CURRENCY CRISES: ORIGIN, PUBLIC DEBT AND CONTAGION

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To Minghua Cun, my Mum

—an ordinary, hardworking

Chinese woman who has always had a BIG dream for her rather slow son.
When I arrived at Heathrow International Airport with an offer letter from University of Surrey on a sunny morning three years ago, I knew I was embarking on a challenging but also exciting journey. What I did not expect at that moment was that there should be so many individuals who, by generously providing me with their assistance and encouraging advice, would turn this journey into a time full of fun and rewards.

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whose presence in my life has, in one way or the other, contributed to my well-being but I know they recognize themselves in these lines. Last, but certainly not least, I would like to thank my motherland China for giving me the willingness to work hard and a spirit of integrity and bravery.
The frequency of currency crises has increased drastically in the last 30 years, and the scale and impact of these episodes has greatly renewed interest in the existing literature and stimulated a growing volume of new literature on both theoretical and empirical sides. These offer quite different explanations for currency crises which the thesis attempts to assess.

Firstly, this project looks at the experiences of countries in Latin America, Europe and East Asia during 1970-2001. The aim is to examine how good the so-called 'three generation' crisis theoretical models are in explaining the origins of different currency crises where they were originally inspired. Instead of imposing an arbitrary threshold a priori, we employ the non-linear SETAR (Self Exciting Threshold Autoregression) model and Markov-switching model to identify the crisis incidents. The Limited Dependent Variable Model—Logit Model is adopted to link the defined crisis variable to a wide range of selected variables based on crisis theories and empirical works. The estimated results show limited evidence that the three generation models efficiently address the different features of currency crises in three regions. The overall imperfect picture lays down the task for our further research.

In the following chapter, we focus on the issue of credibility within the approach of the second-generation models. In an extended model, we show that whether a
strenuous defence of a parity would boost its credibility is ambiguous and depends on the importance of the 'debt-burden' effect relative to the 'signalling' effect. In other words, resisting a crisis could show the authority's resolution not to devalue, and thus enhance the expectation that the regime will be maintained in the future. But at the same time, refraining from inflationary financing increases the debt burden and hence the likelihood of a forced future devaluation. Two testable predictions are derived regarding the impact of the debt maturity on the interest rates in two different types of countries—the debt-burden countries and signalling countries. We distinguish between them in our sample by observing different patterns in the interest rates when countries experience actual crises and/or successful defences. The actual devaluations are identified by the regime changes in the Reinhart-Rogoff index, while the successful defences are periods that are under severe attack (defined by both the SETAR and MRS model) but do not lead to actual crises. By using the GMM (General Method of Moments) for the dynamic panel data model, the estimated results show supportive evidence for the predictions.

Finally, we re-examine the crisis transmission mechanism during the 1997/98 East Asia Crises. It is argued that the tests for contagion based on the conditional cross-market correlation coefficient are inaccurate and biased due to heteroscedasticity. We adopt the proposed formula to adjust for the bias and conduct our tests, for the first time, in the framework of a Markov-switching Vector Autoregressive (MS-VAR) model. It is shown that the adjustment has a significant impact on the tests, and the evidence of contagion that is found in the initial tests based on the conditional correlation coefficients soon disappears, after we adjust for heteroscedasticity. Instead, the highly correlated comovements amongst the East Asian markets during the 1997/98 crises is only a continuation of the strong economic linkages—interdependence between them—which exist in all states of the world.
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In the last 30 years many countries in the world, developed and developing alike, have experienced fierce speculative attacks on their currencies. The European Monetary System (EMS) was severely undermined by such attacks in 1992/93 — the cause of Great Britain's and Italy's exit from the EMS in 1992, and the widening of the Exchange Rate Mechanism (ERM) bands for most of its member nations in 1993.

In December 1994, following a speculative attack on the Mexican peso, the Mexican government abandoned the crawling peg, leaving the currency to devalue. The newly industrializing countries in East Asia, once regarded as the most economically successful in the world, abruptly fell into financial/currency turmoil in 1997. A new phenomenon, which became the most striking aspect of the East Asia crises, was the resultant rapid crash that spreads from the Thai exchange market to other countries in the region, and to Russia and Brazil later in 1998.

More recently, turbulence in Argentina's financial market has pushed its government to default on foreign debts at the beginning of 2002, and also slow its recovering economy. The scale and impact of these events have to a large extent renewed interests in the existing literature on the subject, and stimulated an growing volume of
new literature on both theoretical and empirical work to offer different explanations for these apparently ever-changing phenomena.

Two questions are raised in my research:

1. "How good are the existing crisis theories (the so-called ‘three generation’ models) at explaining the currency crises?"

2. If they are not good enough, then “is there any area where we can make improvement in either/both theories and empirical works?”

The latter question covers two different points:

- Revisit the issue of credibility of an exchange rate regime in second-generation crisis theories, in order to see how current actions affect future policy through the channel of Public Debt.

- Re-examine the crisis transmission mechanism in the 1997 East Asia Currency/Financial Crises, after adjusting for the bias due to heteroscedasticity, as proposed by Forbes and Rigobon (2002).

However, before proceeding to a detailed discussion of currency crises, it is necessary to clarify the concept of ‘currency crisis’ used in this thesis. Popular wisdom refers to currency crises as the episodes when domestic currencies come under severe pressure to be devalued. These episodes may or may not result in actual devaluations; for example, governments facing fierce speculative attacks can avoid devaluation by selling foreign reserves and/or raising interest rates.

Therefore, many papers develop a two-step classification to identify currency crises (see Eichengreen et al. (1996a), Eichengreen et al. (1996b), Frankel and Rose (1996), and Kaminsky and Reinhart (1996) amongst others). Firstly, the vulnerability of a country’s exchange rate regime is gauged by an exchange market pressure index, which comprises changes in the exchange rate, interest rate, and reserves. At the second stage, a crisis is identified if the index exceeds a certain threshold level. The problem is that the choice of the threshold is arbitrary, i.e. 10%, or one
standard deviation. Throughout this thesis, however, the bias will be avoided by adopting non-linear state models that allow the data to reveal whether a given move represents abnormal behaviour and thus signal a crisis.

As will be shown in the Chapter 2 Literature Review, the existing crisis literature has been extensive and offered various explanations for currency crises occurring around the world in the last three decades. The initial approach of 'bad' fundamentals (put forward by Krugman (1979) and modified by Flood and Garber (1984)) indicates that currency crises result from expansionary fiscal and/or monetary policies. These classic speculative attack models have been tested, and are considered good fits to the experiences of many Latin American countries during 1970s and 1980s.

But in the early 1990s, the fact that some speculative attacks took place in European economies without large apparent monetary and fiscal imbalances prompted the emergence of so-called 'second-generation' crisis models (i.e. Obstfeld (1994), Obstfeld (1996)). These models suggest that conflicting goals of governmental policies, when combined with investors' 'rational' perceptions, can create multiple equilibria. A crisis is created when the economy switches from a 'good' equilibrium to a 'bad' one and the process can be self-fulfilling, even when the fundamentals are sound.

Most of the classic models failed to predict the East Asia Currency/Financial Crises which suddenly broke out in 1997/98. As a result, the 'third-generation' models were developed. Krugman (1998) suggests that the problem of 'moral hazard' and the resultant international over-borrowing play an important role, acting as a main force in creating bubbles in asset markets. Change and Velasco (1998)'s illiquidity model shows that a financial (currency) crisis can occur if a large number of agents lose their confidence in the banks and withdraw their money at the same time, because these actions can force the liquidation of more profitable long-term investments to compensate for the temporary shortfall.

But the question is: how good are the three generation crisis models when ad-
dressing the origins of currency crises globally? Chapter 3 will examine the crisis episodes in three different country groups — Latin America, Europe and East Asia — looking at the distinctive features of crises in each region, as defined by the three generation models.

