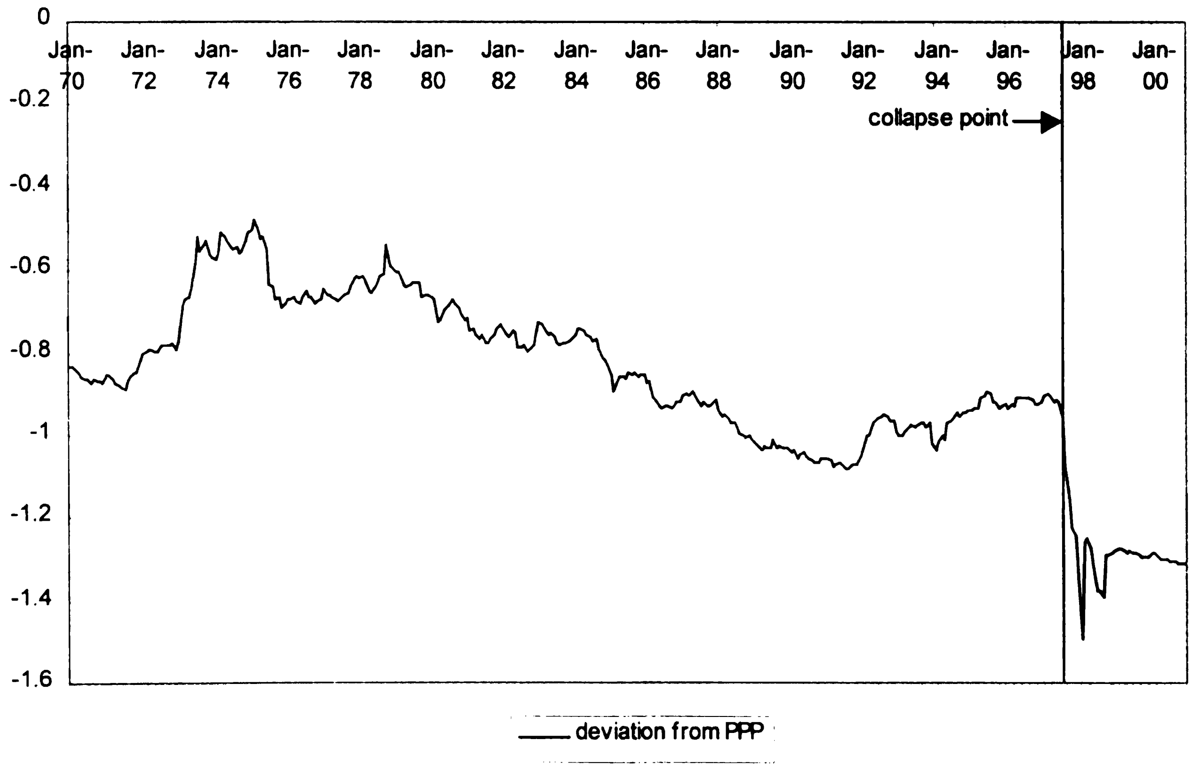


Figure 53. US Interest rate

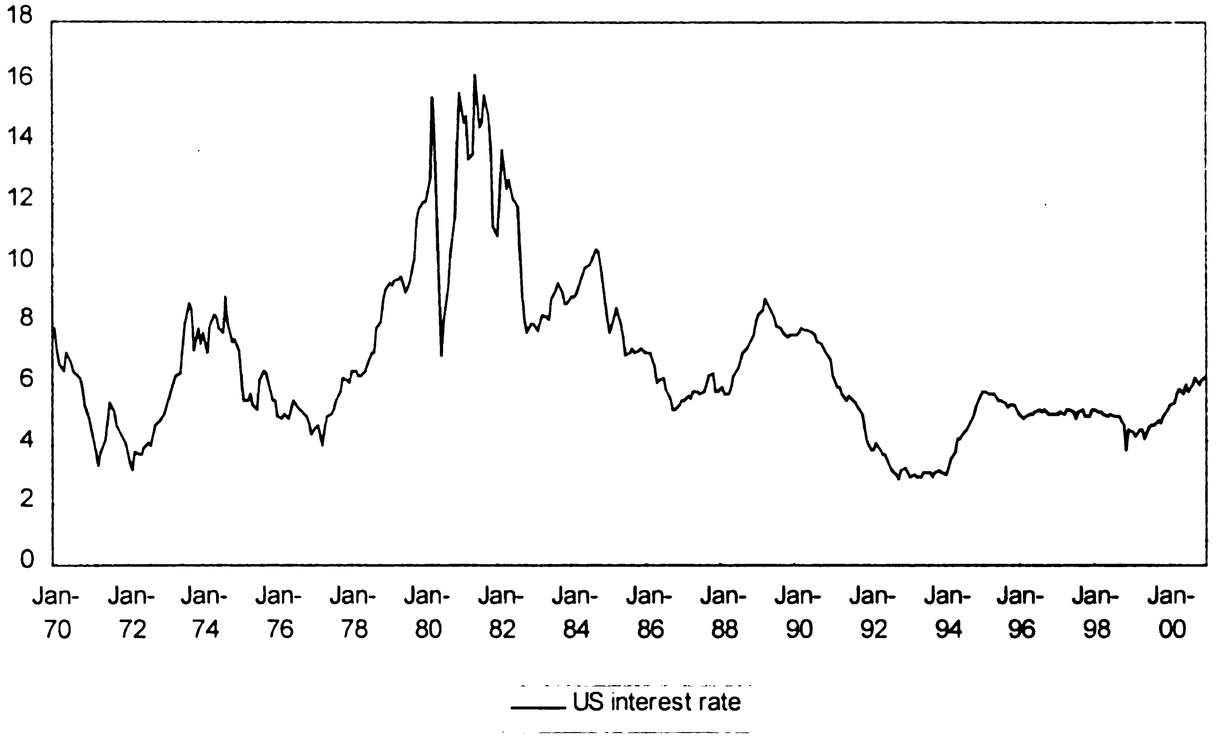


Figure 54. US CPI

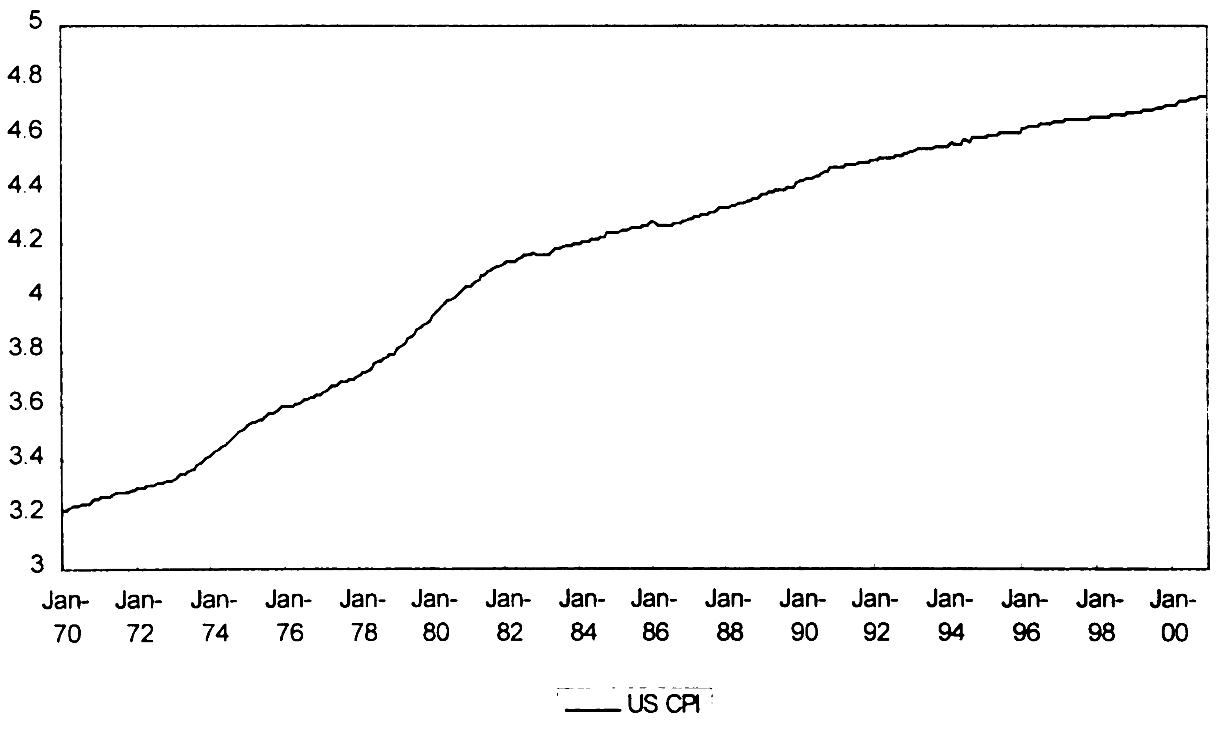


Figure 55. The actual and shadow exchange rates (South Korea)

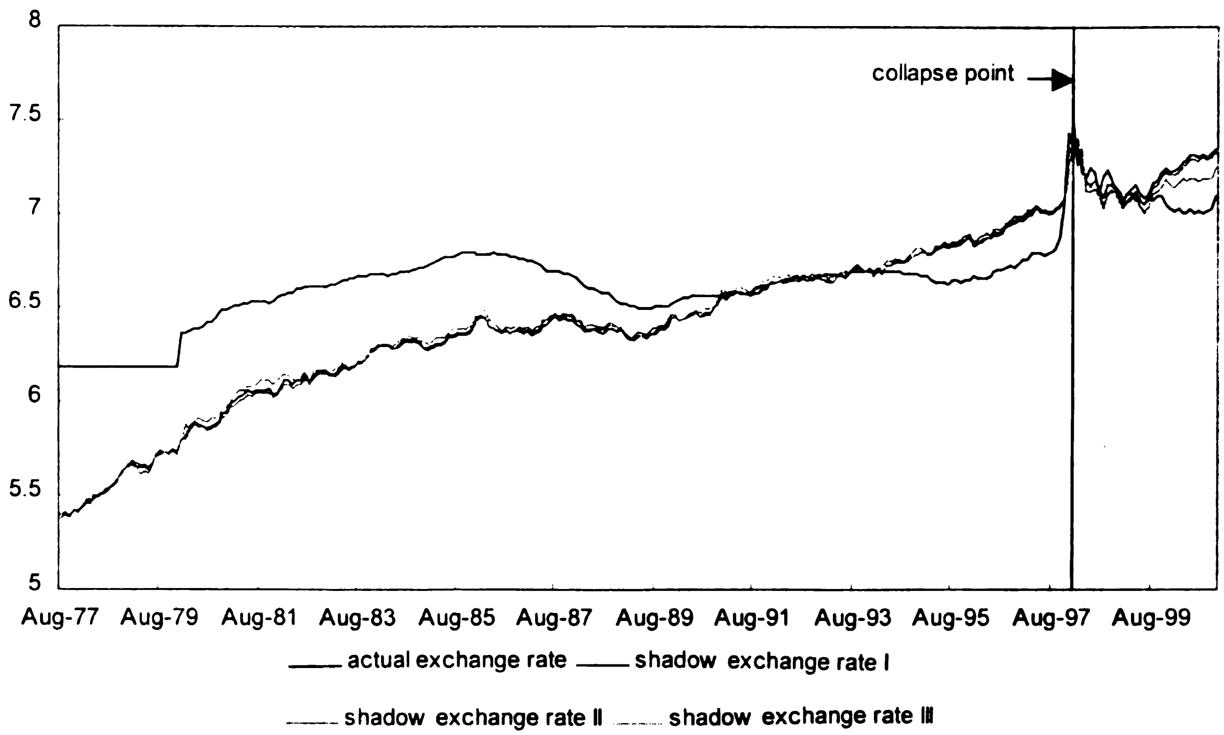


Figure 56. The actual and shadow exchange rates (Malaysia)

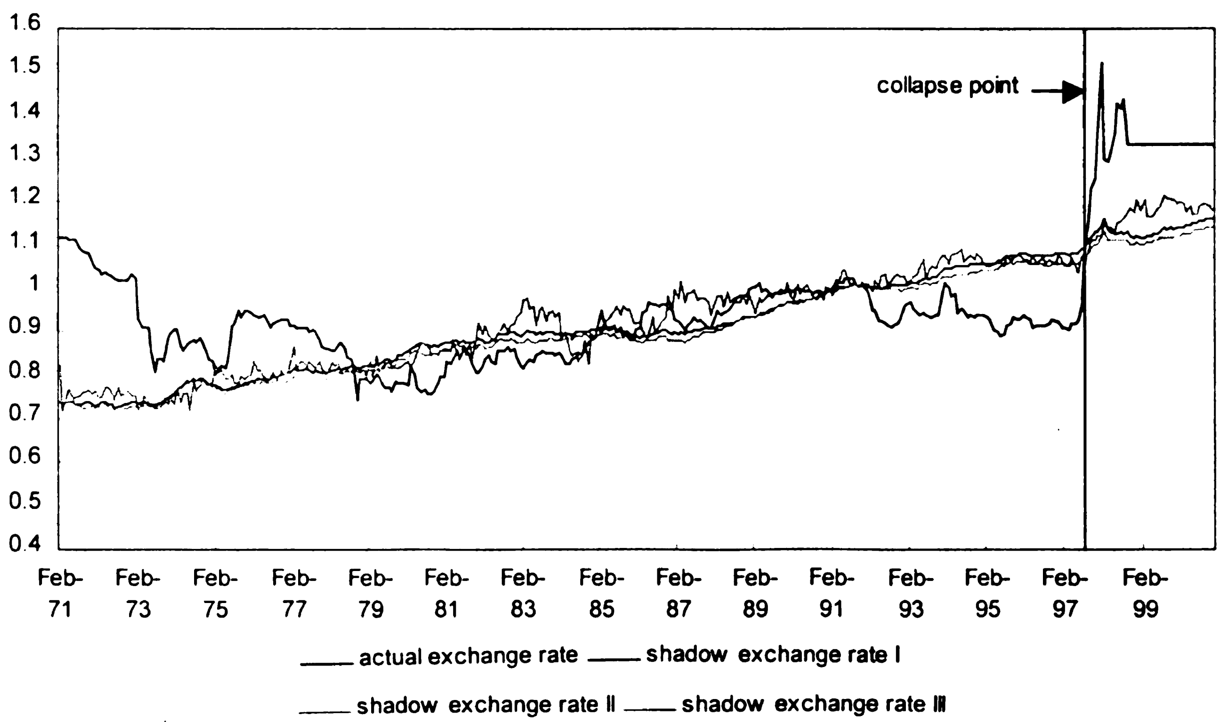


Figure 57. The probability of collapse I (South Korea)

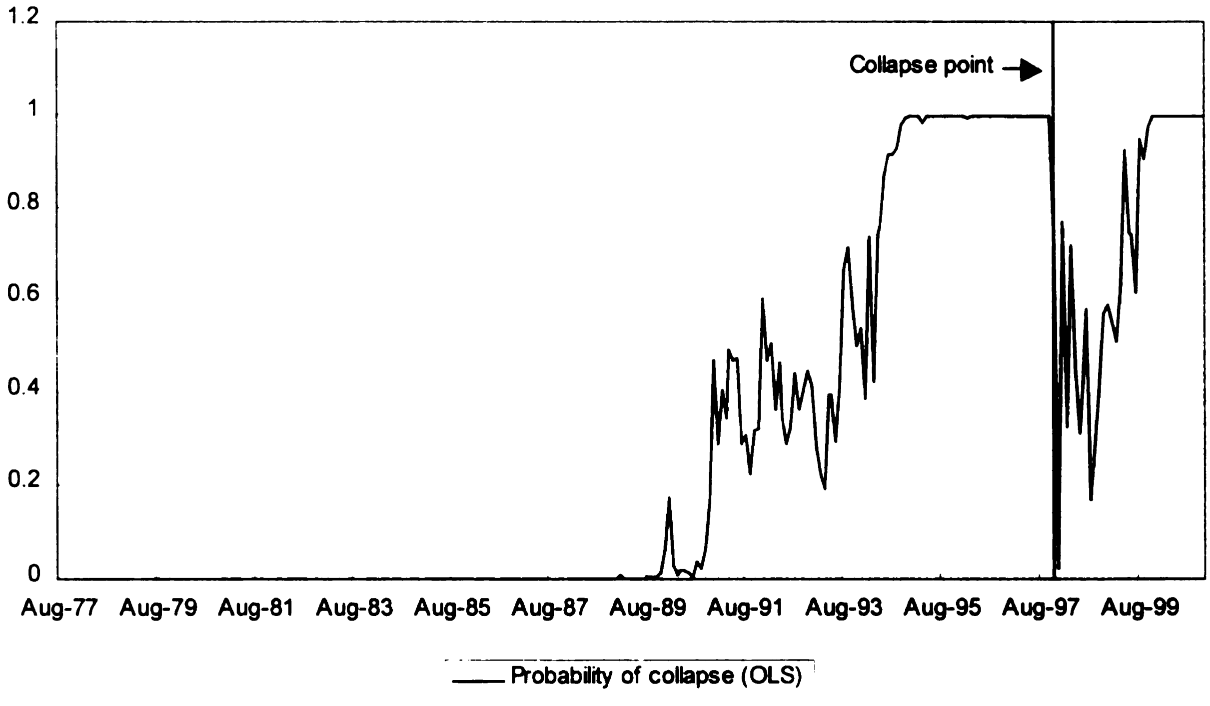


Figure 58. The probability of collapse II (South Korea)

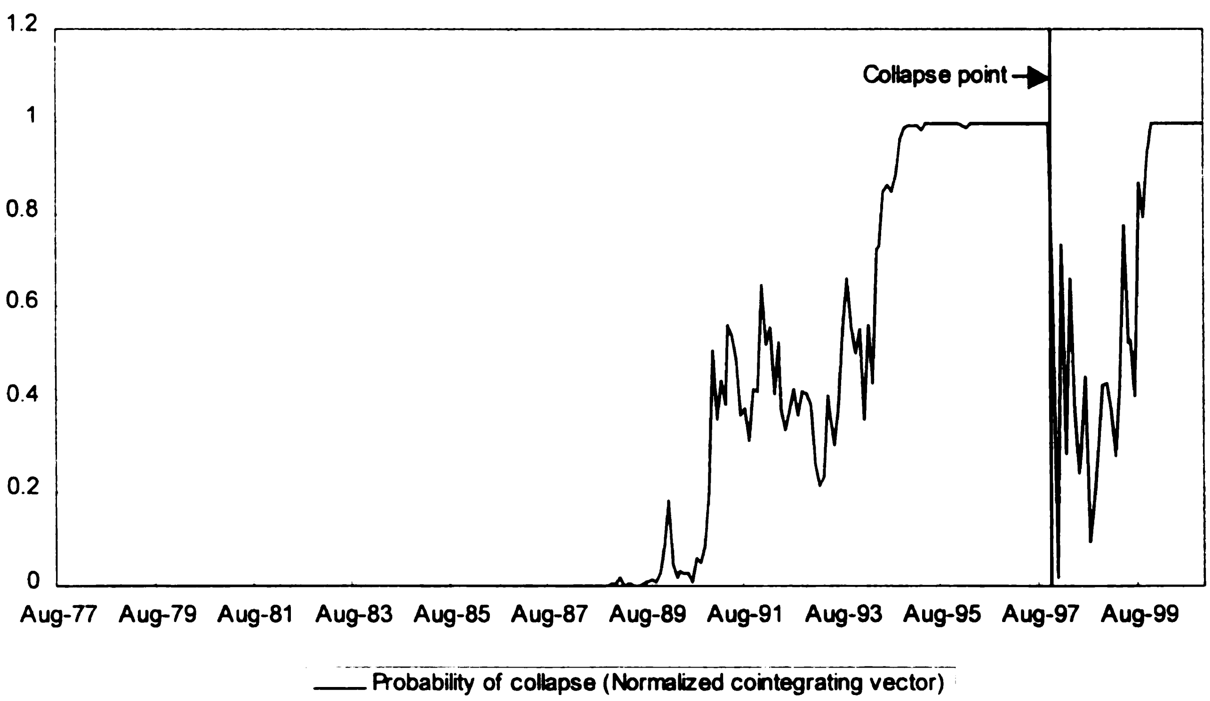


Figure 59. The probability of collapse III (South Korea)

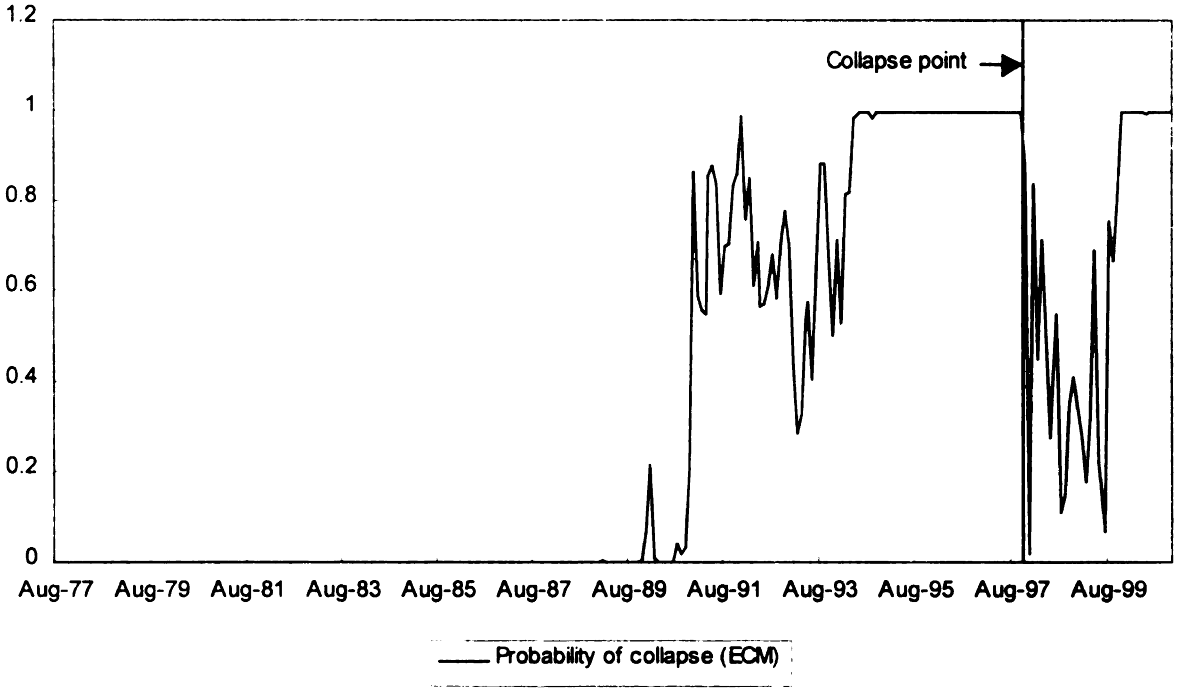


Figure 60. The probability of collapse I (Malaysia)

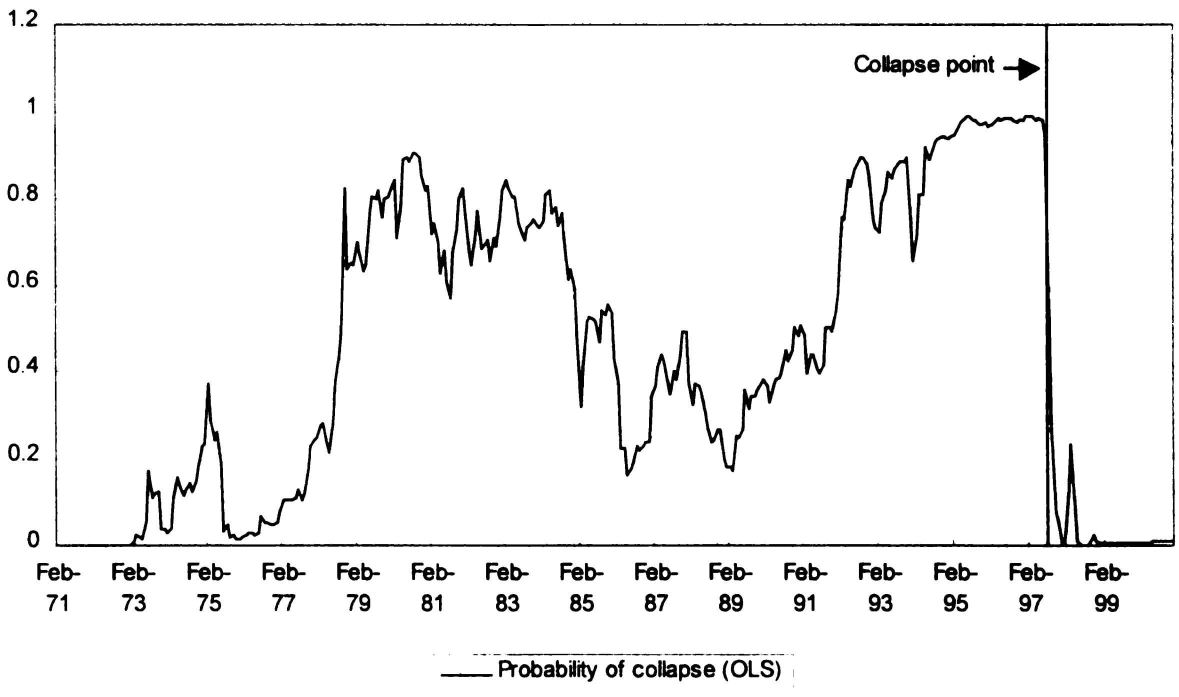


Figure 61. The probability of collapse II (Malaysia)

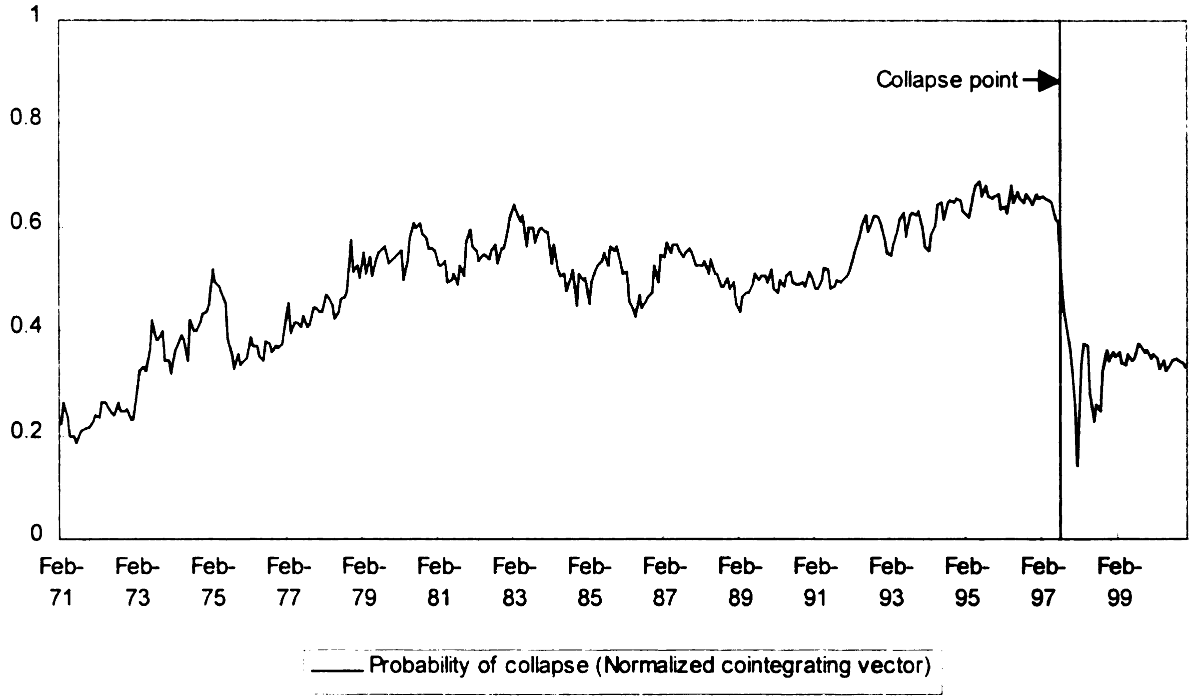
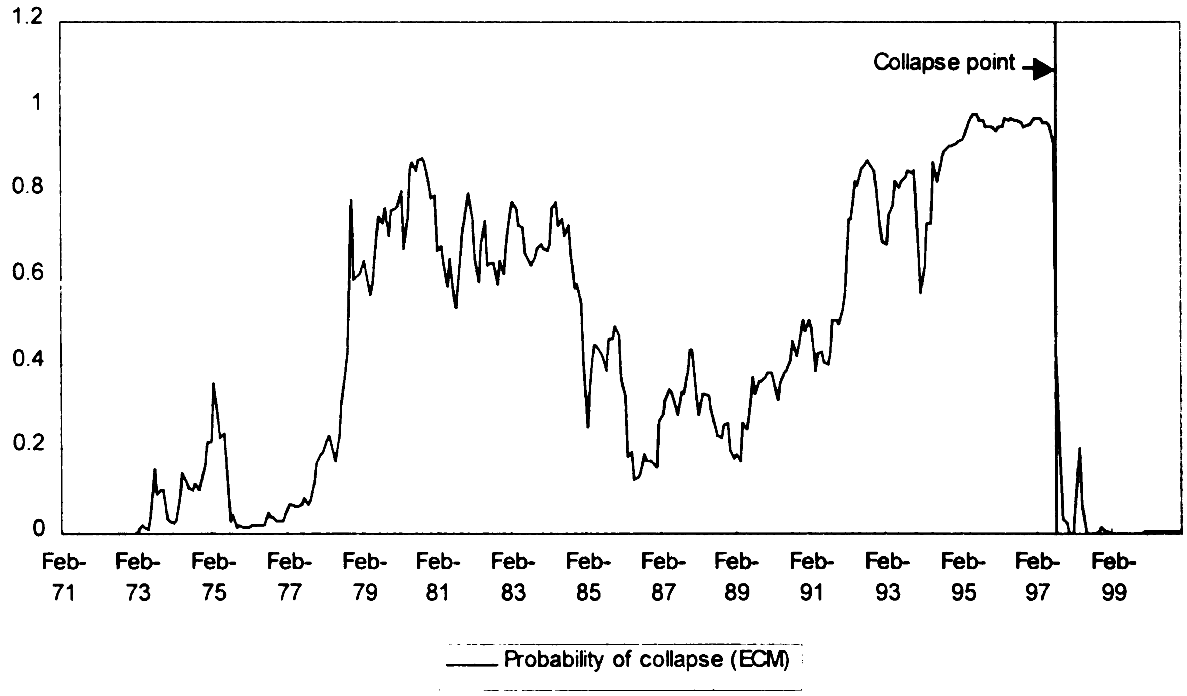


Figure 62. The probability of collapse III (Malaysia)



CHAPTER VII

COMMON FUNDAMENTALS AND CONTAGION EFFECT IN CURRENCY CRISIS

1. Introduction

While the traditional approach to currency crisis stresses the decline in international reserves leading to a collapse of a fixed exchange rate, more recent models focus on additional variables and the possibility of a contagion effect.

In the aftermath of the 1994 Mexican crisis and the 1997 Asian crisis there was a widespread contagion between several emerging markets. Clearly, there may be numerous reasons to expect contemporaneous crises. First, there may be common external factors. An example would be how policies undertaken by industrial countries may have similar effects on emerging markets. This is frequently called the "Monsoon effect". For example, a rise in U.S. interest rates in 1994 or the devaluation of the yen in 1995 would be common external factors. Second, a crisis in an emerging market may affect the macroeconomic fundamentals in other emerging markets. This is usually caused by trade linkages or spillovers related to third market competition. The third factor of the contagion effect is market sentiment. The herd mentality of investors could explain part of the contagion. If investors pay little heed to countries' economic fundamentals, fail to discriminate properly among countries, a crisis in a neighboring country may threaten a future domestic crisis.

Sachs, Tornell and Velasco (1996) and Tornell (1999) seek to identify macroeconomic variables that can help explain which countries were vulnerable to "contagion effects" following the Mexican crisis and the Asian crisis. In addition, Corsetti, Pesenti and Roubini (1998b), and Furman and Stiglitz (1998) try to explain the spread of the Asian crisis by finding the common fundamentals for the eruption of the crisis. Glick and Rose (1998) find that countries with important trade linkages to the country that first experienced a crisis were more likely to experience a crisis. Masson (1998) suggests that the contagion effect is unexplained by the common external effects and that trade linkages played a major role in the Mexican and Asian crises. Table 55 summarizes the findings of 7 selected empirical studies on currency crisis that focus on the contagion effect.

The first objective of this chapter is to see whether the macroeconomic variables that explain the cross-country variation in the severity of the crises in Mexico and Asia generalizes to countries other than emerging market countries. To this end, this study looks at whether the estimated coefficient in each macroeconomic variable is still significant in a regression taken from sample of 29 industrialized and developing countries covering years of the European, Mexican and the Asian currency crises.

The second objective is to examine the extent of contagion effect in all aspects of common external effects, trade linkages and market sentiment. For common external effects, a real exchange rate reflecting the devaluation of yen before the Asian crisis and an interest rate differential indicating the high interest rate of U.S. before the Mexican crisis are controlled for in the model. To control for trade linkage, a trade linkage index, which captures the degree to which the initially attacked country and the home country compete in other markets and the degree to which the two countries directly trade with each other, is added to the model. To control for market sentiment, a dummy variable is included in the model.