As mentioned above, the Market Pressure Index (MPI) is established to gauge the vulnerability of a country’s exchange rate to speculative attacks. As opposed to imposing an arbitrary threshold, i.e. 10% or a one standard deviation move in the MPI, the SETAR (Self Exciting Threshold Auto-regression) model and the Markov-switching model are employed to allow data itself to signal crises. The SETAR model allows the MPI to switch from a state of ‘tranquil’ to a state of ‘turmoil’, once it exceeds the threshold level that is chosen to deliver the lowest standard error in the maximum likelihood estimation. The Markov-switching model assumes that the evolution of the state follows a first order Markov process.

Based on the definition of currency crises above, we examine a wide range of economic factors as potential causes for the crashes in the three regions, looking in particular at the variables recommended by the three generation crisis models.

The work adopts the ‘Limited Dependent Variable’ model, more specifically the ‘Logit’ model, to link the probability of a crisis to the relevant economic variables. As the data used is annual, a one-period lag of the explanatory variables is also included in the model. Finally, by dropping all explanatory variables that are not statistically significant at 10% level from our primary regression, we arrive at a set of final results (discussed in Chapter 3). The study is an attempt to adopt an objective approach by using both the SETAR and Markov-switching model to examine the crisis episodes in Latin America, Europe and East Asia.

The estimated results show limited evidence that the three generation models efficiently address the different features of currency crises in three regions respectively. In other words, there is the possibility for improvement; we start by looking at the issue of credibility of the exchange rate regime within the second-generation framework.
From the traditional view of a monetary authority's reputation, resisting a currency crisis sends a strong signal of resolve and thus enhances the credibility of a fixed exchange rate. But Benigno and Missale (2001) point out that current actions affect future policy. Therefore, defending the currency and refraining from inflationary financing increases the debt burden, impairing the authority's ability to withstand future crises. Therefore, whether an exchange rate regime gains or loses credibility after a successful defence depends on whether the monetary authority's reputation or fundamentals dominates, i.e. the level and maturity of public debt.

Drazen and Masson (1994) suggested in an earlier paper that the persistent effect of unemployment could be another factor, showing that maintaining a fixed parity leads to higher future unemployment thus constraining the scope to manoeuvre in the future.

Following Benigno and Missale (2001)'s approach, we adopt a three-period stochastic version of the Barro and Gordon (1983) model, where the probability of devaluation in each period is derived from the optimal problem facing the monetary authority. Three cases are studied:

Firstly, the case is considered where the central bank can commit itself to a 2-period rule, which is set up in period 0 and publicly known. We obtain the socially optimal solution as a benchmark for following studies. Secondly, we study the case where the monetary authority cannot make a commitment and its type is known to the public ('complete information'). Finally, we examine the case where the monetary authority cannot commit itself but its preferences are not known to private sectors ('asymmetric information').

We also extend the original model to distinguish between deflatable and non-deflatable debt. We show that when the type of monetary authority is not-publicly known, the relative importance of the 'debt-burden' effect and 'signaling' effect are crucial in determining the probability of a future crisis after a current crisis, and the term structure of interest rates.

In a pioneering methodology, we combine the Reinhart-Rogoff index (which gives
actual changes in exchange rate regimes) with the SETAR and Markov-switching model (which could identify the periods of high pressure in exchange rate markets). The periods of severe attacks defined by both the SETAR and Markov-switching model but not accommodated in the Reinhart-Rogoff index are differentiated as successful defences against the actual crises. Based on the different patterns shown by the domestic interest rate differential (relative to the USA, a proxy for the likelihood of a devaluation) after a devaluation and/or a successful defence, 11 emerging countries are split into two groups: one where the 'debt-burden' effect dominates, and the other where the 'signaling' effect prevails. Finally, two predictions of our theoretical models are tested on the two groups by adopting the GMM (Generalized Method of Moments) model for dynamic panel data; the results are quite promising.

The 1997/98 East Asia Crises showed that a sharp devaluation of one country's currency could have a profound impact on the vulnerability of currencies in other economies in the region. In an attempt to make a contribution to the crisis contagion literature, we re-examine whether these highly correlated co-movements in the exchange markets give evidence of contagion. This depends on how we define the term 'contagion'. In the thesis, we follow the conventional approach to give a narrow definition of contagion: a significant increase in cross-market linkages after a shock to one country (or countries). Therefore, if the cross-market correlation does not increase significantly during periods of turmoil compared with the relatively stable periods, then any continued level of correlation suggests that there is an interdependence that exists in all states of the world.

Many economists, basing their tests on estimated conditional correlation coefficients, find there are evidences suggesting that contagions occur in several crisis episodes (including the 1997/98 East Asia Crisis). But as Forbes and Rigobon (2002) argue, the cross-market correlation coefficient is conditional on market volatility. Therefore, during crises when markets are more volatile, the estimators of the correlation coefficient tends to increase and be biased upward. This is the reason that so many tests traditionally find evidence of contagion.
Chapter 5 of this thesis, following Forbes and Rigobon (2002)'s approach, corrects the bias due to heteroscedasticity before testing for contagion. Our analysis focuses on the period from 1991 to 1998 for five East Asian countries, where Thailand is chosen as the country that spawned the original crisis. As mentioned above, the Market Pressure Index (MPI) is adopted to represent not only the successful attacks (actual crises) but also the unsuccessful attacks on local currencies defended by loss in foreign reserves and increased interest rates.

The main contribution of chapter 5 lies in the adoption of the non-linear Markov-switching VAR framework to deliver a better representation of the crisis transmission mechanism. The rationale behind this is that during a long period, the macroeconomic or financial variables, such as Market Pressure Index (MPI), sometimes undergo episodes in which the series behaviour seems to change quite dramatically, e.g. currency crises. Unlike the linear models, the Markov-switching model assumes that observed data is drawn from different distributions conditional on the state-contingent parameter sets. The model switches amongst different states, following the Markov chain, which means that the current state only depends on the state in the last period — ‘let the history tell the future’. The results show that the hypotheses of linear specification are rejected at even 1% significant level in all cases, favouring the Markov-switching VAR models. Several exogenous variables are also included in our models to capture the fundamental changes and/or external shocks.

We find that the adjustment has a significant impact on the estimated cross-market correlation coefficients, and hence on the test results. The evidence of contagion that is found in the initial tests based on the conditional correlation coefficients, soon disappears after we adjust for heteroscedasticity.

The material in the thesis is organized as follows. Chapter 2 surveys the available literature on currency crises. We begin our literature review with a presentation of the so-called three generation crisis theoretical models, followed by four different empirical approaches. We also include some contemporary issues, e.g. crisis contagion. As the volume of work on currency crises has grown so significantly in the last
two decades, presenting the main contributions in a systematic way is a challenging task. However, it is also rewarding, as the progress made in the understanding of currency crises in recent years is particularly encouraging.

Chapter 3 is an empirical section. Firstly, we establish a Market Pressure Index (MPI) to represent both successful and unsuccessful attacks, and define currency crises by using the SETAR model and Markov-switching model. The potential factors for currency crises are chosen according to economic theories and empirical works. The Logit model is employed to identify the distinctive features of currency crises in three regions: Europe, Latin America and Asia.

In Chapter 4, we firstly set up a theoretical framework for a revised Barro-Gordon model. Three cases are then studied respectively: the Commitment case, which gives the benchmark, Discretion under Complete Information, and Discretion under Asymmetric Information. The results are demonstrated by numerical simulation using Matlab. The sample of 11 emerging countries is split into two groups according to their different behaviours—debt-burden countries and signalling countries. Finally, the two predictions drawn up from our theory are tested on the two country groups.

Chapter 5 firstly gives the definition of crisis contagion and review the relevant literature. The bias in the conventional approach using the conditional correlation coefficient to test for contagion is identified and quantified, and an adjustment method is presented. The Markov-switching VAR model is adopted to examine the East Asian countries' experiences during the 1997/98 crises, after adjusting for heteroscedasticity.

The implementation of the econometric analysis in chapters 3, 4 and 5 has involved the collection of a variety of data. Every effort has been made to ensure that we collect the information in a consistent way and from reliable sources. The sources of data extraction are cited in each chapter. Detailed information on the construction of variables is also provided in the relevant chapters. Moreover, at various stages throughout the report the adopted econometric methodologies are
discussed, and their use justified.