The third objective is to predict the currency crisis based on the contagion effect as well as weak fundamentals. To this end, I check for which countries were vulnerable to speculative attack without contagion effects prior to the Asian currency crisis based on the result of prediction of a crisis using estimated coefficients from the Mexican currency crisis in the benchmark regression. Then, the coefficients of variables in the benchmark regression with all the contagion effect variables are estimated using the countries chosen in the first step as the initially attacked countries of the contagion instead of Mexico. Finally, each country's vulnerability before the Asian crisis is predicted using the coefficients of variables for the contagion effect and other variables estimated in the second step.

2. Theoretical framework

To respond to a speculative attack which requires a large supply of foreign exchange in a market, a country runs down reserves; increases its interest rate; and depreciates its currency. The first option may be the least costly politically, but it is an option only available to governments with sufficient reserves to respond to an attack. As such, a country whose short-run liabilities exceed by far their reserves must choose between monetary contraction and currency depreciation. Thigh monetary policy that increases the domestic interest rate makes speculation attack against the currency more costly in the short run. However, those effects may come at the cost of recession. The extent of recession tends to be severe with rapid lending boom. When the banking system has a large share of bad loans because of a lending boom, a higher interest rate leads to a full-scale banking crisis. The existence of both low reserves and a weak banking system may force the government to close the external imbalance through currency depreciation. The more currency has previously appreciated, the more the government should depreciate the currency. This is because it is more likely that firms in the tradable sector have moved to the non-tradable sector. The movement between sectors then lowers the response of tradable sector to a real depreciation.

Therefore, the countries with low reserves, a rapid lending boom and a severe real appreciation are more likely to face a speculative attack. In addition, trade linkages with an initially attacked country or a differential between the domestic and foreign interest rates may signal to speculators the increased likelihood of currency crisis.

2.1 A simple model

Consider an open economy where there are many identical investors who initially hold an aggregate stock M of deposits denominated in domestic currency that pay an interest rate i. The model is static, with the focus on the interaction between an investor's expectation of devaluation and the government's management of the external account in the very short run.

In the model, each investor initially selects the stock of domestic deposits she wishes to hold and the amount she hopes to convert into foreign currency. Then, the government responds to the capital outflow by running down its reserves, increasing interest rate, or depreciating the country's currency. Finally, investors cash their deposits plus the interest accrued.

A risk neutral investor will hold domestic deposits so long as $\frac{1+i}{1+\dot{s}_j^e} \geq 1+i^*$, where i^* denotes the foreign interest rate and $\dot{s}_j^e$ denotes the devaluation rate expected by investor j. In other words, investor j will hold domestic deposits only if the expected devaluation rate is no greater than the threshold value,

$$\bar{\dot{s}} = \frac{i-i^*}{1+i} \quad . \tag{1}$$

Under this assumption, each investor, who initially holds a stock m of deposits, can either continue to hold the deposits or withdraw everything. Hence, an investor j's strategy would be

$$\Delta m_j^d = \begin{cases} 0 & \text{if } \dot{s}_j^e \leq \bar{\dot{s}}(i) \\ -m & \text{if } \dot{s}_j^e > \bar{\dot{s}}(i) \end{cases} \tag{2}$$

where $\bar{\dot{s}}'(i) > 0$. In this model, an increase in the interest rate i will make it more likely that $\dot{s}_j^e$ is less than $\bar{\dot{s}}$. In a symmetric equilibrium, all investors derive the same conclusion from this common information. Thus, the change in aggregate deposits ΔM^d is equal to either $-M$ or 0, where M denotes the aggregate initial stock of deposits.

The government has an initial stock R of international reserves. By taking the behavior of investors, ΔM^d, as given, the government chooses the change in reserves

ΔR, the depreciation rate $\dot{s}$, and the unemployment rate u, to minimize the following social loss function[33] (3), subject to equations (4), (5), (6) and (7).

$$\min_{\Delta R, \dot{s}, u} \ (u + \alpha \dot{s}) \tag{3}$$

$$CA + \frac{\Delta M^d}{1 + \dot{s}} = \Delta R, \quad \Delta R \geq -R \tag{4}$$

$$CA = \varphi(rer)\dot{s} + u(wf) - F(rer) - T(trade), \quad \varphi' > 0, F' < 0 \tag{5}$$

$$0 < u < \bar{u}(wf), \quad \bar{u}'(wf) < 0 \tag{6}$$

where CA = current account, rer = real exchange rate, wf = weakness of financial system and trade = trade linkage with the initially attacked country[34].

Equation (3) says that the government minimizes the sum of the depreciation rate and the unemployment rate, but does not care about the changes in reserves. The parameter α captures how sensitive the government is to nominal depreciation. Equation (4) is the identity linking the current account balance, CA, and the capital account, $\frac{\Delta M^d}{1 + \dot{s}}$, to the changes in reserves. As shown in equation (5), the current account balance is positively affected by nominal depreciation and unemployment but negatively related to the real appreciation, $F(rer)$. The term, $- F(rer)$, captures the negative effect of today's service on the debt associated with past current account deficits caused by previous real appreciation. The coefficient $\varphi(rer)$ indicates how effectively a normal devaluation would improve the current account. The more the real exchange rate appreciated, the

[33] This social loss function is quoted from Obstfeld (1997).

[34] We assume that there is a country which was already attacked by the speculative agents and have been in currency crisis.

lower φ. The term $-T(trade)$ reflects the negative effects from the trade linkage with the initially attacked country. The more the currency of the initially attacked country depreciates, the weaker the home country's competitiveness in the international goods market will be. The existence of an upper bound on the unemployment rate, $\bar{u}(wf)$, in equation (6) captures the idea that one cannot indefinitely increase unemployment without causing bankruptcies and a melt down of the payment system. The share of bad loans in the banks' portfolio caused by a lending boom is one indicator of the weakness of a financial system.

There are three possible solutions to the government's problem, depending on the size of the reserves. First, if international reserves are sufficient to cover any potential capital outflow plus the current account deficit or $R \geq \Delta M^d + F(rer) + T(trade)$, the government will be able to close the external gap by spending its reserves. That is,

$$\begin{aligned} \Delta R^* &= \Delta M^d + F(rer) + T(trade) \\ \dot{s}^* &= 0 \\ u^* &= 0 \end{aligned} \tag{7}$$

Secondly, when reserves may cover the current account deficit, but are not sufficient to cover both the current account deficit and the potential withdrawal of deposits or $F(rer) + T(trade) \leq R \leq \Delta M^d + F(rer) + T(trade)$, then, a government's policy will depend on investors' expectations of devaluations. If $\dot{s}^e \leq \bar{\dot{s}}$, then investors will not attack the currency. Thus, $\Delta R^*, \dot{s}^*, u^*$ are the same as in (7). If $\dot{s}^e > \bar{\dot{s}}$, investors will attack the currency. Since the government prefers to close the external imbalance by spending its reserves, it will wait until $R = 0$ before pursuing the alternatives when reserves are not sufficient to close the external gap. Then, the deficit not covered by reserves must be closed by either depreciation or a recession. Let Φ be the unemployment rate chosen by the government given the existence of an upper limit of the unemployment rate. When Φ is less than the maximum feasible unempolyment rate ($\Phi \leq \bar{u}$), the choices of unemployment and devaluation are given by (9). However, when $\Phi > \bar{u}$, Φ is not feasible any more. Then, as shown in (10), the government sets unemployment at $\bar{u}$, and needs to close the external deficits through additional depreciation than would be the case if $\Phi \leq \bar{u}$.

$$\Delta R^* = -R \tag{8}$$

$$\begin{cases} \pi^* = \sqrt{\dfrac{M}{\varphi - \alpha}} - 1 & \text{if } \Phi \leq \bar{u} \\ u^* = \Phi \end{cases} \tag{9}$$

$$\begin{cases} \pi^* = \dfrac{-(\varphi + Z) + \sqrt{(\varphi + Z)^2 - 4\varphi(M + Z)}}{2\varphi} & \text{if } \Phi > \bar{u} \\ u^* = \bar{u} \end{cases} \tag{10}$$

where $Z \equiv \bar{u} + R - F - T$ and $\Phi \equiv \dfrac{\varphi(\sqrt{\varphi - \alpha} - 2\sqrt{M}) + \alpha}{\sqrt{\varphi - \alpha}} + F + T - R$.

Third, when reserves are too low to cover the current account deficit or $R < F(rer) + T(trade)$, then, reserves will not be sufficient to close an external gap regardless of the fluctuations in the demand for money, ΔM^d. In this scenario, the government will deplete its reserves and then close the external gap through a combination of depreciation and manipulation of the unemployment rate as expressed in equations (9) or (10).

The symmetric rational expectation's equilibria are found by combining the investor's withdrawal policy (2) and the government strategies from (7) to (10). There are three cases. First, if $\dot{s}^*(-M) \leq \bar{s}$, then there is a unique symmetric equilibrium where

$$\Delta \hat{M}^d = 0 \text{ and } \hat{s} = \dot{s}^*(0).$$

Second, if $\bar{s} \in [\dot{s}^*(0), \dot{s}^*(-M)]$, then there are two symmetric equilibria,

$$\Delta \hat{M}^d = 0 \text{ and } \hat{s} = \dot{s}^*(0)$$

$$\Delta \hat{M}^d = -M \text{ and } \hat{s} = \dot{s}^*(-M).$$

Third, if $\bar{s} < \dot{s}^*(0)$, then there is a unique symmetric equilibrium,

$$\Delta \hat{M}^d = -M \text{ and } \hat{s} = \dot{s}^*(-M).$$

In the first case, an attack never occurs since either reserves are high or fundamentals are strong.[35] When reserves are high, the government will respond to any $\Delta \hat{M}^d$ by spending its reserves and setting $\dot{s} = 0$. The second and third cases occur when reserves are low and fundamentals are weak. In the second case, there are multiple equilibria. In the crisis equilibrium, investors believe that the devaluation will be greater than $\bar{s}$ and consequently withdraw their deposits. As a result of the withdrawal, the devaluation is indeed greater than $\bar{s}$. In the non-crisis equilibrium, investors believe do not withdraw their deposits and depreciation is not greater than $\bar{s}$. In the third case, the fundamentals are so weak that the government will have to depreciate more than $\bar{s}$ regardless of investor`s expectations.

In sum, an individual money manager will attack a currency only if it is anticipated that the country will respond with a sizeable depreciation. A sizable depreciation is more likely to occur in countries with low international reserves, a severe current account deficit and a rapid lending boom.

3. Empirical analysis

The theoretical model in the previous section suggests that the countries most vulnerable to a speculative attack have a severe real appreciation, a strong trade linkage with the country initially attacked by currency speculation, a rapid lending boom and low international reserves. Investors concentrate their speculative attacks in countries more likely to respond with an excessive depreciation.

This section shows that the European, Mexican and Asian currency crises did not spread randomly across industrial and emerging markets during the years 1993, 1995 and 1997. In addition, the extent of role of other determinants of currency crisis is examined. These include high government consumption, slowdown in real GDP growth rate, excessive capital inflows and increasing foreign liabilities.

3.1 Defining variables in the empirical model

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[35] There are neither many bad loans nor a sever current account deficit due to a real appreciation or a strong trade linkage with an initially attacked country.

Currency crisis index

The first issue confronted in the analysis is how to measure devaluation pressures on the foreign exchange market. Eichengreen, Rose, and Wyplosz (1996), Sachs, Tornell and Velasco (1996), Frankel and Rose (1996) and Kaminsky and Reinhart (1999) used analogous crisis indices for the measurement of pressures on the market. Their indices are weighted averages of the percentage of depreciation in the nominal exchange rate with respect to the U.S. dollar and the percentage decrease in reserves. The rationale for these indices is as follows. If capital inflows reverse, then the government can depreciate the exchange rate. Alternatively, it can defend the currency by spending its reserves or increasing interest rates. Since the authors assert that there is no reliable and comparable cross-country interest rate data, their indices are constructed without an interest rate.

The extent of currency crisis here is measured with a crisis index (denoted "*Crisis*"),

$$Crisis_{it} = w_{\Delta s}\%\Delta s_{it} - w_{\Delta R}\%\Delta R_{it} \quad (11)$$

where $\%\Delta s_{it}$ and $\%\Delta R_{it}$ are percentage change of the nominal exchange rate and the international reserves respectively. The weights used to derive the crisis index are constructed as

$$w_{\Delta s} = \frac{\frac{1}{\sigma_{\Delta s}}}{\frac{1}{\sigma_{\Delta s}} + \frac{1}{\sigma_{\Delta R}}} \text{ and } w_{\Delta R} = 1 - w_{\Delta s},$$

where $\sigma_{\Delta s}$ and $\sigma_{\Delta R}$ are the standard deviation of the change rates of nominal exchange rate and international reserve. The initial point for the percentage change is the month before the onset of the crisis (August 1992, November 1994 and June 1997). Then, the terminal month is varied over a period of six months starting in October 1992, January 1993, and August 1997. This index is similar to indices used in the extensive prior literature. However, the standard deviation of the change rate of the variable instead of the level to preclude an excessively small weight being given to the international reserves due to its unusually high volatility compared to that of a nominal exchange rate. In addition, unlike Kaminsky and Reinhart (1999) who assign the weight of the nominal exchange rate to be one, $w_{\Delta s}$, which is less than or equal to one, is used to weight the change of nominal exchange rate to prevent an excessively large weight being given to the nominal exchange rate. The values of *Crisis* are listed in Table 42. A higher value of *Crisis* means either higher level of devaluation or a greater fall in reserves. With the exception of Brazil in 1993, all of the countries that experienced a currency crisis in 1993, 1995 or 1997 have higher a crisis index than other countries.

Lending boom

A broad cross-country set of comparable bank balance sheets does not exist. Hence the weakness of the banking sector cannot be assessed directly by comparing the ratio of non-performing loans to total assets. Instead, an indirect measure of financial system vulnerability is used: the magnitude of the increase in bank lending as measured by the percentage change in the ratio of claims on the private sector by deposit money banks and monetary authorities (line 32d) to GDP (line 99b) during the periods 1988-92, 1990-94 or 1992-96. This variable is *LB* and its values are listed in Table 43. The first column of the table shows the *LB*s for European currency crisis. Unlike as was the case with the Mexican and Asian currency crisis in the second and third columns, the *LB*s values in European countries in the first column are not in the high rankings. This may imply that lending booms did not play so crucial of a role in the currency crises of industrial countries relative to the role it played in the currency crises of the developing countries.

Real exchange rate depreciation

A real exchange rate depreciation index is constructed as a weighted average of the bilateral real exchange rates of a given country with respect to the US dollar, the Yen and the Mark. The weights add up to one and are proportional to the shares of bilateral trade in the given country with the US, Japan and Germany, respectively. The extent of real exchange rate misalignment is then measured with the percentage change in this index over the four years prior to the onset of the crisis. This variable is *RER*. A positive value of *RER* signifies that the real exchange rate depreciated relative to the base period, while a negative value indicates appreciation. Table 44 offers the values of *RER* for the currency crises in 1990s. The visual inspection of the first column in the table also shows that the negative relationship between the currency crisis and the *RER* is weaker in the European currency crises, although the rankings of Italy and Spain in 1992 are higher than others in 1994 and 1996.

Reserves adequacy

The government's liquidity is proxied by the ratio of M2 to reserves in the month preceding the onset of the crisis (August 1992, November 1994 or June 1997). The ratio captures the extent to which the liabilities of the banking system are backed by international reserves. If the central bank is unwilling to allow the exchange rate to depreciate, then it must be prepared to cover all the liabilities of the banking system with reserves. Hence, it is *M2*, and not simply the monetary base, that is the relevant proxy for the central bank's contingent liabilities. The values of the ratio of *M2* to reserves for both industrial and developing countries are listed in Table 45 and 46, respectively. The industrial countries' reserve adequacies in Table 45 are excessively lower than the developing countries' in Table 46. This implies that a country whose financial market is well developed and stable is allowed to maintain a larger monetary base than emerging market countries, who are more likely to be exposed to unexpected speculative attack. However, the reserve adequacy of Italy in 1992, Mexico in 1994 and Thailand in 1997, when they were initially attacked, is lower than the reserve adequacies of other countries in their group of sample.

Contagion effects

Contagion effects are the most recent contribution of second-generation models. There are several channels through which they may be transmitted across countries. First, contagion can be explained by common external factors, so called "Monsoon effect". For example, a rise in U.S. interest rates in 1994 or the devaluation of the yen in 1995 could be common external factors. Second, it is also caused by trade linkage or third market competition-related spillovers. The trade linkages between countries with geographic proximity help to explain spillover effects. In addition, an indirect trade linkage due to third market competition may be instrumental in encouraging repeated rounds of competitive devaluation. The third factor of the contagion effect is market sentiment. The herd mentality of investors also may contribute to the contagion effect.

For the common external factors of contagion, this section considers the contagion effect from the devaluation of the major currencies as reflected in the change of the real exchange rate depreciation index, *RER* and a decline in the differential between domestic and U.S. interest rates, *Itrdus*. To control for the trade linkage or third market competition-related spillovers, a trade linkage index (denoted "*Trade*") between the initially attacked and home country is constructed following the same method used in Glick and Rose (1998). *Trade* is a weighted average index between the third market competition index,

$$Indirect = \sum_k \{[\frac{x_{0k} + x_{ik}}{x_{0.} + x_{i.}}] * [1 - \frac{|x_{ik} - x_{0k}|}{x_{ik} + x_{0k}}]\}$$

and direct trade linkage index,

$$Direct = 1 - \frac{|x_{i0} - x_{0i}|}{x_{i0} + x_{0i}}$$

where x_{ik} denotes aggregate bilateral exports from country i to k ($k \neq i, 0$); x_{i0} denotes aggregate bilateral export from the home country i to the initially attacked country 0; $x_{i.}$ denotes aggregate bilateral exports form country i. *Indirect* is the weighted average of the importance of exports to country k for countries 0 and i. The relative importance of country k is strongest when it is an export market of equal importance to both countries 0 and i. The weights are proportional to the importance of country k in the aggregate trade of countries 0 and i. *Direct* measures the equality of bilateral exports between countries 0 and i. A measure of total trade, *Trade*, is the weighted sum of *Indirect* and *Direct*. The weight on the latter term is $(x_{i0} + x_{0i})/(x_{0.} + x_{i.})$. Table 47 lists the values and rankings of the countries. For the European and Asian currency crises, the trade linkages between the initially attacked countries, Italy and Thailand, and home countries are higher than those of other countries. However, for the Mexican currency crisis, the trade linkages between the first victim, Mexico, and home countries are not very high. Based on the lower levels of competition in the third market between the Latin American countries, *Indiect* is low while *Direct* is high. However, the weights for direct trade linkages are low since there is very little direct trade in their whole trade volumes. Finally to control for a market sentiment, *Pure*, a dummy variable equal to one when one of the regional countries is attacked by speculative agents, is included in the regression.

Additional determinants of currency crisis

While the first generation models of currency crisis proved that high government consumption levels (denoted *GOVC*) was a crucial factor for the onset of currency crisis before 1990s, additional factors such as a slowdown in real GDP growth rate (denote *GDP*), excess capital inflows (denoted *CAPI*) and increasing foreign liabilities (denoted by *FORLB*) were identified as important determinants of currency crises in the1990s by the second generation models. Therefore, it is necessary for us to analyze whether these variables help to explain the cross-country variation in the crisis indices after controlling for a lending boom, a real appreciation, and a reserves adequacy ratio. Each variable is measured as the average ratio to GDP (*GOVC*, *CAPI*, and *FORLB*) or the change rate of real GDP growth rate (*GDP*) over the four years to the onset of the crisis (1988-92, 1990-94, and 1992-96).

3.2 Data set

A three-period panel data is used from three different episodes of important and widespread currency crisis in 1990s. The three episodes are: 1) the European currency crisis of 1992-93; 2) the Mexican currency crisis of 1994-95; and 3) the Asian currency crisis of 1997-98. All the variables except bilateral trade are taken from the CD-ROM version of the International Monetary Fund's *International Financial Statistics* (IFS). For the bilateral trade, the IMF's *Direction of Trade* is used. The data set includes data from 29 countries[36]. The countries are grouped as European, Asian and Latin American countries or industrial and developing countries. The sample was chosen based on the existence of free convertibility and financial markets. The sample was also selected to ensure that both industrial and developing countries were included.

For the currency crisis index, monthly data of nominal exchange rate (line rf) and international reserves (line 11.d) were collected from January 1985 through January 1998. To estimate the impact of lending booms, it was necessary to collect annual data of claims on the private sector by deposit money banks and monetary authorities (line 32d) and nominal GDP (line 99b) for the years 1988 through 1996. For the real exchange rate,

[36] The countries are U.S.A., U.K., France, Italy, the Netherlands, Norway, Sweden, Canada, Japan, Finland, Spain, Australia, Germany, Argentina, Brazil, Chile, Colombia, Mexico, Jordan, Sri Lanka, India, Indonesia, Korea, Malaysia, Pakistan, the Philippines, Thailand, Turkey, and Venezuela.

the CPI annual data (line 64) over the 1988-96 period as well as nominal exchange rate were collected. To measure reserve adequacy, monthly data of money (line 34) and quasi-money (line 35) as well as international reserves for August 1992, November 1994 or June 1997 were collected. To capture contagion effect, annual data of money market interest rates (line 60b) or discount rates (line 60) were collected for the years 1988 through 1996. In addition, annual bilateral trade data is used to estimate trade linkage index over the same period as the interest rate. For additional determinants of the currency crises, annual data of government consumption (line 91f), capital accounts (line 78bc), financial accounts (line 78bj), net errors and omissions (line 78 ca), real GDP (line 99 br), and foreign liabilities (line 26c) were collected for the years 1988 through 1996.

3.3 Regression analysis

As discussed in the theoretical framework, a currency crisis occurs when the investors launch an attack because of the weak fundamentals of the country and its relatively low reserves level. The targeted countries for a speculative attack are those countries that are most likely to respond with an excessive depreciation. Following Sachs, Tornell and Velasco (1996), an empirical implementation of these ideas is made by classifying observations into four groups: high and low reserves cases, and strong and weak fundamentals cases. However, since the classification system includes both industrial and developing countries, the country-years with high reserves and strong fundamentals are different from theirs. A country is defined to have a high level of international reserves if the ratio of its *M2* to its reserves is in the lowest quartile for either industrial or developing countries. The dummy variable for high reserves, *dhr*, is equal to one for countries whose money-to-reserves ratio is in the bottom quartile for its group. By contrast, a country has strong fundamentals if its real depreciation is in the highest quartile of sample and its lending boom is in the lowest quartile of sample. The dummy variable for strong fundamentals, *dsf*, is equal to one for countries that have strong fundamentals.

3.3.1 Country effects

In the sample, there are three observations per country for the European, Mexican and Asian currency crisis. As such, we need to check the existence of country effects to determine the correct specification of benchmark regression model. Based on this, the following regression is estimated using pooled OLS, fixed effects and random effects models. The specification consists of

$$Crisis_{it} = \beta_0 + \beta_1 LB_{it} + \beta_2 RER_{it} + \beta_3 dhr \cdot LB_{it} + \beta_4 dhr \cdot RER_{it} + \beta_5 dsf \cdot LB_{it} + \beta_6 dsf \cdot RER_{it} + v_{it} \tag{12}$$

where i indexes the country and t indexes time; $v_{it} = a_i + u_{it}$ and a_i is an unobserved country effect.

To test the null hypothesis of no country effects against the alternative of fixed effects, the unobserved effect a_i is replaced by 28 terms of the form $a_i * d_i$ in equation (12), where d_i is a dummy that equals to one if the observation corresponds to country i. Then, the model is estimated using the pooled OLS model and an F test is performed. Under the null, all coefficients of β_0 and a_i's are equal. The F statistic is

$$F[28, 52] = \frac{(0.3607 - 0.1225)/28}{(1 - 0.3607)/52} = 0.6920.$$

Since the 5% critical value is 1.69, we failed to reject the null hypothesis of no fixed effects.

Next, to test the null hypothesis of no country effects against the alternative of random effects, a Breusch Pagan test is performed after the model in equation (12) is first estimated using a random effects model. The null hypothesis means that the variance of a_i is zero. The test statistic for a Breusch Pagan test is 1.29 and we failed to reject the null hypothesis at the 5% critical level.

The model in equation (12) is then estimated using the pooled OLS and fixed effects models.[37] As shown in Table 48, the point estimates of *LB* and *RER* have the same signs regardless of the specification of the models. However, while the estimated *LB* coefficient is significantly different from zero at 5% level in the pooled OLS model, it is not in the fixed effects model.

The results of the tests indicate that a pooled OLS model is an appropriate specification. Based on this, the pooled OLS model will be the benchmark regression in the remainder of this chapter.

3.3.2 Benchmark regression

In the benchmark regression, 87 observations for the 1992, 1994, and 1997 crises are stacked and the following regression using ordinary least squares is estimated.

$$Crisis_{it} = \beta_0 + \beta_1 LB_{it} + \beta_2 RER_{it} + \beta_3 dhr \cdot LB_{it} + \beta_4 dhr \cdot RER_{it} + \beta_5 dsf \cdot LB_{it}$$

$$+ \beta_6 dsf \cdot RER_{it} + u_{it} \qquad (13)$$

The effects of a lending boom and real appreciation with weak fundamentals and low reserves are reflected in β_1 and β_2, respectively. The signs of β_1 and β_2 are expected to be positive and negative, respectively. In addition, $\beta_1 + \beta_3$ and $\beta_2 + \beta_4$ indicate the effects of a lending boom and real appreciation with high reserves. The effects of a lending boom and real appreciation with strong fundamentals are likewise indicated by $\beta_1 + \beta_5$ and $\beta_2 + \beta_6$, respectively. This study's expectation is that $\beta_1 + \beta_3$ =0, $\beta_2 + \beta_4$ = 0, $\beta_1 + \beta_5 = 0$ and $\beta_2 + \beta_6 = 0$.

The currency crisis index used here is obtained with data from five months after the eruption of the crisis. For the European crisis, the time period is from September 1992 through January 1993. For the Mexican and Asian crises, it is from November 1994 through April 1995 and June 1997 through November 1997, respectively. The estimated regression is

$$Crisis_{it} = \underset{(1.85)}{3.90} + \underset{(0.04)}{0.08} LB_{it} - \underset{(0.13)}{0.18} RER_{it} - \underset{(0.11)}{0.03} dhr \cdot LB_{it} + \underset{(0.27)}{0.01} dhr \cdot RER_{it} - \underset{(0.21)}{0.24} dsf \cdot LB_{it}$$

$$- \underset{(0.24)}{0.21} dsf \cdot RER_{it}$$

$$R^2 = 0.12, \quad \bar{R}^2 = 0.06, \quad \text{N} = 87 \qquad (14)$$

[37] I do not present the results of estimation using the random effects model since they are as same as the pooled OLS estimation.

Heteroskedasticity robust standard errors are presented in parentheses. The point estimates in equation (14) indicate that the estimated coefficients of *LB* and *RER* have positive and negative signs as expected. A one unit increase in the *LB* or a one unit decrease in the *RER* for a country-year with low reserves and weak fundamentals leads to 0.08 or 0.18 unit increase in the crisis index, respectively. In addition, the estimated coefficients of *LB* and *RER* are significantly different from zero at the 5% and 10% level, respectively. This justifies the inclusion of both variables in the equation (14). The fourth column of Table 49 presents the results of estimation with the sample only including industrial countries. In this sample, the signs of *LB* and *RER* are negative and both variables are not significant at the10% level. In the strong and stable financial systems of industrial countries, a lending boom does not contribute to the variation of the crisis index. By contrast, the estimation results from the sample of developing countries in the fifth column shows that the signs of *LB* and *RER* are positive and negative and only *LB* is significant at the 5% level. Therefore, the existence of a lending boom has a stronger impact on the crisis index in developing countries than the industrial countries. Since the lending boom before the currency crisis is a common experience in the developing countries,[38] this result is not puzzling. As an additional check for this, the terms *LB*ddev* and *RER*ddev* were added to equation (13) where *ddev* takes the value of one for observations that correspond to the years, 1994 and 1997. The sixth column of Table 49 shows that the estimated coefficient of *LB*ddev* is significant at the 5% level whereas the *RER*ddev*'s coefficient is not significantly different from zero.

[38] See Chapter II for details.

Since the lending boom's effect on the currency crisis is different depending on the crisis episodes, we need to check whether the same model that explains the crises of industrial countries in 1992-93 also explains the cross-country variation in the 1994-95 and 1997-98 crises. To test the hypothesis that the coefficients in the equation (13) are the same in both periods, I perform a Chow test. The test statistic is

$$F[7, 73] = \frac{(17606 - 5946 - 6599)/7}{(5946 + 6599)/73} = 4.21$$

Since the critical value at the 5% level is 2.14, we can reject the null hypothesis that the coefficients are the same for the crises of industrial and developing countries.

Following a confirmation of the initial theoretical implications, *F*-tests indicate that the hypothesis $\beta_1 + \beta_3 = 0$ and $\beta_2 + \beta_4 = 0$ failed to be rejected in the third column of Table 49. Therefore, in countries with higher levels of reserves, neither *LB* nor *RER* affect the severity of a crisis. In addition, for the countries with strong fundamentals, $\beta_1 + \beta_5 = 0$ and $\beta_2 + \beta_6 = 0$ cannot be rejected. Hence, neither changes in *LB* or *RER* affect the severity of a crisis in the countries with strong fundamentals.

3.3.3 Contagion effects

To find contagion effects at the onset of the currency crisis, a regression with the trade linkage index variable, *Trade*, interest rate differential, *Itrdus*, and a pure contagion variable, *Pure* that reflects market sentiment. The estimated regression is

$$
\begin{aligned}
Crisis_{it} = & -6.58 + 0.08LB_{it} - 0.04RER_{it} + 0.08dhr \cdot dcg \cdot LB_{it} - 0.13dhr \cdot dcg \cdot RER_{it} \\
& \quad (3.04) \quad (0.03) \qquad (0.12) \qquad\quad (0.10) \qquad\qquad\qquad (0.22) \\
& - 0.28dsf \cdot dcg \cdot LB_{it} - 0.01dsf \cdot dcg \cdot RER_{it} + 26.65Trade_{it} + 6.48Pure_{it} \\
& \quad (0.29) \qquad\qquad\quad (0.29) \qquad\qquad\qquad (9.30) \qquad\quad (3.90) \\
& - 0.40Itrdus_{it} \\
& \quad (0.47)
\end{aligned}
$$

$$
R^2 = 0.35, \quad \overline{R}^2 = 0.28, \quad N = 87 \tag{15}
$$

A country is defined as not exposed to contagion effects if its trade linkage index is in the lowest quartile of the sample or its pure contagion index is zero. In addition, if the interest rate differential is in the highest quartile of the sample, a country is not under the influence of the contagion effects. Thus, the dummy variable for the contagion effect, *dcg*, is equal to one for countries not exposed to the contagion effects. Therefore, the point estimates of the coefficients of *LB* and *RER* in equation (15) reflect the effect of a unit increase of *LB* and *RER* on the crisis index under all conditions for countries who do not have high reserves, strong fundamental, and have been exposed to contagion effects. In equation (15), the estimated coefficients of *Trade* and *Pure* have positive signs and *Itrdus* coefficient has a negative sign as expected. The scale of *Trade* and *Pure* goes from zero to one. In this framework, a 0.1 unit increase in *Trade* and *Pure* leads to 2.67 and 0.65 unit increases in the crisis index, respectively. A one unit decrease in *Itrdus* also increases the crisis index by 0.40 units. However, while the estimated coefficients for *Trade* and *Pure* are significantly different from zero at the 5% level, *Itdrus* coefficient is not significantly different from zero at the 5% level. This indicates that the contagion effects from the trade linkage and market sentiment played a crucial role in the currency crises after 1990. The third column of the Table 50 indicates that the estimated coefficients of *LB* and *RER* are still significant at the 5% and 10% level even with the dummy *dcg*. The fifth column of the Table 50 indicates that the estimated coefficients of *RER*, *Trade*, *Pure* and *Itrdus* has the expected signs but all the coefficients except *Itrdus* and *Constant* are no longer significantly different from zero. Thus, with the exception of *Itrdus*, the relationships between the crisis index and explanatory variables in the currency crisis of industrial countries are not reliable. By contrast, the sixth column, where the estimated coefficients of *Trade* and *Pure* are significant at the 5% level, shows that the contagion effects played a key role in the onset of the Mexican and Asian currency crises.

3.3.3 Additional determinants of currency crisis

In this subsection, it is analyzed whether higher government consumption, a slowdown in real GDP growth rate, excess capital inflows, and increasing foreign liabilities help to explain the cross-country variation in the crisis indices after controlling for a lending boom, real appreciation, reserves adequacy and contagion effects. Table 51 presents the estimated coefficient for each variable.

The third column of Table 51 presents the estimated coefficients of government consumption, denoted by *GOVC*. The regression results indicate that government consumption does not significantly effect the crisis index. This coincides with the literature[39]. The literatures has found that government consumption had a weaker effect on the currency crises in 1990s relative to the crises in 1980s. The insignificant coefficient on *GOVC* may be explained by the changing nature of crises.