Finally, the main findings of this research project are discussed and summarized in Chapter 6. We also acknowledge the limitations of our analysis and outline directions for future research.
CHAPTER 2

Literature Review

2.1 Introduction

During the last three decades, the frequency of currency crises has increased dramatically (see Bordo et al. (1998)). Events such as the Latin America Crises in the late 1970's and 1980's, the collapse of the European Exchange Rate Mechanism (ERM) in 1992/93, the 1994 Mexican peso crisis, the East Asia Financial/Currency Crises in 1997/98, and the more recent turmoil in Argentina are all good examples. The scale and impact of these incidents has inspired interest in the existing crisis literature, and stimulated a large volume of new literature (including both theories and empirical works) to provide different explanations for currency crises—the apparently ever-changing phenomena.

The initial speculative attack models were inspired by the crisis episodes experienced by the Latin American countries during the late 1970's and 1980's. This approach of 'bad' fundamentals (put forward by Krugman (1979) and modified by Flood and Garber (1984)) indicates that crises could be caused by expansionary fis-
cal and/or monetary policies, resulting in unsustainable credit expansion that was common amongst many Latin American economies at that time. However, the fact that some crises took place in Europe without either weak economic fundamentals or irresponsible credit expansion in the early 1990's, prompted the emergence of the so-called 'second generation' crisis models (Obstfeld (1994), Obstfeld (1996)). They emphasize that the conflicting goals of government policies when combined with investors' rational expectations can create multiple equilibria, where a crisis can be self-fulfilled even when the fundamentals are sound—move from a 'good' equilibrium where the fixed exchange rate regime is feasible to a 'bad' equilibrium where it is not.

The debate on the causes of crises continued until 1997 when the third generation models emerged to explain the factors that pushed the East Asian economies, once regarded as the world's most dramatic economic successes, into deep turmoil. The problems of moral hazard (international over-borrowing) and the liquidity crunch experienced once the currency regime collapsed are believed to have played a major part in the crises. The speed that crisis spread from one country to another has also raised interest in studies of contagion.

On the empirical side, there are in general four different approaches analyzing currency crises. The first is the 'signalling' method, where the behaviours of a number of individual variables are evaluated against certain threshold levels. Once any of these indicators moves beyond its respective threshold, it signals a potential crisis in waiting. The 'optimal' threshold is selected to balance out the risks of failing to predict the crisis and giving a false signal of an impending crisis.

The second methodology is the discrete-choice technique that analyzes the probability of having a currency crisis. This has been the most widely-used method until recent years when the non-linear changing regime (state) models are introduced into crisis literature. Different countries and time periods are firstly sorted into two discrete states: a crisis and a stable period. Then, a Logit or Probit model is employed to link the binary crisis variable to a range of potential factors chosen on the basis
of *a priori* economic theories. The likelihood of having a crash can therefore be evaluated, given a set of economic variables. The third method is largely descriptive and often based on specific case studies, where authors try to establish structural relationships between the selected variables and crisis incidents.

The last approach that is increasingly popular in the crisis empirical literature is the changing regime models. In models like the Markov-switching model, we are allowed to assume that the time series variable could be drawn from different distributions, conditional on state-contingent parameter sets. By adopting the Markov-switching model, we could have a continuous dependent variable, and also avoid the arbitrary choice of a threshold level for currency crises.

The literature grows at such a speed that it is very difficult to keep track of all theoretical and empirical developments. So while this chapter gives an overview of the theoretical approaches and empirical methodologies that appear to be relevant, it concentrates on key articles in each area. The rest of the chapter is organized as follows. In Section 2, we review the three generation theoretical models, noting their implications in the using of certain variables as potential factors in crises. Section 3 presents the four different approaches adopted by economists in building up empirical models. Finally, we summarize and give direction to our research in the following chapters.

### 2.2 The Three Generation Crisis Models

There are three generations of theoretical models in the currency crisis literature: the initial approach indicates that irresponsible government policies and unsound economic fundamentals are potential causes for crises; the second generation models emphasize the contingent nature of the government that may give rise to multiple equilibria, when combined with private sector's perfect foresight; and the latest method blames the problems of moral hazard (international over-borrowing) and the liquidity crunch for the currency regime collapses.
2.2.1 The First-generation Crisis Models

The first-generation crisis model is inspired by the literature on government price-fixing schemes in exhaustible resource markets (Salant and Henderson (1978), Salant (1983)). In the so-called 'New International Economic Order', Salant argues that any policy made by an official price stabilization board aiming to buy or sell the exhaustible resource (e.g. gold) at a fixed price would be subject to devastating speculative attacks. It is assumed in his paper that because gold in the ground yields no service flow and costs nothing to extract, its price must rise in line with the real rate of interest.

Suppose the government pegs the price of gold at some level between the initial price and the zero-demand choke price. At the beginning, gold owners will sell their entire stock to the government because they can earn a rate of return—the real interest rate—by placing their wealth in bonds instead of gold. For a time, industrial and personal demands therefore will be supplied entirely by the government, who must sell any amount of its reserve at the fixed price. It is clear, however, that this situation is unsustainable: eventually the stock will be depleted and the equilibrium price will have to be at its choke level.

In the model, the shadow free-market price of gold is introduced to show the competitive market price that would prevail in the absence of price fixing, whose path is derived by the well-known Laissez-faire perfect-foresight solution. The date at which the two prices become equal is the date at which the price-fixing scheme collapses: it does so after a speculative attack in which private market participants acquire all of the remaining official gold stock at the fixed price. The increase in private stock is justified, 

The initial approach of the speculative attack model in crisis literature was first put forward by Krugman (1979) and modified by Flood and Garber (1984). Krugman (1979) applies similar logic to speculative attacks not on a government using
a stockpile of an exhaustible resource to stabilize its price, but on a central bank using foreign reserves to peg an exchange rate. In his model, the central bank has to intervene to support the fixed rate under the regime, and a speculative attack on a government’s reserves can be viewed as a process by which investors change the composition of their portfolios. In other words, investors would reduce the proportion of domestic currency which they hold and buy the foreign currency from the authority, once they expect a potential possibility of a depreciation of domestic currency.

If the government runs excessively expansionary fiscal policies and large budget deficits are financed by monetary expansion, the economy will face a problem: the resulting increase in domestic credit will lead to a gradual growth in money supply and a decrease in foreign exchange reserves. Ultimately, pegging the exchange rate becomes impossible when the reserves are depleted.

We can formalize these arguments in the following model. The money demand function is as usual:

$$\frac{M(t)}{P(t)} = \alpha_0 - \alpha_1 i(t) \quad (2.1)$$

where $M$ is the nominal money demand, $P$ is the price level and $i$ is the interest rate. The money supply function is:

$$M(t) = R(t) + D(t) \quad (2.2)$$

where $M$, $R$ and $D$ are the nominal money supply, book value of foreign reserves and domestic credit respectively. Suppose that because of the excessively expansionary fiscal and monetary policies, the domestic credit expands at a positive fixed rate $\mu$ to meet the government deficits:

$$\Delta D(t) = \mu \quad (2.3)$$

Moreover, we assume PPP (Purchasing Power Parity) and UIP (Uncovered In-
terest Parity) hold,

\[ P(t) = P^*(t) \times S(t) \]  
\[ i(t) = i^*(t) + \frac{\Delta S(t)}{S(t)} \]

where \( S \) is the spot exchange rate and an asterisk (*) denotes the foreign variables, which are assumed exogenous for the small open country case.