Similarly, a slowdown of real GDP growth rate is assumed to increase the policymaker's incentive to switch to a more expansionary policy, which can be achieved through a nominal devaluation of the currency. Therefore, the real GDP growth rate, denoted by *GDP*, should capture the escape-clause interpretation developed in various second-generation models of currency crisis.[40] The fourth column of Table 51 shows that the estimated coefficient of *GDP* has an expected negative sign but it is not significantly different from zero at the 5% level. Hence, a decline in the real GDP growth rate does not appear to contribute to the currency crises. Table 2, in Chapter II, shows how Asian countries continued to have relatively high GDP growth rate in the 1990s before the 1997 crisis although the growth rate slowed slightly prior to the crisis

For the capital inflows, denoted by *CAPI*, the estimated coefficients are reported in the fifth column of Table 51. Excessive capital inflows are regarded as a main factor for the onset of currency crisis. This is because the short time span of excessive inflows prevents them being efficiently channeled to productive projects and eventually lead to a shortage of returns to repay investors. The fifth column of Table 51 shows that the estimated coefficient of *CAPI* has positive sign as we expect. It is also significantly different from zero at the 5% level. Hence, even after controlling for all the other

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[39] Sachs, Tornell and Velasco (1996), Pazarbasioglu and Otker (1996), and Corsetti, Giancalro, Pesenti, and Roubini (1998)

[40] For example, Obstfeld (1994, 1996).

contributors to a crisis, a one unit increase in the capital inflow index leads to 0.56 unit increase in the crisis index.

The sixth column of Table 51 presents the estimated coefficients of foreign liabilities, denoted by *FORLB*. We expect the bank's foreign liabilities as a ratio of GDP to represent the extent to which the banking system is exposed to international capital flow. However, the point estimate sign goes against expectations but also is not significant at the 5% level. Hence, the foreign liabilities do not contribute significantly to the cross-country variation of the crisis index.

3.4 Robustness

To analyze whether the results are robust over the periods in which the crisis index is measured, equation (13) is estimated again using six different crises indices. For all indices, the starting point is the month preceding the onset of the crisis (i.e. August 1992 for the European crisis, November 1994 for the Mexican crisis, and June 1997 for the Asian currency crisis). Then, the terminal month is varied over a period of six months starting with October 1992, January 1995, and August 1997. As Table 52 shows, in columns three, four, six and eight, the point estimates of *LB* are similar to the benchmark regression point estimates (the fourth column). Moreover, they are significantly different from zero at both the 5% or 10% level with the exception of the coefficient reported in the eighth column. The point estimates for *RER* show some variation across specifications but is always significantly different from zero at the 5 or 10% level. Other variables show the same tendency in their values and significances.

In the benchmark regression, a country-year is classified as having high reserves if its ratio of M2 to reserves is in the lowest quartile for industrial or developing countries at the onset of the crisis. The threshold values are 8.0 (10 country-years) and 2.8 (12 country-years), respectively. A country-year is also classified as having strong fundamentals if lending boom index is in the lowest quartile of sample and real appreciation index is in the highest quartile of the sample where the threshold value for the lending boom is 8.0 and for the real appreciation is –9.0 (9 country-years). The fourth and fifth columns of Table 53 show the estimates for different thresholds concerning the high reserves dummy, while keeping the strong fundamentals dummy unchanged. The thresholds are 6.7 (8 country-years) and 2.0 (8 country-years) for the fourth column and 9.5 (11 country-years) and 3.2 (14 country-years) for the fifth column. Column 6 and 7 also indicate the estimates based on different thresholds for the strong fundamentals dummy, while keeping the high reserves dummy stayed at the same level. The thresholds of *RER* and *LB* are 12.6 and -10.9 (6 country-years) for the sixth column and 3.8 and –4.0 (15 country-years) for the seventh. The last column shows the estimates after all the thresholds for the dummies are changed. The thresholds of the ratio of M2 to reserves, *RER* and *LB* are 6.7 (8 country-years) and 2.0 (8 country-years), 12.6 and –10.9 (6 country-years) respectively. As seen in the columns of Table 53, the coefficient's values and significance levels do not substantially differ from each other based on the threshold values selected.

3.5 Predicting the Asian currency crisis

3.5.1 Prediction without the contagion effects

Based on the benchmark regression's estimates, we may predict the currency crisis indices in the Asian crisis, 1997. To this end, the following regression can be estimated with data from the 1994 crisis:

$$Crisis_{it} = \beta_0 + \beta_1 LB_{it} + \beta_2 RER_{it} + \beta_3 CAPI_{it} + \beta_4 dhr \cdot LB_{it} + \beta_5 dhr \cdot RER_{it} + \beta_6 dsf \cdot LB_{it} + \beta_7 dsf \cdot RER_{it} + u_{it} \qquad (16)$$

This equation includes the capital inflow index denoted by *CAPI.* Its inclusion is based on its ability to explain the cross-country variation of the currency index. Then, an out-of-sample predicted crisis index is constructed by substituting into equation (16) the estimated coefficients of a regression that are significantly different from zero and the values of the explanatory variables that correspond to the Asian currency crisis. The fourth column of Table 54 presents the resulting predicted crises indices without the contagion effects according to descending order.

First, the predicted crisis indices is a dotted line in Figure 63. As shown in Figure 63, the predicted crisis indices follows closely the actual crisis indices well. Second, each 5 countries is then grouped in descending order of predicted crisis indices. This chapter then checks how many rankings of countries in each group of predicted crisis indices coincide with the rankings of the countries in the matched group ranked by the actual indices as reported in column 2 of Table 54. A total of 8 countries have rankings that match in column 2 and 4. Third, the actual crisis indices of 1997 are regressed on the predicted out-of-sample crisis indices. The regression result is

$$Actual\ 97crisis_i = 6.55 + 0.42 \cdot [\Pr edicted\ 97crisis_i]$$
$$(2.17)\ (0.12)$$

$$R^2 = 0.26\ , \quad \overline{R}^2 = 0.24, \quad \mathrm{N} = 29 \qquad (17)$$

The regression coefficient is 0.42, and significantly different from zero at the 5% level.

Fourth, a Root Mean Square Errors, denoted by $RMSE_f$ is estimated;

$$RMSE_f = \sqrt{\frac{1}{N}\sum_{i=1}^{N}(Actual\ 97crisis_i - \Pr edicted\ 97crisis_i)^2}\ , \mathrm{N} = 29$$

The estimated $RMSE_f$ is 16.02. This value will be compared with $RMSE_c$, the Root Mean Square Errors when the predicted crisis indices includes the impact of contagion effects as shown in the following subsection.

3.5.2 Prediction with the contagion effects

The previous section's predicted currency indices for the Asian currency crisis was obtained with the benchmark regression. The benchmark regression proved to be dependable in predicting the out of sample movements of the Asian currency crisis indices. However, as shown in Table 54, the predicted crisis indices of the countries, Indonesia and Korea, which were severely attacked during the Asian crisis, were not predicted well. To improve our ability to predict the crisis indices, the following regression is estimated

$$Crisis_{it} = \beta_0 + \beta_1 LB_{it} + \beta_2 RER_{it} + \beta_3 CAPI_{it} + \beta_4 dhr \cdot dcg \cdot LB_{it} + \beta_5 dhr \cdot dcg \cdot RER_{it}$$
$$+ \beta_6 dsf \cdot dcg \cdot LB_{it} + \beta_7 dsf \cdot dcg \cdot RER_{it} + \beta_8 Trade_{it} + \beta_9 Pure_{it} + \beta_{10} Itrdus_{it} + u_{it} \qquad (18)$$

where the contagion effects are captured by *dcg*, *Trade*, *Pure* and *Itrdus*. After the Philippines and Thailand were predicted as the countries that are most likely to be attacked by the speculators by the benchmark regression without the contagion effects, the trade linkage index, *Trade*, and market sentiment variable, *Pure*, for the Asian crisis are reconstructed based upon the prediction of the benchmark regression.[41] Then, an out-of-sample predicted crisis index with contagion effects is constructed with the same procedure used in the previous subsection.

The sixth column of Table 54 presents the resulting predicted crises indices with the contagion effects according to descending order. As shown in Figure 63, the new predicted crisis indices closely follows the actual crisis indices. In addition, the new predicted crisis indices appears to fit the actual crisis indices better than the prediction made from the benchmark regression in Figure 63. Furthermore, a total of 11 as opposed to 8 countries have the same ranking based on the new predicted series as when they are ranked with the actual crisis indices.

Next, the actual crisis indices of 1997 are regressed on the new predicted out-of-sample crisis indices. The regression result is

$$\underset{}{Actual\ 97crisis_i} = \underset{(1.41)}{4.23} + \underset{(0.13)}{0.51} \cdot [Predicted\ 97crisis_i^c]$$

$$R^2 = 0.56\ , \quad \bar{R}^2 = 0.53, \qquad \mathrm{N} = 29 \tag{19}$$

The point estimate of the predicted crisis indices with contagion effects is now 0.51 instead of 0.42. It is also significantly different from zero at the 5% level. A larger

[41] The new trade linkage is a weighted sum of the trade linkages with the Philippines and Thailand which

portion of the variation in the actual crisis indices is explained by the new predicted crisis indices. This is shown by how the adjusted-R^2 increases from 0.24 to 0.53 and the estimated *RMSE* goes from 16.02 to 12.08. This indicates that the prediction based upon contagion effects improves our ability to predict an upcoming currency crisis.

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are weighted by the predicted crisis indices of the Philippines and Thailand, respectively.

Table 42. Currency crisis index*

Country	1993	Country	1995	Country	1997
Brazil	69.7	Mexico	70.9	Thailand	47.9
Finland	29.7	Brazil	19.1	Indonesia	33.5
Spain	25.2	Argentina	16.7	Malaysia	32.0
Sweden	22.8	Philippines	7.2	Philippines	29.7
U.K.	21.5	Italy	5.8	Korea	16.4
Italy	18.7	Spain	5.7	Colombia	16.3
Norway	18.1	Venezuela	4.3	Turkey	12.4
Venezuela	14.7	Colombia	1.8	Pakistan	8.4
Turkey	11.4	Indonesia	1.1	Brazil	7.3
Australia	11.1	Pakistan	0.7	Australia	4.6
U.S.A.	9.7	Sri Lanka	0.1	Japan	4.4
France	9.4	Canada	-0.1	India	3.8
Sri Lanka	8.9	Malaysia	-0.6	Finland	3.8
Canada	8.8	India	-0.9	Canada	2.9
Jordan	3.9	Australia	-1.2	Netherlands	1.9
Colombia	3.2	Jordan	-1.4	Germany	1.3
Pakistan	3.1	Thailand	-2.4	U.S.A	1.2
India	2.9	U.K.	-2.7	U.K.	0.8
Malaysia	2.1	Sweden	-3.3	Sri Lanka	0.3
Indonesia	1.9	Korea	-4.5	Sweden	0.3
Germany	1.8	France	-7.3	Mexico	0.1
Thailand	0.0	Germany	-7.4	Chile	-0.5
Chile	-0.3	Chile	-7.5	Argentina	-0.7
Japan	-0.5	Finland	-7.8	France	-1.5
Korea	-0.7	Norway	-9.5	Jordan	-2.3
Philippines	-1.6	Netherlands	-9.5	Spain	-2.9
Netherlands	-2.3	Turkey	-12.4	Norway	-4.5
Mexico	-2.4	Japan	-18.7	Italy	-7.7
Argentina	-12.3	U.S.A	-23.4	Venezuela	-11.0

* The currency crisis index (*Crisis*) is a weighted average of the percentage depreciation of nominal exchange rate with respect to the U.S. dollar and the percentage decrease in reserves.

Table 43. Lending boom*

Country	1988-1992	Country	1990-1994	Country	1992-1996
Mexico	231.1	Mexico	124.6	Philippine	137.2
Turkey	181.0	Philippines	51.0	Colombia	40.3
Indonesia	67.7	Thailand	40.9	Thailand	38.4
Australia	44.4	Brazil	30.8	Turkey	38.3
Thailand	41.4	Colombia	25.5	Malaysia	25.6
Philippines	28.1	Sri Lanka	24.7	Chile	23.3
Malaysia	21.6	Canada	20.7	Netherlands	22.8
Canada	20.0	Argentina	17.1	Jordan	22.3
Korea	17.7	Netherlands	10.7	Indonesia	21.8
Finland	17.0	Indonesia	10.6	Canada	20.2
U.K.	14.9	Chile	8.0	Argentina	16.7
France	10.0	Italy	7.1	Germany	15.7
Italy	5.8	Korea	6.4	Korea	14.3
Netherlands	4.5	Australia	6.1	Sri Lanka	12.1
Germany	4.3	Malaysia	4.9	Australia	7.6
Japan	4.3	Germany	3.9	U.K.	3.6
Sweden	4.2	Jordan	1.9	U.S.A	3.1
Sri Lanka	1.3	Spain	-0.9	Norway	0.8
Spain	1.1	Pakistan	-3.1	Pakistan	0.5
Colombia	-1.6	Japan	-4.2	Spain	-3.9
India	-2.9	U.K.	-4.8	Japan	-0.7
Norway	-10.2	India	-5.9	India	-4.0
Jordan	-10.8	Turkey	-8.6	Italy	-9.5
Pakistan	-11.9	France	-10.6	France	-15.5
Chile	-20.0	U.S.A	-11.3	Finland	-29.2
Brazil	-13.1	Finland	-13.3	Sweden	-32.1
U.S.A.	-16.0	Norway	-13.5	Mexico	-39.6
Argentina	-23.1	Sweden	-30.1	Brazil	-51.6
Venezuela	-37.9	Venezuela	-44.5	Venezuela	-56.7

* Lending boom (*LB*) is the percentage change in the ratio of claims on the private sector by deposit money banks and monetary authorities (line 32d) to GDP (line 99b).

Table 44. Real depreciation*

Country	1988-1992	Country	1990-1994	Country	1992-1996
Argentina	-50.0	Argentina	-41.1	Colombia	-33.5
Mexico	-27.6	Brazil	-27.4	Brazil	-24.1
Turkey	-25.2	Japan	-27.0	Philippines	-19.6
Spain	-21.2	Colombia	-25.0	Chile	-16.1
Chile	-18.2	Philippines	-19.7	Thailand	-11.8
Sweden	-17.3	Mexico	-19.3	Indonesia	-11.7
Brazil	-17.3	Chile	-12.4	Sri Lanka	-10.8
Philippines	-16.6	Venezuela	-9.5	Japan	-10.1
Italy	-14.2	Sri Lanka	-8.8	Malaysia	-9.3
Germany	-11.2	Malaysia	-8.5	Australia	-9.2
U.K.	-10.2	Thailand	-7.9	Korea	-9.2
France	-9.7	Indonesia	-5.8	Argentina	-8.5
Korea	-8.5	Germany	-2.2	Venezuela	-8.1
Netherlands	-7.4	Jordan	-1.9	Germany	-5.3
Sri Lanka	-7.3	Korea	-1.6	Netherlands	-5.0
Thailand	-6.6	U.S.A.	-0.9	Jordan	-3.9
Norway	-5.3	Netherlands	0.6	U.S.A	-2.5
Canada	-4.9	Pakistan	4.3	France	-2.4
Malaysia	-3.6	France	5.3	Pakistan	0.8
U.S.A	-2.6	Australia	11.0	Norway	3.8
Finland	1.4	U.K.	14.6	India	4.2
Australia	1.4	Norway	15.3	Finland	5.3
Japan	1.6	Canada	18.9	Turkey	9.6
Indonesia	2.4	Spain	20.5	U.K.	10.7
Colombia	7.8	Sweden	22.2	Sweden	12.7
Pakistan	10.7	Italy	25.2	Spain	12.8
Venezuela	17.9	Turkey	32.0	Italy	14.3
Jordan	22.3	India	34.5	Canada	14.4
India	44.5	Finland	39.2	Mexico	22.8

* Real depreciation of the exchange rate (*RER*) is the percentage change in the real exchange rate index over the four years prior to the onset of the crisis

Table 45. Reserve adequacy (Industrial countries)*

Country	Aug. 1993	Country	Nov. 1994	Country	June 1997
U.S.A	57.9	U.S.A.	63.6	U.S.A	81.5
Japan	54.0	Japan	42.1	U.K.	40.6
Italy	33.7	France	34.7	France	32.4
France	26.3	U.K.	24.9	Japan	22.3
U.K.	24.4	Canada	24.6	Canada	17.7
Canada	19.9	Italy	23.0	Germany	16.6
German	17.9	Australia	18.4	Australia	16.5
Netherlands	15.9	German	16.3	Italy	14.9
Finland	13.4	Spain	9.7	Netherlands	11.3
Australia	13.2	Netherlands	8.6	Sweden	7.1
Sweden	6.9	Finland	5.9	Spain	6.3
Spain	6.6	Sweden	4.2	Finland	5.9
Norway	4.7	Norway	3.7	Norway	2.8

* Reserve adequacy is the ratio of M2 to reserves in the month preceding the onset of the crisis (August 1992, November 1994 or June 1997).

Table 46. Reserve adequacy (Developing countries)*

Country	Aug. 1993	Country	Nov. 1994	Country	June 1997
India	19.8	Mexico	9.0	Pakistan	20.9
Pakistan	18.9	Pakistan	8.7	India	7.5
Korea	6.9	India	7.5	Korea	6.8
Jordan	6.3	Korea	6.7	Indonesia	6.2
Turkey	6.3	Indonesia	6.1	Thailand	4.9
Philippines	5.7	Philippines	4.8	Philippines	4.9
Indonesia	4.8	Turkey	4.4	Jordan	4.3
Mexico	4.4	Argentina	4.1	Mexico	4.1
Thailand	4.0	Brazil	4.0	Malaysia	4.0
Brazil	3.7	Jordan	3.9	Brazil	3.7
Sri Lanka	3.3	Thailand	3.8	Argentina	3.6
Argentina	3.2	Norway	3.7	Turkey	3.2
Malaysia	2.8	Malaysia	2.1	Norway	2.8
Chile	1.7	Colombia	1.9	Sri Lanka	2.6
Venezuela	1.6	Sri Lanka	1.9	Colombia	2.0
Colombia	1.0	Venezuela	1.8	Chile	1.8

* Reserve adequacy is the ratio of M2 to reserves in the month preceding the onset of the crisis (August 1992, November 1994 or June 1997).

Table 47. Trade linkage*

Country	1992	Country	1994	Country	1996
Italy	1.00	Mexico	1.00	Thailand	1.00
France	0.86	Canada	0.54	Malaysia	0.83
Netherlands	0.80	Korea	0.45	Indonesia	0.81
U.K.	0.79	Japan	0.36	Australia	0.70
Germany	0.62	Malaysia	0.34	India	0.68
Spain	0.56	U.K.	0.34	Brazil	0.66
Japan	0.44	Venezuela	0.32	Philippines	0.65
Sweden	0.43	Brazil	0.31	Korea	0.63
U.S.A	0.38	Thailand	0.30	Finland	0.47
Korea	0.38	U.S.A	0.26	Sweden	0.46
Norway	0.35	Germany	0.25	Chile	0.44
Brazil	0.32	Italy	0.23	Norway	0.37
Malaysia	0.29	Philippines	0.23	Turkey	0.37
Thailand	0.27	Indonesia	0.22	Venezuela	0.37
Canada	0.25	India	0.22	Colombia	0.33
Finland	0.24	France	0.21	Argentina	0.31
Australia	0.23	Sweden	0.20	U.K.	0.30
Mexico	0.23	Colombia	0.18	Spain	0.30
Indonesia	0.22	Chile	0.18	Italy	0.29
India	0.27	Australia	0.17	Pakistan	0.29
Venezuela	0.16	Spain	0.16	France	0.27
Turkey	0.14	Argentina	0.15	Canada	0.26
Philippines	0.13	Norway	0.15	Netherlands	0.24
Chile	0.12	Finland	0.14	Mexico	0.24
Argentina	0.12	Netherlands	0.14	Japan	0.22
Colombia	0.09	Pakistan	0.11	Germany	0.20
Pakistan	0.07	Turkey	0.11	Sri Lanka	0.17
Sri Lanka	0.03	Sri Lanka	0.07	U.S.A	0.12
Jordan	0.01	Jordan	0.00	Jordan	0.03

* *Trade* is a weighted average index between the third market competition index, *Indirect*, and direct trade linkage index, *Direct*.

Table 48. Country effects

Dependent variable: *Crisis*			
Estimated coefficients	Independent variable	Pooled OLS	Fixed effects
β_1	*LB*	0.084**	0.054
		(0.041)	(0.110)
β_2	*RER*	-0.183*	-0.211
		(0.134)	(0.196)
β_3	*LB*dhr*	-0.035	0.161
		(0.106)	(0.320)
β_4	*RER*dhr*	-0.012	-0.056
		(0.268)	(0.484)
β_5	*LB*dsf*	-0.241	-0.230
		(0.209)	(0.579)
β_6	*RER*dsf*	-0.211	-0.393
		(0.241)	(0.476)
β_0	*Constant*	3.901**	4.854**
		(1.853)	(2.640)
Sample size		87	87
R^2		0.123	0.361
$\overline{R}^2$		0.060	0.000

Note: Heteroscedasticity robust standard errors in parenthesis.
Significance at the 10 percent level is denoted by *; at the 5 percent level by **

Table 49. Benchmark regression

Dependant variable: *Crisis*					
Estimated coefficients	Independent variable	Benchmark	Industrial countries crisis (1992)	Developing countries crisis (1994, 1997)	Changes of *LB* and *RER*
β_1	*LB*	0.084**	-0.043	0.330**	-0.014
		(0.041)	(0.063)	(0.112)	(0.052)
β_2	*RER*	-0.183*	-0.009	-0.180*	-0.179
		(0.134)	(0.293)	(0.123)	(0.253)
β_3	*LB*dhr*	-0.035	0.425	-0.236**	-0.224**
		(0.106)	(0.677)	(0.114)	(0.126)
β_4	*RER*dhr*	-0.012	-0.388	0.030	0.089
		(0.268)	(0.469)	(0.192)	(0.276)
β_5	*LB*dsf*	-0.241	-1.073	-0.343**	-0.225**
		(0.209)	(0.839)	(0.183)	(0.121)
β_6	*RER*dsf*	-0.211	-0.945	-0.192	-0.123
		(0.241)	(0.554)	(0.153)	(0.198)
β_7	*LB*ddev*+				0.316**
					(0.112)
β_8	*RER*ddev*				0.036
					(0.220)
β_0	*Constant*	3.901**	10.156**	-0.026	2.904*
		(1.853)	(3.843)	(1.737)	(1.993)
Sample size		87	87	87	87
R^2		0.123	0.068	0.492	0.290
$\bar{R}^2$		0.060	0.000	0.397	0.217
$\beta_1 + \beta_3 = 0$		0.31 [0.58]	0.33 [0.57]	7.54 [0.01]	
$\beta_2 + \beta_4 = 0$		0.55 [0.46]	1.79 [0.19]	0.01 [0.93]	
$\beta_1 + \beta_5 = 0$		0.00 [0.91]	1.76 [0.20]	1.64 [0.21]	
$\beta_2 + \beta_6 = 0$		1.75 [0.20]	3.69 [0.06]	2.78 [0.11]	

Note: Heteroscedasticity robust standard errors in parenthesis.
Significance at the 10 percent level is denoted by *; at the 5 percent level by **
P-values in brackets.

+ Dummy variable which is one when year is 1994 or 97

Table 50. Contagion effects

Dependant Variable: *Crisis*					
Estimated coefficients	Independent variable	Benchmark with *dcg*	Contagion effects	Industrial countries crisis (1992)	Developing countries crisis (1994, 1997)
β_1	*LB*	0.083**	0.076**	-0.051	0.217**
		(0.037)	(0.028)	(0.028)	(0.080)
β_2	*RER*	-0.173*	-0.037	0.075	-0.004
		(0.124)	(0.117)	(0.316)	(0.108)
β_3	*LB*dhr*dcg*	0.010	0.076	0.619	0.027
		(0.122)	(0.101)	(0.583)	(0.098)
β_4	*RER*dhr*dcg*	-0.066	-0.127	-0.227	-0.019
		(0.265)	(0.218)	(0.379)	(0.178)
β_5	*LB*dsf*dcg*	-0.267	-0.279	-1.169*	-0.273**
		(0.235)	(0.293)	(0.705)	(0.099)
β_6	*RER*dsf*dcg*	-0.184	-0.006	-0.736*	-0.048
		(0.256)	(0.286)	(0.458)	(0.176)
β_7	*Trade*		26.650**	7.331	32.952**
			(9.304)	(9.677)	(7.650)
β_8	*Pure*		6.478*	7.441	7.076**
			(3.892)	(6.522)	(3.891)
β_9	*Itrdus*		-0.399	-2.253**	-0.293
			(0.470)	(1.175)	(0.394)
β_0	*Constant*	3.904**	-6.582**	8.617**	-11.636**
		(1.706)	(3.040)	(6.613)	(2.453)
Sample size		87	87	87	87
R^2		0.123	0.352	0.184	0.724
$\bar{R}^2$		0.057	0.276	0.000	0.672

Note: Heteroscedasticity robust standard errors in parenthesis.
Significance at the 10 percent level is denoted by *; at the 5 percent level by **

Table 51. Additional determinants

		Dependent variable: *Crisis*			
Estimated co-efficients	Independent variable	*GOVC* (Government consumption)	*GDP* (Real GDP)	*CAPI* (Capital inflow)	*FORLB* (Foreign liability)
β_1	*LB*	0.080** (0.027)	0.075** (0.028)	0.062** (0.022)	0.072** (0.029)
β_2	*RER*	-0.043 (0.117)	-0.040 (0.119)	-0.051 (0.112)	-0.023 (0.110)
β_3	*LB*dhr *dcg*	0.073 (0.093)	0.085 (0.107)	0.058 (0.101)	0.080 (0.111)
β_4	*RER*dhr *dcg*	-0.134 (0.210)	-0.107 (0.235)	-0.029 (0.197)	-0.114 (0.225)
β_5	*LB*dsf *dcg*	-0.293 (0.296)	-0.289 (0.320)	-0.329 (0.295)	-0.256 (0.296)
β_6	*RER*dsf *dcg*	-0.038 (0.285)	0.033 (0.323)	-0.162 (0.253)	0.010 (0.289)
β_7	*Trade*	26.434** (9.572)	26.232** (9.545)	26.745** (8.979)	27.527** (9.452)
β_8	*Pure*	6.481* (3.871)	6.236* (3.917)	6.649* (3.971)	6.656** (3.861)
β_9	*Itrdus*	-0.321 (0.514)	-0.462 (0.502)	-0.545 (0.494)	-0.464 (0.500)
β_{10}	*Added Variable*	0.157 (0.210)	-0.176 (0.327)	0.563** (0.319)	-0.059 (0.047)
β_0	*Constant*	-8.938** (3.452)	-6.246** (3.383)	-7.916** (3.488)	-5.769** (3.102)
Sample #		87	87	87	87
R^2		0.354	0.354	0.370	0.359
$\bar{R}^2$		0.269	0.269	0.287	0.275

Note: Heteroscedasticity robust standard errors in parenthesis.
Significance at the 10 percent level is denoted by *; at the 5 percent level by **

Table 52. Robustness for the crisis index

		Dependant Variable: *Crisis*					
Estimated co-efficients	Indepen-dent variable	Aug-Oct*** Nov-Jan June-Aug	Aug-Nov Nov-Feb June-Sep	Aug-Dec Nov-Mar June-Oct	Aug-Jan Nov-Apr June-Nov	Aug-Feb Nov-May June-Dec	Aug-Mar Nov-June June-Jan
β_1	*LB*	0.071**	0.074**	0.101**	0.084**	0.069*	0.086
		(0.018)	(0.016)	(0.034)	(0.041)	(0.054)	(0.081)
β_2	*RER*	-0.051*	-0.122**	-0.179**	-0.183*	-0.256*	-0.350*
		(0.033)	(0.049)	(0.093)	(0.134)	(0.171)	(0.268)
β_3	*LB*dhr*	-0.039	-0.068	-0.088	-0.035	-0.038	-0.068
		(0.050)	(0.081)	(0.094)	(0.106)	(0.123)	(0.144)
β_4	*RER*dhr*	0.086	-0.043	-0.054	-0.012	0.106	0.228
		(0.116)	(0.173)	(0.220)	(0.268)	(0.329)	(0.434)
β_5	*LB*dsf*	-0.117	-0.048	-0.082	-0.241	-0.204	-0.201
		(0.095)	(0.153)	(0.173)	(0.209)	(0.217)	(0.201)
β_6	*RER*dsf*	-0.101	0.013	-0.060	-0.211	-0.286	-0.401
		(0.112)	(0.168)	(0.209)	(0.241)	(0.277)	(0.324)
β_0	*Constant*	1.595**	2.489**	3.023**	3.901**	6.635**	10.114**
		(0.559)	(0.812)	(1.457)	(1.853)	(2.569)	(3.953)
Sample #		87	87	87	87	87	87
R^2		0.145	0.167	0.165	0.123	0.081	0.055
$\bar{R}^2$		0.080	0.105	0.102	0.060	0.012	0.000

Note: Heteroscedasticity robust standard errors in parenthesis.
Significance at the 10 percent level is denoted by *; at the 5 percent level by **
*** 1992, 1994 and 1997 crisis, respectively.

Table 53. Robustness for the dummies

		Dependant Variable: *Crisis*					
		Benchmark	M2/R (20%)	M2/R (30%)	RER & LB (20%)	RER & LB (30%)	M2/R, RER & LB (20%)
Estimated coefficients	Independent variable	M2/RI < 8.0 M2/RD<2.8 RER > 8.0 LB < -9.0	M2/RI < 6.7 M2/RD<2.0 RER > 8.0 LB < -9.0	M2/RI < 9.5 M2/RD<3.2 RER > 8.0 LB < -9.0	M2/RI< 8.0 M2/RD<2.8 RER > 12.6 LB < -10.9	M2/RI < 8.0 M2/RD<2.8 RER > 3.8 LB < -4.0	M2/RI< 6.7 M2/RD<2.0 RER > 12.6 LB < -10.9
β_1	*LB*	0.084**	0.083**	0.087**	0.084**	0.085**	0.082**
		(0.041)	(0.040)	(0.041)	(0.041)	(0.040)	(0.040)
β_2	*RER*	-0.183*	-0.191*	-0.181*	-0.185*	-0.156	-0.195*
		(0.134)	(0.131)	(0.134)	(0.136)	(0.140)	(0.133)
β_3	*LB*dhr*	-0.035	-0.027	-0.056	-0.032	-0.033	-0.024
		(0.106)	(0.094)	(0.112)	(0.106)	(0.102)	(0.093)
β_4	*RER*dhr*	-0.012	0.047	-0.027	0.009	-0.019	0.079
		(0.268)	(0.259)	(0.269)	(0.290)	(0.233)	(0.282)
β_5	*LB*dsf*	-0.241	-0.261*	-0.225	-0.271	-0.173	-0.305*
		(0.209)	(0.185)	(0.215)	(0.223)	(0.171)	(0.197)
β_6	*RER*dsf*	-0.211	-0.258	-0.199	-0.279	-0.194*	-0.355
		(0.241)	(0.236)	(0.243)	(0.280)	(0.146)	(0.272)
β_0	*Constant*	3.901**	3.984**	3.865**	3.935**	4.182**	4.015**
		(1.853)	(1.785)	(1.831)	(1.800)	(1.827)	(1.728)
Sample #		87	87	87	87	87	87
R^2		0.123	0.123	0.124	0.123	0.123	0.124
$\bar{R}^2$		0.060	0.060	0.060	0.060	0.060	0.060