In order to get the money market equilibrium, we substitute (2.4) and (2.5) into (2.1), and obtain

\[
M(t) = \beta S(t) - \alpha \Delta S(t) 
\]

where \( \beta = [\alpha_0 P^*(t) - \alpha_1 i^*(t) P^*(t)] \) and \( \alpha = \alpha_1 P^*(t) \). Thus, the equilibrium in the money market is given by

\[
M(t) = R(t) + D(t) = \beta S(t) - \alpha \Delta S(t) 
\]

Under a fixed exchange rate regime, \( S(t) = S, \Delta S(t) = 0 \). Therefore, the foreign reserves are equal to

\[
R(t) = \beta S - D(t) 
\]

It follows that if the central bank tries to peg its exchange rate, continuous domestic credit expansion will eventually dry up the reserves in finite time as

\[
\Delta R(t) = -\Delta D(t) = -\mu 
\]

In this case, the price level will immediately begin to rise when reserves are
exhausted. There are two reasons behind that: domestic residents may still be
dissaving, and try to reduce their holdings of domestic as well as foreign money;
and, if the government continues running a deficit as before, the nominal money
supply must begin rising.

In Krugman (1979)'s original model, this rise is reflected in the private sector's
rational expectation. If an investor correctly anticipates that the price of foreign
exchange will jump instantly by a discrete amount when the central bank sells out
all its reserves, he can earn an infinite rate of return by exchanging his domestic
currency for foreign currency just a little bit earlier than that point. If everyone
tries to do this, of course, the government's reserves will be eliminated so that all
investors can only avoid windfall capital losses, and the speculative attacks will not
lead to a discrete change in the exchange rate.

Krugman (1979) then suggests that if the foreign reserves the central bank holds
initially are larger, then the time before the crisis is due is longer.

However, Flood and Garber (1984) take another approach to find the collapse
time for the fixed-rate regime. In their model, the 'shadow floating exchange rate' is
introduced as the floating rate that would prevail if the exchange rate was not fixed
(see Figure 2.1). The upward trend in the 'shadow rate' is defined by the assumption
that the central bank is engaged in a steady, but uncontrollable issuing of money to
finance the budget deficits. As before, the limited stock of foreign reserves will be
exhausted since the central bank is assumed to use it to hold the exchange rate peg.

Suppose the post-attack exchange rate is free floating. Let $T$ be the time at
which the fixed exchange rate regime collapses, $T_+$ and $T_-$ be the instant right after
and before the attacks. Take for instance, at $t = T_+$, we have

$$M(T_+) = \beta S(T_+) - \alpha \Delta S(T_+)$$  \hspace{1cm} (2.10)

At the instant right after the attacks reserves will be depleted completely, $R(T_+) =
0$, so $M(T_+) = D(T_+)$, and $\Delta M(T_+) = \Delta D(T_+) = \mu$. If we conjecture that a solu-
tion to equation (2.10) as a first order differential equation is $S(t) = \lambda_0 + \lambda_1 M(t)$,
then $\Delta S(t) = \lambda_1 \mu$. Combining this with (2.10), we have

$$S(t) = \frac{\alpha \lambda_1 \mu}{\beta^2} + \frac{M(t)}{\beta}$$

(2.11)

By comparing 2.11 with our conjectured solution, we find $\lambda_1 = \frac{1}{\beta}$ and $\lambda_0 = \frac{\sigma \mu}{\beta^2}$. Therefore, the 'shadow exchange rate' equals

$$S(t) = \frac{\alpha \mu}{\beta^2} + \frac{M(t)}{\beta}, \quad t \geq T$$

(2.12)

If the fixed-rate regime is to last without a speculative attack until the last penny of reserves leaks away, there will be a discrete rise in the exchange rate at the point of abandonment, implying an instantaneously infinite return on speculation. Such a path, as in Krugman's original model, can not be an equilibrium. Instead, foresighted speculators, realizing that such a jump is in prospect, will sell domestic currency just before the collapse—and in doing so advance the date of the abandonment, leading speculators to sell even earlier, and so on... the result is that when the shadow exchange rate equals the fixed rate ($S(T) = S'$), there will be an abrupt speculative attack that quickly drives all the central bank's remaining reserves to zero and forces it to abandon the fixed exchange rate regime (Figure 2.1).

The remarkable finding of the classical crisis model is its surprising prediction that the exchange rate will have to be abandoned, before the central bank's foreign reserves gradually decrease to zero on their own for monetizing the deficits.

Substitute this condition and the fact that $M(t) = D(t) + \mu t$ into (2.11), we have

$$T = \frac{\beta S - D(0)}{\mu} - \frac{\alpha}{\beta} = R(0) - \frac{\alpha}{\beta}$$

(2.13)

From (2.13) we see that the larger the initial foreign reserve holdings are, the longer the fixed regime will last, same as in Krugman (1979)'s original model. Also, a decrease in credit expansion (smaller $\mu$) will put off the speculative attacks. Overall, the Krugman-Flood-Gerber model suggests that fiscal variables such as the budget...
Figure 2.1: Anatomy of a Speculative attack
deficits to GDP ratio, and financial variables such as the domestic credit to GDP ratio and international reserves are all crucial factors explaining currency crises.

In an extended model (Agénor et al. (1992)), expansionary fiscal and monetary policies lead to higher domestic demand for both tradable and non-tradable goods. The former causes a deterioration in the trade balance, while the latter causes a real appreciation of the currency. Thus, external variables such as trade and current account balances should also be considered as potential causes of currency crises. We will examine the effects of all these variables suggested by the first generation models on currency crises in Chapter 3.

The Krugman-Flood-Garber model was originally inspired by the experience of many Latin American countries in the late 1970's and 1980's. During that period, countries like Argentina, Chile and Brazil embarked on stabilization programmes, including pegging their currencies, in order to curb high inflation rates. But their failure to tackle large fiscal deficits at the same time resulted in the same scenario: the collapse of the fixed rate regimes and a massive depletion of the central banks’ foreign exchange reserves. For this reason, the first generation crisis models are also described as the ‘seignorage-driven’ models by Krugman (1996).

2.2.2 The Second-generation Crisis Models

Under intense speculative attacks, member countries of the European Monetary System (EMS) agreed in August 1993 to widen the fluctuation bands for most ERM rates from ±2.25 percent around par to ±15 percent. This shows that factors other than weak economic fundamentals and/or irresponsible credit expansions could lead to currency crises, prompting the emergence of the so-called ‘second-generation’ crisis models that emphasize the contingent nature of government’s economic policies.

Obstfeld (1986), adopting the Flood and Garber (1984) framework, demonstrates the existence of circumstances in which balance-of-payments crises might indeed be purely self-fulfilling events rather than the inevitable results of unsustainable macroeconomic policies. In the paper, he suggests that there are an infinite number
of alternative equilibria, each representing a set of public beliefs about the probability of a run on the central bank’s reserves. They are all self-fulfilling equilibria in the sense that if private agents expect a speculative attack on the fixed exchange rate, then the central bank will validate the expectation by shifting to an inflationary policy. Moreover, it is shown that the nominal interest rate can at times exceed the world rate even when the exchange rate remains fixed, in order to compensate domestic bond holders for the possible capital losses in the case of a currency devaluation.

Based on that, Obstfeld (1994) and Obstfeld (1996) develop the theory in two different ways, showing how crises (realignments) can result from the interaction between the rational private agents and the government that pursues ‘well-defined’ policy goals. Note that in his models, foreign reserves can be freely borrowed in the world capital market, so that there is no special role assigned to reserve level in generating currency crises.

In one model that is later widely cited by most economists, it is shown that devaluation could be triggered by the government’s desire to offset negative output shocks by loosening monetary policies—the perfect foresight formed by private sectors could feed into wages, lower competitiveness and raise unemployment, then therefore sap the government’s resolve to resist the realignment, which in return validates the expectation. Obstfeld (1996) draws the basic framework from the Barro and Gordon (1983) model, but assumes an open economy. The government’s loss function is given by:

$$\ell = (y - y^*)^2 + \beta \epsilon^2 + C(\epsilon)$$  \hspace{1cm} (2.14)

where $y$ is output, $y^*$ is the government’s output target, and $\epsilon \equiv e - e_{-1}$ the change in the exchange rate (the price of foreign currency). $C(\epsilon)$ is the political cost born by the government in the case of a collapses of the fixed exchange rate: any upward changes in the rate (devaluation) produce $C(\epsilon) = \overline{c}$, whereas downward (revaluation) changes cost the government $C(\epsilon) = \underline{c}$. Output is determined by the expectations-
augmented Phillips curve

\[ y = \bar{y} + \alpha(\epsilon - \bar{c}) - u \]  \hspace{1cm} (2.15)

where \( \bar{y} \) is the 'natural' output level, \( \epsilon \) is the market's expectation of \( \epsilon \) based on lagged information, and \( u \) is an i.i.d. mean-zero shock. Firstly, ignore the term \( C(\epsilon) \) in (2.14). The government chooses the currency's exchange rate \( e \) in each period to minimize its loss function after observing \( u \) (unlike private sectors), taking the expectation of exchange rate as predetermined (e.g., through nominal wages agreed in the last period). The assumption of \( y^* > \bar{y} \) implies that the government always tries to boost the output above the natural level, which results in a systematic inflation bias under the rational expectations.