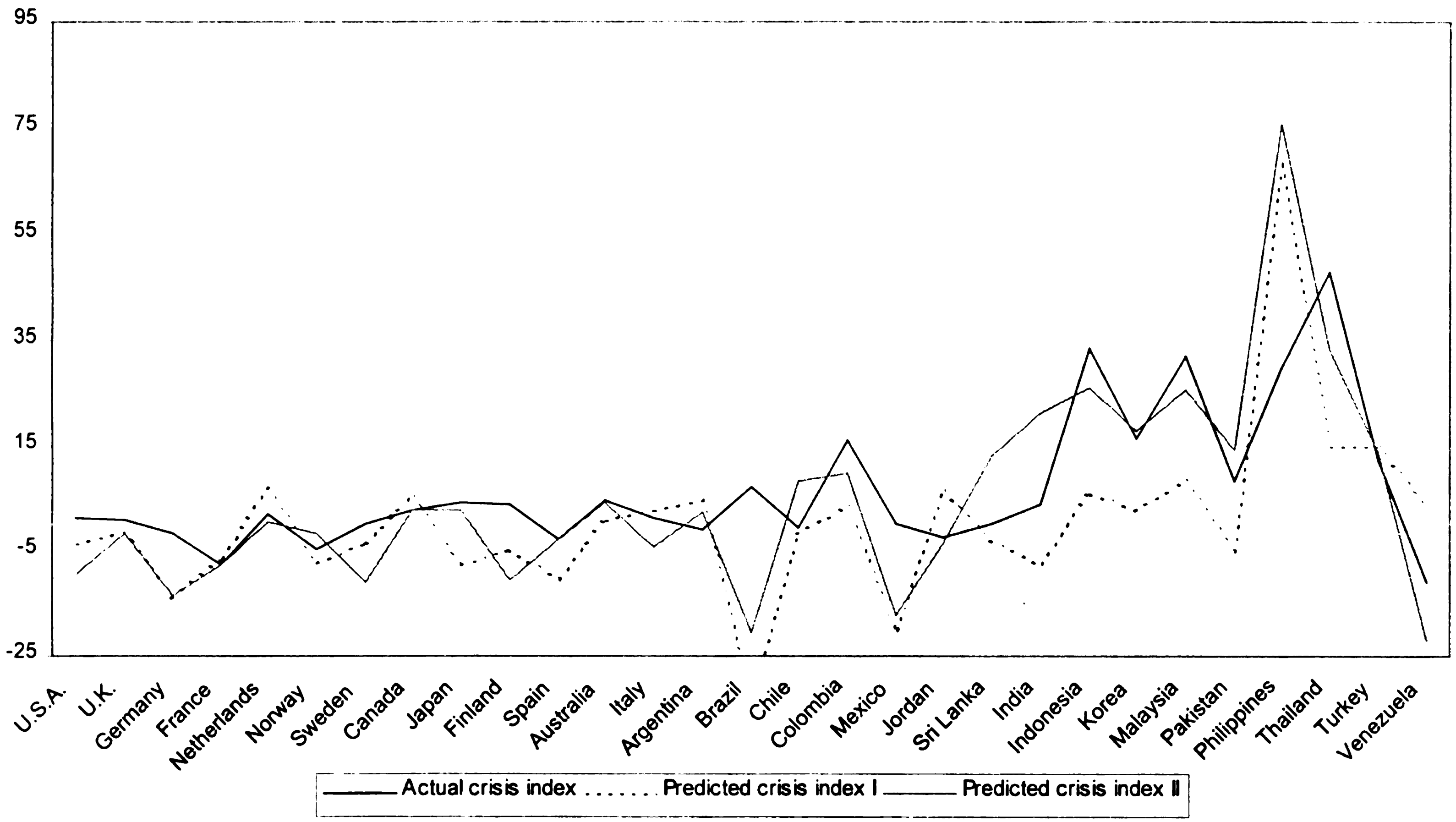
Note: Heteroscedasticity robust standard errors in parenthesis.
Significance at the 10 percent level is denoted by *; at the 5 percent level by **

Table 54. Actual and predicted currency crisis index*

Actual crisis index		Predicted crisis index without contagion effects			Predicted crisis index with contagion effects		
Country	1997	Country	1997		Country	1997	
Thailand	47.9	Philippines	68.3	X**	Philippines	75.6	X
Indonesia	33.5	Thailand	15.1	X	Thailand	33.2	X
Malaysia	32.0	Turkey	15.1		Indonesia	26.2	X
Philippines	29.7	Malaysia	8.3	X	Malaysia	25.9	X
Korea	16.4	Netherlands	6.8		India	21.3	
Colombia	16.3	Jordan	6.5		Korea	17.8	
Turkey	12.4	Indonesia	6.2		Pakistan	14.3	X
Pakistan	8.4	Canada	5.4		Turkey	14.0	X
Brazil	7.3	Argentina	4.5		Sri Lanka	13.1	
Australia	4.6	Colombia	3.4	X	Colombia	10.0	X
Japan	4.4	Venezuela	3.3		Chile	8.5	
India	3.8	Germany	2.9		Australia	4.4	
Finland	3.8	Korea	2.9		Canada	2.9	X
Canada	2.9	Australia	1.0		Japan	2.6	X
Netherlands	1.9	U.K.	-1.4		Argentina	2.6	
Germany	1.3	Chile	-1.5		Netherlands	0.7	
U.S.A	1.2	Sri Lanka	-2.7	X	Norway	-1.8	
U.K	0.8	U.S.A	-3.6	X	U.K.	-1.8	X
Sri Lanka	0.3	Sweden	-3.7	X	Spain	-2.8	
Sweden	0.3	Finland	-4.7		Jordan	-3.4	
Mexico	0.1	Pakistan	-5.1		Germany	-4.1	
Chile	-0.5	Italy	-6.5		Italy	-7.5	
Argentina	-0.7	Norway	-7.3		U.S.A.	-9.3	
France	-1.5	Japan	-7.6		Finland	-10.7	
Jordan	-2.3	India	-7.7		Sweden	-10.9	
Spain	-2.9	Spain	-10.6	X	France	-13.6	
Norway	-4.5	France	-13.8		Mexico	-17.1	
Italy	-7.7	Mexico	-20.3		Brazil	-20.6	
Venezuela	-11.0	Brazil	-33.3		Venezuela	-22.0	X

* The currency crisis index (*Crisis*) is a weighted average of the percentage depreciation of nominal exchange rate with respect to the U.S. dollar and the percentage decrease in reserves.

** The country which can be matched with the one in the group of 5 countries which is clustered by the descending order of crisis index



* Predicted crisis index I and II are those which are predicted without and with contagion effects, respectively.

Figure 63. Actual and predicted currency crisis index

Table 55. Previous empirical studies

Study	Sample and frequency	Country coverage	Empirical method	Comments
Baig and Goldfajn (1998)	Jan. 1, 1995 - May 18, 1998, Daily	5 countries	VAR	- Cross-country correlations among currencies are found to increase significantly during the currency crisis.
Glick and Rose (1998)	1971- 1998, Annual	161 countries	Probit and OLS	- They show that currency crisis affect clusters of countries tied together by international trade. - The common external effects and trade links are studied.
Eichengreen, Rose and Wyplosz (1999)	1959 - 1993, Quarterly	20 industrial countries	Probit	- They seek to test for contagion in foreign exchange markets. - Trade links and macroeconomic similarities are tested separately. - The common external effects and trade links are investigated.
Masson (1998)	1994 - 1996, annual	13 emerging market countries	Index analysis	- The result of study suggests that the contagion effect unexplained by the common external effects and trade links played a major role in the Mexican and Asian crises.
Sachs, Tornell, and Velasco (1995)	1986 - 1995, Monthly and annual	20 emerging market countries	OLS	-The emphasis is on explaining why some countries were more affected by the Mexican crisis than others - The common external effects of contagion are investigated.
Tornell (1999)	1986 - 1997 Monthly and annual	22 emerging market countries	OLS	- The cross-country variation in the severity of the crisis in Mexico and Asia are explained. - The common external effects of contagion are investigated.
Valdes (1997)	Mar. 1986 - Aug. 1994 Monthly	7 Latin American countries	OLS	- Using debt prices and country credit ratings, this paper provides evidence of contagion in emerging markets.

CHAPTER VIII

CONCLUSION

This dissertation has studied the causes of the Asian currency crisis and improved the performance in predicting actual currency crises. The empirical analyses presented in this dissertation examined the Asian currency crisis using higher frequency data and more refined models than previous studies.

First, an extension of the structural currency crisis model is presented to derive shadow exchange rate and the probability of an exchange rate regime change. The model is a stochastic version of the monetary approach to exchange rate determination. While the nonstructural studies' results are not robust and do not forecast crises well, the results of our structural study indicate that weak fundamentals in the selected Asian countries, South Korea and Malaysia, prior to the Asian currency crisis already had been predicting the upcoming currency crisis.

Second, to analyze currency crises extensively, a pooled OLS using panel data is estimated, focusing on the importance of contagion effects on the eruption of a currency crisis. The empirical results show that a lending boom and contagion effects sufficiently explain the cross-country variation in the severity of the crisis of emerging markets. In addition, the prediction of currency crisis based upon the contagion effect is found to improve our ability to predict an eruption of currency crisis.

In Chapter II, an overview of the beginning and development in the Asian crisis is presented with a focus on the movements of the macroeconomic variables and the structural conditions of financial system. The evidence given in the overview indicates that deterioration in macroeconomic fundamentals and poor economic policies were a root cause of the crises. Nonetheless, the evidence is not convincing enough to establish that fundamentals had deteriorated so severely that the outbreak of the Asian currency crisis was inescapable.

Chapter III provides a survey of the theoretical and empirical literature on currency crises and introduces an extended structural currency crisis model. The theoretical literature about the currency crises contains a number of models which either support the 'fundamentalist' or 'non-fundamentalist' views as to the causes of currency crises. These two views are represented by first and second generation models of currency crises.

Based upon the logics of the theoretical models, the empirical literature has attempted to determine the actual sources of various currency crises. The empirical literature grouped into two categories: nonstructural and structural analyses. Nonstructural analyses exploited the high variability associated with cross-country information with the limitation by the lack of robustness to various sensitivity tests and poor performance in predicting actual crises. Beginning with the Blanco and Garber's (1986) study, structural analyses have presented strong evidence suggesting that domestic macroeconomic indicators play a key role in determining a currency crisis.

To determine whether the Asian crisis was distinct relative to other crises and improve the performance in the prediction of crises, the currency crisis is modeled with a structural model. As a result of the modeling, the influence of pure macroeconomic fundamentals on exchange market pressures for the Asian currencies can be evaluated. Furthermore, a spurious regression problem caused by the non-stationarities of relevant processes is resolved by using an error correction model (ECM) for the extended structural model.

Chapter IV then offers an analysis of the time series properties and forecasts of each variable of the structural currency crisis models introduced in Chapter III for the derivation of shadow exchange rates and probabilities of collapse. Unlike most previous studies, to capture the properties of economic and financial time series that exhibit long memory in both their conditional mean and variances, the ARFIMA(p,d,q)-FIGARCH(P,δ,Q) model is included in the selection of models in Chapter IV. Based on the Wald tests, the U.S. inflation rate, the percentage changes of deviations from PPP in Indonesia, the percentage changes of domestic credits in Malaysia and Thailand and the change rates of real GDP in Malaysia appear to have estimated long memory parameters d and δ which lie in the ranges of $-0.5 < d < 0.5$ and $0 < \delta < 1.0$, respectively. However, Wald tests do not find evidence for dual long memory behavior in other processes. Following the analysis of time series properties, the estimated parameters are used to forecast each variable. The forecasted variables are used to generate shadow exchange rates and probabilities of collapse in Chapter VI.

In Chapter V, long and short-run real money demand functions are estimated in the structural model used to derive the shadow exchange rates and probabilities of collapse. The previous structural analyses of currency crises estimated a real money demand function without taking into consideration the variable's non-stationarity. Therefore, the estimated money demand functions faced a spurious regression problem whereby conventional t-ratio and F significance tests could not be applied. To avoid these problems, cointegration and error correction techniques are applied to model a real money demand. The unit-root tests presented in the following sections detect non-stationarity of real money balance, real GDP, interest rate for South Korea and real money balance for Malaysia. Then, a residual based test was used to test for the number of cointegration relations and to estimate the cointegrating vectors. These analyses suggest that both long and short-run models can be specified in South Korea and Malaysia. Nonetheless, we cannot determine whether a model without a deterministic time trend can explain real money demand in South Korea. The use of monthly data enables this study to determine that long-run real money demand is unstable when modeled as a cointegrating relationship without a deterministic time trend in South Korea.

Chapter VI derives the shadow exchange rates and the probabilities of an exchange rate regime change for South Korea and Malaysia. Two countries experienced a severe devaluation during the Asian currency crisis. The shadow exchange rates and probabilities of collapse rely on earlier forecasts for relevant variables and estimates of the real money demand functions.

As shown in Figures 49 and 50, the estimated shadow exchange rate of each country signals the possibility of an upcoming severe depreciation by following the behavior of fundamentals. In particular, South Korea's shadow exchange rate was far above the regulated exchange rate since 1995. For Malaysia, the gap between the shadow and controlled exchange rates became noticeable around 1994. It seems that the gap between the shadow and actual exchange rates attracted speculators by the potential profit to be gained following a change in exchange regimes. The changes in the probability of collapse, which is positively correlated with the difference between shadow and actual exchange rate, indicate that both South Korea and Malaysia were under severe depreciation pressure starting in the early 1990s and continuing up until the Asian currency crisis.

In Chapter VII, a pooled OLS is estimated on panel data with a focus on the influence of contagion effects on the currency crises. The first result found is that lending booms impact the crisis index much more among developing countries than industrial countries. Under the strong and stable financial systems of industrial countries, the lending booms appears not to contribute to the variation of the crisis index. The second result is that contagion effects as represented by trade linkage between the initially attacked and home country as well as market sentiment are a significant source of the eruption of currency crises during the 1990s. Furthermore, the contagion effects are stronger in emerging markets than in industrial countries. Thirdly, evidence such as adjusted-R^2, $RMSE$, and matched number of countries supports that controlling for contagion effects improves the ability to predict a currency crisis.

Currency crises in the 1970s and 1980s were rooted in the dynamics of domestic credit extended by the central bank to the government. However, in the 1997 Asian currency crisis, domestic credit to the government did not play a crucial role. As shown in Chapter II, the domestic credit to the private sector enhanced by the domestic banks fueled by the foreign liabilities played a pivotal role in the latter currency crisis. In particular, as shown in Table 9, domestic bank lending to the private sector increased among Asian countries prior to the crises, leading to a sever lending boom. To make matters worse, many of the loans made by banks were invested in risky and low profit projects or used for real estate, property and the purchase of equity funds due to the moral hazard problem. The shadow exchange rates and probabilities of collapse derived in Chapter VI using *M2* as a money supply and the empirical results in Chapter VII that indicate the significance of lending booms confirm that lending booms played a key role in the eruption of the Asian currency crisis.

The 1994 Mexican crisis and the 1997 Asian crisis had a widespread contagion to several emerging markets. Currency crises were a regional phenomenon in the 1990s. As expected, the contagion effects are found to play a crucial role in the Asian currency crisis based on the empirical results in Chapter VI. In particular, strong trade linkage and market sentiment were essential in the eruption of Asian currency crisis.

In conclusion, given the evidence presented by the estimated shadow exchange rates and the probabilities of collapse of South Korea and Malaysia and by the empirical results in Chapter VII, weak fundamentals and contagion effects in the Asian countries had been indicating that a currency crisis could erupt whenever unexpected events such as bank failure, corporate failure and unstable political condition may trigger it in the Asian countries. Therefore, the implications for economic policy based on the Asian currency crisis are that countries that support sound macroeconomic and monetary policies are not readily vulnerable to a currency crisis and a country that has an initially attacked neighbor should perform more stable macroeconomic and monetary policies while trying to weaken the channel through which the contagion effects are delivered.

BIBLIOGRAPHY

Agenor, P.-R., J. S. Bhandari, and R. P. Flood (1992), "Speculative attacks and models of balance of payments crises," IMF Staff Papers 39, 357-394.

Akerlof, G. A. (1980), " The short run demand for money," The Economic Journal 90, 885-900.

Akerlof, G. A., and R. D. Milbourne (1980), "Irving Fisher on his head II: The Consequences of the timing of payments for the demand for money," The Quarterly Journal of Economics 95, 145-57.

Amsler, C. (1999), "Size and power: Lower tail KPSS tests and anti-persistent alternatives," Applied Economics Letters 6, 693-95.

Arize, A. C. (1994), "A Re-Examination of the demand for money in small developing economies", Applied Economics 26, 217-28.

Ariz, J., F. Caramazza, and R. Salgado (2000), "Currency crisis: In search of common elements", IMF Working Papers No. 00/67.

Baillie, R.T. (1996), "Long memory processes and fractional integration in econometrics", Journal of Econometrics 73, 5-59.

Baillie, R.T., C.-F. Chung, and M.A. Tieslau (1996), "Analyzing inflation by the fractionally integrated ARFIMA-GARCH model", Journal of Applied Econometrics 11, 23-40.

Baillie, R.T., and R.A. Pecchenino (1991), "Equilibrium relationships in international finance", Journal of International Money and Finance 10, 582-593.

Baillie, R.T., and T. Bollerslev (1992), "Prediction in dynamic models with time-dependent conditional variances", Journal of Econometrics 52, 91-113.

Baillie, R.T., and T. Bollerslev (1994), "Cointegration, fractional cointegration and exchange rate dynamics", Journal of Finance 49, 737-745.

Baillie, R.T., T. Bollerslev and H.O. Mikkelsen (1996), "Fractionally integrated generalized autoregressive conditional heteroskedasticity", Journal of Econometrics 74, 3-30.

Barro, R. J., and D. B. Gordon (1983), "Rules, discretion and reputation in a model of monetary policy", Journal of Monetary Economics 12, 101-21.

Barro, R. J., and S. Fischer (1976), "Recent development in monetary theory", Journal of Monetary Economics 2, 133-67.

Baumol, W. J. (1952), "The transactions demand for cash: An inventory theoretic approach," The Quarterly Journal of Economics 66, 545-56.

Bollerslev, T. (1986), "Generalized autoregressive conditional heteroskedasticity", Journal of Econometrics 31, 307-27.

Blanco, H., and P. M. Garber (1986), "Recurrent devaluation and speculative attacks on the Mexican Peso," Journal of Political Economy 94, 149-166.