Because of its sticky nature, we assume that the nominal wage is time-invariant \( c^e \). If we substitute (2.15) into (2.14), then the optimal problem for the government can be solved by differentiating (2.14) by \( \epsilon \):

\[ \epsilon = \frac{\alpha(y^* - \bar{y} + u) + \alpha^2 c^e}{\alpha^2 + \beta} \]  \hspace{1cm} (2.16)

The output is then given by substituting (2.16) back into (2.15):

\[ y = \bar{y} + \frac{\alpha^2(y^* - \bar{y}) - \beta u - \alpha \beta c^e}{\alpha^2 + \beta} \]  \hspace{1cm} (2.17)

The government's loss in a flexible exchange rate regime is

\[ \xi^{\text{FLEX}} = \frac{\beta}{\alpha^2 + \beta} (y^* - \bar{y} + u + \alpha c^e)^2 \]  \hspace{1cm} (2.18)

But if the government has no option whatsoever to change the exchange rate (e.g. the fixed rate regime), its loss is instead

\[ \xi^{\text{FIX}} = (y^* - \bar{y} + u + \alpha c^e)^2 \]  \hspace{1cm} (2.19)

by setting \( \epsilon = 0 \).
Then, let us reintroduce the assumption that the government faces a fixed cost of realignment (devaluation or revaluation) into the model. Clearly, a realignment will occur whenever the government’s loss under a fixed exchange rate exceeds its loss if it realigns its exchange rate, e.g., when $u$ is so high that $\ell^{\text{FLEX}} + \bar{c} < \ell^{\text{FIX}}$ or so low that $\ell^{\text{FLEX}} + \bar{c} < \ell^{\text{FIX}}$. Devaluation thus occurs for $u > \bar{u}$, and revaluation for $u < \bar{u}$, where

$$
\bar{u} = \frac{1}{\alpha} \sqrt{c(\alpha^2 + \beta)} - y^* + \bar{y} - \alpha \varepsilon^* \quad (2.20)
$$

$$
u = -\frac{1}{\alpha} \sqrt{c(\alpha^2 + \beta)} - y^* + \bar{y} - \alpha \varepsilon^* \quad (2.21)
$$

Suppose $u$ is uniformly distributed on $[-\mu, \mu]$. Given market’s expectation $\varepsilon^*$, the rational expectation of $\varepsilon$ in the next period is

$$
E \varepsilon = E(\varepsilon|u < \bar{u}) \Pr(u < \bar{u}) + E(\varepsilon|u > \bar{u}) \Pr(u > \bar{u}) 
$$

$$= \frac{\mu}{\alpha^2 + \beta} \left[ \left( 1 - \frac{\bar{u} - u}{2\mu} \right) (y^* - \bar{y} + \alpha \varepsilon^*) - \frac{\bar{u}^2 - u^2}{4\mu} \right] \quad (2.22)
$$

Figure 2.2 graphs Equation (2.23) together with a 45° line, which shows the equilibrium $E \varepsilon = \varepsilon^*$.

This ‘escape-clause’ arrangement allows the government to escape (change to the floating exchange rate) in some extreme situations (large negative shocks), while restraining the inflationary proclivity by imposing a realignment cost. But the trigger point at which the escape option is exercised depends on prior expectation of depreciation, which in turn depends on market perception of where the realignment trigger point lies. The element of circularity creates the potential for multiple equilibria, which is shown in Figure 2.2. A relatively low credibility of the exchange rate regime could be self-fulfilling: a high expectation of depreciation feeds into wage inflation, which creates a shortfall in competitiveness, then the resulting high unemployment is so painful that the government will be forced to devalue its currency,
which in return justifies the expectation. The process is always accompanied by a sharp rise in domestic interest rates and a loss in foreign reserves. This approach will be employed to set up a theoretical model in Chapter 4 to examine the impact of the public debt, especially its debt maturity, on the credibility of the exchange rate regime.

In an earlier paper, Obstfeld (1994) suggests that the high nominal interest rates associated with the high expectation of devaluation can force the government to devalue its currency, although the peg would be viable under another set of private expectations.

The world is again assumed to last for two periods, and the government can only levy taxes on output to balance its budgets in the second period. Subject to the inter-temporal budget constraint, the government’s objective function is minimized to yield the optimal solutions as in the former model. According to the concept of perfect foresight, the depreciation rate that the market expects must equal the rate the government finds optimal given the market’s expectation. Therefore, the intersections of the government reaction function and the interest parity curve determine

Figure 2.2: Three Possible equilibria in the expected depreciation rates
the possible equilibria.

Once again, the government faces a dynamic inconsistency problem (multiple equilibria): much as it would like to, it cannot credibly promise not to validate expectations if the bond market settles on the high-inflation equilibrium interest rate. In the model, the high interest rate induces the government to realign through its impact on the government's fiscal position.

In summary, these more sophisticated second generation models require three ingredients. First, there must be a reason why the government would like to abandon its fixed exchange rate. Second, there must be a reason why the government would like to defend the exchange rate—so that there is a tension between the two motives. Finally, in order to create the circular logic that drives a crisis, the cost of defending a fixed rate must itself increase when people expect (or at least suspect) that the rate might be abandoned. All the three elements combine to give rise to multiple equilibria—under one equilibrium, the fixed exchange rate regime is feasible, but not under another equilibrium.

It is recommended by the second generation models that the real business cycle variables, such as unemployment and real GDP growth rate, and the market's expectation variables (e.g. interest rate spreads) are important factors in triggering currency crises. A large reserve loss certainly accompanies a crisis, but is not considered as a factor that ultimately leads the government to devalue.

In another paper by Benigno and Missale (2001), a three-period stochastic version of the Barro and Gordon (1983) model is adopted to show that defence of the exchange rate in the current period would worsen the fundamentals by increasing public debt, and thus impair the ability of the authority to withstand crises in the future.

It is shown that whether the exchange rate regime gains or loses credibility after a successful defence is uncertain. On one hand, resisting a crisis sends a signal that the government is tough on inflation (devaluation) and thus enhances the expectation that the parity will be maintained in the following period—the 'signal' effect.
On the other hand, defending the parity and refraining from inflationary financing increases the debt burden, hence the likelihood of a forced future devaluation—the 'debt-burden' effect. Which effect dominates depends on the importance of the government's credibility relative to the public debt level.

In an earlier paper, Drazen and Masson (1994) suggest that the persistence of unemployment could provide another channel whereby current actions affect future policy. An extended theoretical model based on Benigno and Missale (2001) will be presented in Chapter 4 with its results demonstrated by numerical simulation using Matlab, and the two derived predictions will be tested empirically.

2.2.3 The Third-generation Crisis Models

In the most part of the 1980's and the first half of the 1990's, the newly industrializing countries in East Asia were held up as the world's most dramatic economic success story. They were characterized by exceptionally rapid rates of economic growth and human development, by relatively low inflation and by an absence of balance of payment difficulties. Then, suddenly in 1997/1998, it seemed that the success turned sour and the economies of East Asia experienced an economic reversal of such proportions that it would generally be regarded as a currency/financial crisis. The crisis saw a sharp depreciation in exchange rates, a large turnaround in capital flows from inflows to outflows, high interest rates, rising unemployment, reduced confidence in policies, and a big reduction in the rate of economic growth.

But the essential features that characterize the earlier crisis models—weak fundamentals, excessive monetary expansion, and conflicting government policy goals—did not appear to exist in these East Asian countries before the crises occurred. Also surprising is the vulnerability of some (but not all) of the East Asian economies to crisis contagion. As a result, the so-called 'Third-generation Models' were developed. Amongst them, the problems of moral hazard (international over-borrowing) and the liquidity crunch experienced by the private sector once a currency regime collapses are in particular believed to be the main causes of these currency crises.
Moral Hazard Models

The Moral Hazard Models, laid out by Krugman (1998) and McKinnon and Pill (1998), suggest that factors like moral hazard and international over-borrowing played a key role in the 1997/98 East Asia Crises, acting as main forces for bubbles in asset markets. According to Krugman (1998), renewed emphasis needs to be put on two important issues: the role of financial intermediaries (and of the moral hazard associated with such intermediaries when they are poorly regulated), and the prices of real assets such as capital and land.

In the East Asian countries, the foreign creditors always believe that their investment to local financial institute are guaranteed implicitly by the governments, as there exist strong connections between the owners of these local institutions and the East Asian governments. These local financial intermediaries that are able to raise money at safe interest rates from international financial markets, intend to lend that money at premium rates to finance speculative investments. In other words, since the local intermediaries are essentially unregulated, they have an incentive to undertake excessively risky investments. As a result, there appears to be a serious problem of moral hazard. If the supply of assets is completely inelastic—the intermediaries have their impact not on quantities but on prices, then the excessive risky lending from these intermediaries creates inflation—not of goods but of asset prices.

Krugman (1998) emphasizes that with moral hazard, crises could be self-fulfilling: if the pre-crisis asset values are in a ‘meta-stable’ state (such as the boom in property and equity markets in Malaysia and Thailand in the early 1990s), a small shock could cause a slump in asset values that will make the insolvency of intermediaries visible and force them to cease operations, which in turn would justify the drop in asset prices by leading to further deflation.

Corsetti et al. (1998) then go further to argue that because of the special relationship between the governments and the local financial institutions, the former might at last have to step in and guarantee the outstanding stock of external liabilities of the latter. To satisfy demand for solvency requirement, the governments
may possibly resort to seigniorage revenues through money creation, which therefore causes a first-generation crisis.

It is also worth noting that access to the world capital markets might actually have made the East Asian economies worse off, by allowing moral hazard in financial sectors to be translated into excess capital accumulation instead of driving up the interest rates which is what happens when the local financial institutions have to rely on fixed supplies of domestic savings.

In short, the moral hazard models consider the East Asia story as a bubble in and subsequent collapse of values in asset markets, with the crisis rather a symptom than a cause of the underlying real (in both senses of the word) malady. Therefore, the domestic credit provided by the banking sector would be an important factor in triggering currency crises in East Asia.

Illiquidity

In contemporary banking theory, one of main reasons of the existence of financial intermediaries is to provide liquidity for private agents by issuing demand deposit contracts. As Diamond and Dybvig (1983) argue, there is a co-ordination problem between the bank and depositors. If a large number of depositors attempt to exit since they individually believe that there will be a bank run, they will then impose huge pressure on the bank. As a result, the behaviours of the depositors who first require withdrawals are validated as the bank hardly has enough liquidity to pay back the requested large deposits and is finally forced to default. This vicious circle is enforced by the fact that the depositors who are most reluctant to withdraw their deposits will expect there will be no money left in the bank in the event of bank run, and thus join other depositors to hurry out. By contrast, all the depositors will be better off collectively if they stay put. Hence, banks are subject to the risk of illiquidity and a bank run can be self-fulfilling.

Change and Velasco (1998) extend the Diamond-Dybvig framework into a small open economy context, and emphasize the significant role that international illiquidi-
ity played in the recent East Asia financial crises. In a 3-period Diamond-Dybvig setup, the domestic agents discover their 'type' at $t = 1$. Each can be either an 'impatient' type who only derives utility from early consumption (at $t = 1$) or a 'patient' type who only derives utility from late consumption (at $t = 2$). It is also assumed that the agents, who also happen to be the investors, are faced with two investment choices at $t = 0$: a short-term investment in the world capital market with a low rate of return at $t = 1$, or a long-term domestic investment with a higher rate of return at $t = 2$. The long-term investment, however, is illiquid in the sense that early liquidation (at $t = 1$) will yield relatively little. Each agent is born with an endowment but can also borrow a limited amount from abroad.

If we suppose that there is no co-ordination problem in the model, then investors will act collectively as 'a commercial bank' that pools its resources (including each investor's maximum foreign borrowing and the endowment that he/she inherits at the period 0) in order to maximize the welfare of all agents by providing them with optimal liquidity solutions.

However, a financial (currency) crisis will materialize precisely if all agents lose confidence in the bank and demand immediate withdrawals. This would force the liquidation of the more profitable long-term investments. Such collective behaviour turns out to be individually optimal—it forces the bank to run out of resources and fail to meet all the claims on it. It is also shown that the collapse of the financial intermediaries can lead to a full-blown financial (currency) crisis, when combined with other problems such as weak fundamentals, moral hazard and excess foreign borrowing. Thus, factors that would affect the liquidity of banking sector (e.g. ratio of bank liquid reserves to bank assets) are regarded significant in indicating the financial fragility.

Chui et al. (2001) set up a model to explore the policy ramifications of sovereign liquidity crises. One central feature of their analysis is that they focus on the collective action problem amongst creditors, and the interaction between fundamentals and strategic behaviours. They identify the unique liquidity crisis equilibrium of
the liquidity game and are, thus able to characterize and calibrate the welfare costs of sovereign liquid crises. Some specific policy proposals are evaluated—namely, prudent debt management, information disclosure, and capital controls.

Contagion

Besides the severity of the East Asia crisis, the speed with which it spread from the Thai crash to other countries in the region also gave most commentators a big surprise, as it did in the case of Russia and Brazil in the following year. The literature has striven to explain these episodes in both theoretical and empirical works. The term ‘contagion’ is first defined systematically by Masson (1998) as having three distinctive effects: ‘monsoonal’, ‘spillovers’ and ‘jumps’. The monsoonal effect refers to a cluster of crises due to a common external cause, while spillovers relate to the interdependence amongst countries that are linked by trade and/or finance. The last category of jump or pure contagion refers to a shift in private agents’ expectations, but not to changes in the country’s macroeconomic fundamentals.

But Forbes and Rigobon (1999) suggest a slightly different taxonomy of the crisis transmission mechanism: crisis-contingent (shift-contagion) and non-crisis contingent (real linkages) contagion. For the former, a structural shift in the economy is caused by a crisis via a channel that does not exist pre-crisis. An obvious example is the so-called ‘common lender’ effect—during a crisis, if the common lenders fail to cash their claims for liquidity in one country, they will seek for it in another country, and... then in other countries. Shortly after Finland devalued the markka in the early 1990’s, the German banks that had relatively heavy exposures to Finland (as well as other European countries) were forced to re-evaluate their portfolios, creating a Euro-wide liquidity crisis. For the latter type of contagion, it is assumed that the transmission channels shown after an initial shock are just a continuation of the ‘real linkages’, which has already existed before the crisis. Scenario include Masson’s monsoonal effect, and the situation when a currency crisis in one country can trigger a beggar-thy-neighbor type of devaluation in its neighboring countries.
with close trade links. It is argued that only those countries that are affected by shift-contagion can qualify for multilateral bailouts.

One of the well-defined contagion models is that of information 'cascades', which is based on the herding model developed by Banerjee (1992). Chari and Kehoe (2000) point out in their model that asymmetrically informed investors acquire information sequentially by observing the actions of others who precede them. So if the first \( n \) investors receive bad signals about the state of the economy and sell the currency, the \( n + 1 \) investor will infer that all those investors have information favouring that action and sell the currency as well, although his private information alone would imply another action. Then such behaviours, whatever their initial causes, could be magnified through sheer imitation and turn, quite literally, into a stampede out of the currency. Calvo and Mendoza (1998) get the same results even after relaxing the assumption of sequential decision-making. They assume that investors form their decisions simultaneously instead. The idea of information externality is still applied to the uncertainty about policymaker’s commitment to a fixed exchange rate. It is shown that small rumors can trigger herding behaviours amongst investors who have few incentives to collect country-specific information. This collective action can shift the country and its neighbors from an equilibrium where the fixed exchange rate regime is feasible, to another where it is not.