Boorman, J. T. (1976), "The evidence on the demand for money: theoretical and empirical results," in current issues in monetary theory and policy, ed. by T. M. Havrilesky and J. T. Boorman (Arlington Heights, Illinois: AHM Public Corporation, 1976), 315-60.

Boughton, J. M. (1979), "Demand for money in major OECD countries," OECD economic outlook: occasional studies (Paris: Organization for economic cooperation and development, January 1979).

Boughton, J. M. (1981), "Recent instability of the demand for money: An international perspective," Southern Economic Journal 47, 579-97.

Calvo, G. A., and E. G. Mendoza (1997), "Rational herd behavior and the globalization of securities markets," mimeo, University of Maryland.

Chiang, C. Y. (1999), "The long-run demand for money functions in Taiwan (1961:4-1997:3) : Cointegration and stability," Ph.D. dissertation, Department of Economics, Michigan State University.

Chow, G. C. (1966), "On the long-run and short-run demand for money," Journal of Political Economy 74, 111-31.

Cooper, R. (1996), "Comment," Brookings Paper on Economic Activity 1, 203-208.

Corsetti, G., P. Pesenti, and N. Roubini (1998a), "Fundamental determinants of the Asian crisis: A preliminary empirical assessment," mimeo.

Corsetti, G., P. Pesenti, and N. Roubini (1998b), "Paper tigers?: A model of the Asian crisis," NBER Working Paper No. 6783.

Corsetti, G., P. Pesenti, and N. Roubini (1998c), "What caused the Asian currency and financial crisis? Part I and II," NBER Working Papers No. 6833 and 6834.

Craig, B., M. Eichenbaum, and S. Rebelo (1998), "Perspective deficits and the Asian currency crisis," NBER Working Papers No. 6758.

Cumby, R. E., and S. V. Wijnbergen (1989), "Financial policy and speculative runs with a crawling peg: Argentina 1979-1981," Journal of International Economics 27, 111-127.

Davidson, R., and J. G. MacKinnon, (1993), Estimation and inference in econometrics (Oxford: Oxford University Press, 1993).

Dickey, D.A., and W.A. Fuller (1979), "Distribution of the estimators for autoregressive time series with a unit root", Journal of American Statistical Association 74, 427-31.

Dickey, D.A., and W.A. Fuller (1981), "Likelihood ratio statistics for autoregressive time series with a unit root", Econometrica 49, 1057-72.

Dornbusch, R. (1998), "Asian crisis themes," mimeo, MIT.

Eichengreen, B., A. K. Rose, and C. Wyplosz (1996), "Speculative attacks on pegged exchange rates: An empirical exploration with special reference to the European monetary system," in M. B. Canzoneri, W. J. Ethier, and V. Grilli (eds.), *The New Transatlantic Economy*, New York: Cambridge University Press, 191-235.

Engel, R.F. (1982), "Autoregressive conditional heteroskedasticity with estimates of the variance of UK Inflation", Econometrica 50, 987-1008.

Engle, R.F., and C.W.J. Granger (1987), "Co-integration and error correction: Representation, estimation, and testing," Econometrica 55, 251-76.

Engel, R.F., and T. Bollerslev (1986), "Modeling the persistence of conditional variances", Econometric Reviews 5, 1-50.

Fair, R. C. (1987), "International evidence on the demand for money," The Review of Economics and Statistics 69, 473-80.

Feige, E. L., and D. K. Pearce (1994), "The substitutability of money and near-monies: A survey of the time-series evidence," Journal of Economic Literature 15, 439-69.

Fisher, I. (1922), The purchasing power of money, its determination and relation to credit interest and crises by I. Fisher assisted by H. G. Brown. (New York: Macmillan, 1922).

Fischer, A. M. (1993), "Is money really exogenous? Testing for weak exogeneity in Swiss money demand," Journal of Money, Credit, and Banking 25, 248-58.

Fischer, S. (1975), "The demand for index bonds," Journal of Political Economy 83, 509-34.

Flood, R. P., and R. J. Hodrick (1986), "Real aspects of exchange rate regime choice with collapsing fixed rates," Journal of International Economics 21, 215-32.

Flood, R. P., and N. Marion (1998), "Perspectives on the recent currency crisis literature," NBER Working Paper No. 6380.

Flood, R. P., and P. M. Garber (1984a), "Collapsing exchange-rate regimes: Some linear examples," Journal of International Economics 17, 1-13.

Flood, R. P., and P. M. Garber (1984b), "Gold monetization and gold discipline," Journal of Political Economy 92, 90-107.

Frankel, J. A., and A. K. Rose (1996), "Currency crashes in emerging markets: An empirical treatment," Journal of International Economics 41, 351-66.

Friedman, M. (1956), "The quantity theory of money-a restatement," in studies in the quantity theory of money, ed. by M. Friedman (Chicago: University of Chicago Press, 1956).

Fuller, W.A. (1976), Introduction of statistical time series, (Wiley: New York, N.Y. 1976).

Furman, J., and J. E. Stiglitz (1998), " Economic crises: Evidence and insights from Asia", Brookings papers on economic activity 0, 1-114.

Ghysels, E.(1990), "Unit-root tests and the statistical pitfalls of seasonal adjustment: The case of U.S. postwar real gross national product", Journal of Business and Economic Statistics 8, 145-348.

Glick, R., and A. K. Rose (1998), "Contagion and trade: Why are currency crisis regional?", NBER Working Paper No. 6806.

Goldberg, L. S. (1988), "Collapsing exchange rate regimes: A theoretical and empirical investigation," Ph.D. dissertation, Department of Economics, Princeton University.

Goldberg, L. S. (1994), "Predicting exchange rate crises: Mexico revisited," Journal of International Economics 36, 413-430.

Goldfeld, S. M. (1973), " The demand for money revisited," Brookings Papers on Economic Activity 3, 577-646.

Goldfeld, S. M., and D. E. Sichel (1990), "The demand for money," in Handbook of Monetary Economics, Volumn I, ed. by B.M. Friedman and F.H. Hahn (Amsterdam: Elsevier Science Publishers B.V, 1990), 300-56.

Gonzalo, J. (1994), "Five alternative methods of estimating long-run equilibrium relationships", Journal of Econometrics 60, 203-233.

Goodfriend, M. (1985), "Reinterpreting money demand regressions," Carnegie-Rochester Conference Series on Public Policy 22, 207-42.

Gordon, R. J. (1984b), "The short-run demand for money: A reconsideration," Journal of Money, Credit, and Banking 16, 403-34.

Granger, C.W.J. (1980), "Long memory relationships and the aggregation of dynamic models", Journal of Econometrics 14, 227-238.

Granger, C.W.J. (1981), "Some properties of time series data and their use in econometric model specification", Journal of Econometrics 16, 121-130.

Granger, C.W.J. (1986), "Developments in the study of cointegrated economic variables", Oxford Bulletin of Economics and Statistics 48, 213-28.

Granger, C.W.J., and R. Joyeux (1980), "An introduction to long memory time series models and fractional differencing", Journal of Time Series Analysis 1, 15-39.

Hansen, B.E. (1992), "Efficient estimation and testing of cointegrating vectors in the presence of deterministic trends", Journal of Econometrics 53, 87-121.

Hansen, B.E. (1992), "Efficient estimation and testing of cointegrating vectors in the presence of deterministic trends", Journal of Econometrics 53, 87-121.

Hendry, D. F. (1986), "Econometric modeling with cointegrated variables: An overview", Oxford Bulletin of Economics and Statistics 48, 201-12.

Hurst, H.E. (1951), "Long-term storage capacity of reservoirs", Transactions of the American Society of Civil Engineers 116, 770-799.

Hylleberg, S. (1995), "Tests for seasonal unit roots: General to specific or specific to general?", Journal of Econometrics 69, 5-25.

Jarque, C.M., and A.K. Bera (1980), "Efficient tests for normality, homoscedasticity and serial independence of regression residuals", Economic Letters 6, 255-59.

Jeanne, O, (1997), "Are currency crises self-fulfilling? A test," Journal of International Economics 43, 263-286.

Johansen, S. (1988b), "Statistical analysis of cointegration vectors," Journal of Economic Dynamics and Control 12, 231-54.

Johansen, S. (1991), "Testing weak exogeneity and the order of cointegration in UK money demand data", Journal of Policy Modeling 14, 313-334.

Johansen, S. (1992), "Estimation and hypothesis testing of cointegration vectors in gaussian vector autoregressive models", Econometrica 59, 1551-80.

Johansen, S., and K. Juselius (1990), "Maximum likelihood estimation and inference on cointegration – with application to the demand for money", Oxford Bulletin of Economics and Statistics 52, 169-210.

Johansen, S., and K. Juselius (1992), "Testing structural hypotheses in a multivariate cointegration analysis of the PPP and the UIP for the UK", Journal of Econometrics 53, 211-44.

Judd, J. P., and John L. Scadding (1982), "The search for a stable money demand function: A survey of the post- 1973 Literature," Journal of Economic Literature 20, 993-1023.

Kaminsky, G. L., and C. M. Reinhart (1999), "The twin crises: The cause of banking and balance-of-payments problems," American Economic Review June, 473-500.

Kaminsky, G. L., S. Lizondo, and C. M. Reinhart (1997), "Leading indicators of currency crisis," IMF Working Papers No. 97/79.

Keynes, J. M. (1930), A treatise on money (London and New York: Macmillan, 1930).

Keynes, J. M. (1936), The general theory of employment, interest, and money (London and New York: Macmillan, 1936).

Kim, S. (2000), "What caused the Asian currency crisis?," Ph.D. dissertation, Department of Economics, Michigan State University.

Klein, M. W., and N. P. Marion (1997), "Explaining the duration of exchange-rate pegs," Journal of Development Economics 54, 387-404.

Krugman, P. (1979), "A model of balance-of-payments crises," Journal of Money, Credit and Banking 11, 311-325.

Krugman, P. (1998), "What happened to Asia?," mimeo, MIT.

Kwiatkowski, D., P.C.B. Phillips, P. Schmidt, and Y. Shin (1992), "Testing the null hypothesis of stationary against the alternative of a unit root: How sure are we that economic time series have a unit root?", Journal of Econometrics 54, 159-178.

Kwon, T. (1997), "Long memory modeling of conditional means and variances with applications to asset pricing," Ph.D. dissertation, Department of Economics, Michigan State University.

Kydland, F.E., and E.C. Prescott (1977), "Rules rather than discretion: The inconsistency of optimal plans", Journal of Political Economy 85, 473-91.

Laidler, D. E.W. (1985), The demand for money: Theories and evidence (New York: Dunn-Donnelley, 3rd edition, 1985).

Lee, H.S., and C. Amsler (1997), "Consistency of the KPSS unit root test against Fractionally integrated alternative," Economic Letters 55, 151-60.

Ljung, G., and G. Box (1979), " On a measure of lack of fit in time series models", Biometrika 66, 265-270.

Lucas, R. E., Jr. (1980), "Equilibrium in a pure currency economy," Economic Inquiry 18, 203-20.

Marshall, A. (1923), Money, credit and commerce (London: Macmillan, 1923).

Masson, P. R. (1998), "Contagion effects: monsoonal effects, spilllovers, and jumps between multiple equilibria," mimeo, IMF.

Milbourne, R. (1983), "Optimal money holding under uncertainty," International Economic Review 24, 685-98.

Milbourne, R. (1988), "Disequilibrium buffer stock models: A survey," Journal of Economic Surveys 2, 187-208.

Milbourne, R., P. Buckholtz, and M.T. Wason (1983), "A theoretical derivation of the functional form of short-run money holdings," The Review of Economic Studies 50, 531-41.

Miller, M. H. (1968), " The demand for money by firms: Extensions of analytic results," Journal of Finance 23, 735-59.

Miller, M. H., and D. Orr (1966), " A model of the demand for money by firms," The Quarterly Journal of Economics 80, 413-35.

Morris, S., and H.S. Shin (1995), "Informational events that trigger currency attacks," Working paper No.95-24, Federal Reserve Bank of Philadelphia.

Nelson, C.R. and C.I. Plosser (1982), "Trends and random walks in macroeconomic time series", Journal of Monetary Economics 10, 139-162.

Obstfeld, M. (1986), "Rational and self-fulfilling balance-of-payments crises," American Economic Review 76, 72-81.

Obstfeld, M. (1994), "The logic of currency crises," Cahiers Economiques et Monetaires 43, 189-213.

Obstfeld, M. (1996), "Models of currency crises with self-fulfilling features," European Economic Review 40, 1037-1047.

Obstfeld, M. (1997), "Destabilizing effects of exchange-rate escape clauses," Journal of International Economics 43, 61-77.

Otker, I., and C. Pazarbasioglu (1996), "Speculative attacks and currency crisis: The Mexican experience," Open Economies Review 7, 535-552.

Otker, I., and C. Pazarbasioglu (1997a), "Likelihood versus timing of speculative attacks: A case study of Mexico," European Economic Review 41, 837-845.

Otker, I., and C. Pazarbasioglu (1997b), "speculative attacks and macroeconomic fundamentals: Evidence from some European countries," European Economic Review 41, 847-860.

Park, D., and C. Rhee (1998), "Currency crisis in Korea: Could it have been avoided?," mimeo.

Park, Y. C. (1996), "East Asian liberalization, bubbles, and the challenges from China," Brookings Papers on Economic Activity 2, 357-371.

Patinkin, D. (1965), Money, interest, and prices: An integration of monetary and value theory (New York: Harper and Row, 2nd edition, 1965).

Perron, P. and T.J. Vogelsang (1992), "Nonstationarity and level shifts with an application to purchasing power parity", Journal of Business and Economic Statistics 10, 301-320.

Phillips, P.C.B. (1987), "Time series regression with a unit root", Econometrica 55, 277-301.

Phillips, P.C.B. and B.E. Hansen (1990), "Statistical inference in instrumental variables regression with I(1) Processes", Review of Economic Studies 57, 99-125.

Phillips, P.C.B. and P. Perron (1988), "Testing for a unit root in time series regression", Biometrika 75, 335-346.

Phillips, P.C.B. and S. Ouliaris (1990), "Asymptotic properties of residual based tests for cointegration", Econometrica 58, 165-193.

Pigou, A.C (1917), "The value of money", The Quarterly Journal of Economics 37, 38-65.

Radelet, S., and J. D. Sachs (1998a), "The east Asian financial crisis: Diagnosis, remedies, prospects," Brooking Papers on Economic Activity 1, 1-74.

Radelet, S., and J. D. Sachs (1998b), "The onset of the east Asian financial crisis," NBER Working Paper No. 6680.

Rose, A. K. (1985), "An alternative approach to the American demand for money", Journal of Money, Credit, and Banking 17, 439-55.

Sachs D. J., A. Tornell, and A. Velasco (1996), "Financial crises in emerging markets: The lessons from 1995," Brookings Papers on Economic Activity 1996, 147-198.

Salant, S.W., and D.W. Henderson (1978), "Market anticipation of government gold policies and the price of gold", Journal of Political Economy 86, 627-648.

Samuelson, P. A. (1958), " An exact consumption loan model of interest with or without the social contrivance of money", Journal of Political Economy 66, 467-82.

Schwert, G.W. (1989), "Test for unit roots: A Monte Carlo investigation", Journal of Business and Economic Statistics 7, 147-59.

Sowell, F.B. (1992), "Maximum likelihood estimation of stationary univariate fractionally integrated time series models", Journal of Econometrics 53, 165-188.

Sriram, S. S. (1999a), "Demand for M2 in an emerging-market economy: An error-correction model for Malaysia", IMF Working Paper No. 99/173.

Sriram, S. S. (1999b), "Survey of literature on demand for money: Theoretical and empirical work with special reference to error-correction models", IMF Working Paper No. 99/64.

Stock, J.H., and M.W. Watson (1988), "Testing for common trends", Journal of the American Statistical Association 83, 1097-1107.

Tan, E.U. (1997), "Money demand amid financial sector developments in Malaysia," Applied Economics 29, 1201-1215.

Tobin, J. (1956), "The interest-elasticity of transactions demand for cash," The Review of Economics and Statistics 38, 241-47.

Tobin, J. (1958), "Liquidity preference as behavior towards risk," The Review of Economic Studies 25, 65-86.

Tornell, A. (1999), "Common fundamentals in the tequila and Asian crises." NBER Working Paper No. 7139.

Whalen, E. L.(1966), "A rationalization of the precautionary demand for cash," The Quarterly Journal of Economics 80, 314-24.

Willman, A. (1988), "The collapse of the fixed exchange rate regime with sticky Wages and imperfect substitutability between domestic and foreign bonds," European Economic Review 32, 1817-1838.

Wooldridge, J. M., (2000), Introductory Econometrics: A modern approach. South-Western College Publishing.

Yoshida, T. (1990), "On the stability of the Japanese money demand function: Estimation results using the error correction model," Bank of Japan Monetary and Economic Studies 8, 1-48.