The models of ‘trade spillovers’ put forward by Gerlach and Smets (1995) and Eichengreen et al. (1996a), show that an attack-induced devaluation in one country enhances its export price competitiveness, leading to increases in trade deficits (and a drastic decline in foreign reserves) in countries which it competes with in a third country. As a result, its trading partners find that their currencies become more vulnerable. Valdés (1997) develops a simple model of capital flows to show that some country-specific fundamentals (i.e. the liquidity of one country) could affect the abilities of other countries to repay their debts during a crisis. Financial spillovers can also be applied to individual agents, as the third-country investors try to liquidate their positions in other countries when facing a crisis-induced loss in one
country. Drazen (1998) studies the political pressure on one country to commit to
the fixed exchange rate arrangement, if other potential members of the union place
less weight on meeting the required conditions. It is shown that in that case, the
country is less likely to maintain the fixed exchange rate either.

2.3 The Relevant Empirical Literature

In this section, we focus on the leading indicator literature where statistical models
are set up to help rank the vulnerability of countries given a set of potential factors.

There are in general four different approaches to set up leading indicator frame-
works examining currency crises: the signalling approach, the discrete-choice ap-
proach, the structural approach, and the Markov Regime Switching approach.

2.3.1 Signalling Approach

Kaminsky et al. (1998) are amongst the first to adopt this approach to analyze
currency crises. A broader definition of crises is adopted in their framework—in
addition to devaluations, they include episodes of unsuccessful speculative attacks,
i.e., attacks that are averted without a devaluation but at the cost of a large increase
in domestic interest rates and/or a sizable loss of international reserves. Therefore,
a Market Pressure Index (MPI) for the exchange rate regime is created, which is
a weighted average of monthly percentage changes in the exchange rate and the
negative monthly percentage changes in gross international reserves, measured in
US Dollars.

Based on theoretical considerations and the availability of monthly data, 15
variables are chosen as potential factors for crises. Their warning system essentially
involves monitoring the evolution of the economic variables that tend to systemati-
cally behave differently prior to a crisis. Every time that a variable exceeds a certain
threshold value, it is interpreted as ‘signalling’ a warning that a currency crisis may
take place within the following 24 months. The ‘optimal’ set of threshold values are
chosen to minimize the noise-to-signal ratio, so as to strike a balance between the risk of having too many false signals (if a signal is issued at the slightest possibility of a crisis) and the risk of missing too many crises (if the signal is issued only when the evidence is overwhelming). Furthermore, one can identify the source(s) of the crisis from the information provided by the group of variables that are issuing signals.

Finally, they conclude that the variables that have the best track record in anticipating crises within their framework include exports, deviation of the real exchange rate from its long-term trend, ratio of broad money to gross international reserves, output, and equity prices.

Kaminsky and Reinhart (1996) also suggest that problems in the banking sector normally precede a currency crisis, and that a currency crisis in return could increase the risk of a bank run.

2.3.2 Discrete-Choice Approach

Discrete-Choice Approach is the most common method that economists adopt to establish the potential factors relating to currency crises. It allows a quantification of the probability of having a crisis, given a set of economic variables. (see Frankel and Rose (1996) and Kumar et al. (1998))

Based on pooled panel data across countries and time periods, the discrete-choice technique—Logit/Probit model is employed to analyze crisis incidents. Let \( y \) denote the crisis variable that takes a value of either 1 (if a crisis occurs) or 0 (otherwise), \( X \) be a vector of explanatory variables and \( \beta \) be a vector of parameters. Then we can obtain the probability of having a crisis as:

\[
Pr(y = 1) = f(\beta'X)
\]

where \( f(.) \) is a Logistic/Probability distribution function. If we assume a Probability distribution and a standard normal distribution with unit variance, then
This is known as the Probit model.

Alternatively, if we assume a logistic distribution, then

$$\Pr(y = 1) = \frac{1}{1 + \exp(-X'\beta)}$$  \hspace{1cm} (2.26)$$

This model is known as the Logit model. The only difference between the Probit and Logit model is that the underlying latent variable that is assumed to generate the discrete events in two models has slightly different distributions—being a little more fat-tailed in the Logit case than in the Probit case. The parameter vector $\beta$ can be estimated by Maximum Likelihood in either way.

Within this approach, the crisis definition normally differs amongst studies and are subject to arbitrary choices. The crisis variable ranges from the simple large nominal depreciation to a significant increase in the crisis index (e.g. MPI in Kaminsky et al. (1998)), while the latter incorporates the changes in exchange rate, foreign reserves and interest rates. The threshold level chosen to qualify for a crisis is arbitrary and varies from study to study, but is typically 10% or a one standard deviation movement. Furthermore, in the case of countries with high inflation rates and consequently high expectation of depreciation, as Frankel and Rose (1996) suggest, a substantial increase in the nominal depreciation or in upward movement in the index compared with the prior period is also required for a currency crisis to be identified.

The choices of sample frequency and size also vary amongst projects. Frankel and Rose (1996) choose a panel of annual data for over 100 countries from 1971 through 1992 to characterize currency crashes. Kumar et al. (1998) choose monthly data because they focus on how to 'forecast' currency crashes, and covering emerging markets only.

Despite all these differences, empirical works following the discrete-choice ap-
proach produce the most fruitful results in identifying the potential factors for currency crises. Frankel and Rose (1996) examine the variables that are most suggested by the first generation models, especially fiscal and debt variables. They find that currency crashes tend to occur when domestic credit growth is high, the real exchange rate appears to be over-valued and reserves are low. A low ratio of FDI to debt is also consistently associated with a high likelihood of a crisis. They also attempt to associate crashes with sharp recessions, but fail in at the end as the causal linkage is not clear. Moreover, surprisingly neither current account balance nor government budget deficits appear to play an important role in a typical crisis.

Jeanne (1998) studies the respective role of the fundamentals and self-fulfilling speculation in currency crises. She first presents a model of a fixed exchange rate system in which self-fulfilling speculation can arise following a bifurcation in the fundamentals. She then estimates the model in the case of French Franc in the period 1991-93. She shows that the model—in which the relationship between the fundamentals and devaluation expectation is non-linear and may give rise to multiple equilibria—gives a better account of the data than the usual linear models. In other words, she finds evidence that self-fulfilling speculation is at work.

Kumar et al. (1998) select their explanatory variables based on an a priori judgement drawing on the three generation crisis theories. Their findings confirm that the most important explanatory variables for crashes are declining reserves and exports according to the first generation models, and weakening economic activities for the second generation models. Contagion also plays an important role in triggering the occurrence of a crash, both as a regional factor and through export growth correlation. Surprisingly, they find that the capital flows variables suggested by the third generation models mainly have counter-intuitive signs, although not statistically significant.

The journey will go on in Chapter 3, but in a slightly different fashion. Variables recommended by all three generation crisis models will be selected as potential factors explaining currency crises in Latin America, Europe and East Asia. Our
aim is to see how good the three generation crisis models are in addressing the different features of currency crises in different regions respectively, by which they were originally inspired.

2.3.3 Structural Approach

While the above two approaches aim to establish the likelihood of impending currency crises, the structural approach is developed to explain the causes of currency crises in terms of what characteristics of a country make it more vulnerable to speculative attacks. Based on crisis theories, structural models are applied to case studies of the specific devaluation episodes.

Sachs et al. (1996a) (STV) analyze the financial incidents following the devaluation of Mexican peso in 1994 by adopting a simple structural model. They identify three factors that could determine whether a country is vulnerable to a financial (currency) crisis: a large appreciation of the real exchange rate, a weak banking system, and a low level of foreign exchange reserves.

The cross-sectional data of 20 emerging markets in 1994 is examined in STV. In order to measure the pressure in foreign exchange markets, an index \( (INP) \) is created as a weighted average of the depreciation rate with respect to US Dollar and the negative percentage change in foreign exchange reserves. The extent of real exchange rate misalignment \( (RER) \) is represented by the percentage changes in the average of the real exchange rate index in 1990-94 compared with the period 1986-1989. The banking sector vulnerability \( (LB) \) is measured in terms of whether or not the economy experiences a lending boom immediately before the crisis. Finally, the yardstick with which to evaluate the abundance of reserves is a broad measure of money \( (M2) \) compared with the stock of foreign exchange reserves \( (M2/R) \).

The basic framework is a statistic model regressing the (continuous) crisis index \( (IND) \) on the levels of \( RER \) and \( LB \), taking into account the strength or weakness of these fundamentals and the adequacy of foreign exchange reserves. Sachs et al. (1996a) find that countries with strong fundamentals but low reserves are not likely
to be attacked. But for countries who have weak fundamentals and low reserves, a more overvalued real exchange rate or a larger lending boom will lead to a higher crisis pressure.

Corsetti et al. (1998) adopt a similar framework to examine the East Asia crises and find that variables, such as the non-performing loans in the banking system, the current account balance, and M1 (as a proxy for the strength of fundamentals and the position of reserves), are also significant explanatory variables for currency crises.

2.3.4 Markov Regime Switching Approach

Recently, it has become increasingly popular to analyze currency crises as regime changes in a non-linear model, and in particular the evolution of a regime is assumed to follow a first-order Markov chain\(^1\). The literature is currently not extensive, but it is clear that interest in this approach is increasing rapidly.

Engel and Hakkio (1996) use the method of identifying outliers to show that the European Monetary System (EMS) exchange rates over the period 1979-1993 seem to be drawn from a mixture of distributions, one with high variance (which coincides with speculative attacks and realignments) and one with low variance (tranquil periods). They also model the probabilities of switching between states as a function of the distance of the exchange rate from the upper band, and find the negative relationship between the probabilities of staying in a volatile state and the distance of the exchange rate to the top of the band.

Jeanne and Masson (1998) propose to use the Markov-switching model to bring the second generation approach to the data. They characterize the properties of a class of second generation models, and then show that financial markets are subject to multiple equilibria that link the markets to fundamentals. They also find that what are commonly called 'sunspot equilibria' exist, where extrinsic variables (such as market sentiment) might influence the equilibrium outcomes. A liberalization of

\(^1\)See Appendix A for a detailed technical review about the Markov-switching model.
their model gives a Markov-Switching model the probability of devaluation, where the switch between regimes corresponds jumps amongst different equilibria. Finally, they estimate the Markov-Switching model for the French franc over the period 1987-93, and find that a model allowing for so-called sunspots performs better than a purely fundamental-based model.

Hsieh (1994) employs an AR(4) version of Hamilton’s regime switching model to identify speculative attacks on EMS currencies in the period 1979-93. He models the behaviour of exchange rates, reserves, and interest rate differentials of EMS countries as time series subject to discrete shifts in regimes. But he assumes that the transition probability is constant over time. He then compare these episodes with those captured by indices of speculative pressure in Eichengreen et al. (1996b).

Similar to Hsieh (1994), Peria (1998) models the components of the Reinhart and Rogoff (2004) index as autoregressive time series subject to shifts in regimes. While Hsieh (1994) only assumes the mean and variance to be different across regimes, Peria (1998) also allows the coefficients of the autoregressive terms to be potentially different between regimes. Instead of fitting the switching model to each of exchange rates, reserves, and interest rate differentials, Peria (1998) estimates a regime switching VAR model that allows him to analyze the attacks associated with the combined behaviours of three variables. Furthermore, Peria (1998) allows the transition probability to and from the more volatile regime to be time-varying, so that he can explicitly examine the behaviours of determinants of crises. Finally, Peria (1998) also includes the market expectation in his analysis in causing speculative attacks.

We will follow the latest trend in empirical works to employ the non-linear regime changing models, such as Markov-switching model and the SETAR (Self Exciting Threshold Autoregression) model to examine currency crises. While the Markov-switching model assumes that the evolution of the state follows a first-order Markov process, the state is supposed to change to the period of turmoil once the variable exceeds its threshold in the SETAR model.
2.3.5 Contagion

In general, there are two types of approaches in the empirical works to studying the contagion effect. The first approach is to examine the impact of contagion on the country's vulnerability to speculative attacks, and establish the probability of having a crisis given the contagion effect. The second approach takes one step back, and focuses on the tests of whether contagion actually exists by examining the cross-market co-movements pre and post-crisis.

In the discrete-choice approach discussed above, many economists include a proxy for contagion in their models and have found that it could affect the likelihood of a future crisis. The contagion variables are either chosen by statistical means or designed to capture regional linkages.

Glick and Rose (1999) estimate a binary Probit model across countries via maximum likelihood. The cross-sectional data is from five different episodes of important and also widespread currency instability. The default measure of trade linkages is given by the degree to which the 'ground zero' country competes with other countries in the foreign (third country) export markets. It is shown in their results that currency crises would affect a cluster of countries tied together by international trade. By way of contrast, macroeconomic and financial influences are not closely associated with the cross-country incidences of speculative attacks.

On the other hand, as noted above, some economists step back and ask whether contagion really exists or not in the world. The conventional definition of a contagion is a significant increase in cross-market linkages after a shock to one country (or group of countries). In other words, it is contagion only if the correlation between two markets increases significantly during crisis periods compared with periods of stability. If the cross-market comovement does not increase significantly after the shock, then any continued level of market correlation only suggests the interdependence between the two economies, which exists in all states of the world.

The most widely-used approach to test for contagion is examining the cross-market correlation coefficient. Economists measure the correlation in assets, equity
returns and exchange rates between two markets during a tranquil period, and then test for a significant increase after a shock. Calvo and Reinhart (1996) adopt this method to test for contagion in stock prices and brady bonds after the 1994 Mexican peso crisis. They find that cross-market correlations increase for many emerging markets during the crisis. Baig and Goldfajn (1998) also find statistically significant increases in the correlations of asset returns during the East Asia crises, and therefore evidence of contagion.

However, Forbes and Rigobon (2002) argue that the conventional tests based on the conditional correlation coefficient, are biased and inaccurate due to heteroscedasticity. During crises when markets are more volatile, estimators of the cross-market correlation coefficients tend to increase and be biased upward. When tests do not adjust for this bias, they traditionally find evidence of contagion. In Chapter 5, we will follow Forbes and Rigobon (2002)'s approach to base our test for contagion upon the unconditional cross-market correlation coefficient for 1997/98 East Asia Crises. A non-linear Markov-Switching VAR framework will also be adopted for the first time to give a better description of the crisis transmission mechanism.

2.4 Summary

The frequency of currency crises has increased significantly in the last 30 years. Correspondingly, the crises literature has grown and evolved to offer different explanations for these apparently ever-changing phenomena. The initial approach of 'bad' fundamentals (e.g. Krugman (1979)), indicating that crises could be avoided with sound fiscal and monetary policies, is followed by game-theoretic models with multiple equilibria (e.g. Obstfeld (1996)). The latter predict that the conflicting goals of economic policies when combined with investors' perfect perceptions can lead to a self-fulfilled crisis even when the fundamentals are good.

More recently, as a result of the failure of existing theories to predict the East Asia crises, third generation models are developed where the problems of moral hazard and the liquidity crunch experienced by the private sector once the currency regime
collapsed are believed to be the main causes. Each generation crisis models have their own recommendations regarding what economic factors play an important role in triggering currency crises. Therefore, one would expect that the three generation crisis models are best at identifying the potential causes of the different currency crises that initially inspired them. We will look at this issue in three groups of countries—Latin America, Europe and East Asia—in the following chapter.

Similarly, the empirical works also develop drastically to account for the latest crisis episodes. The signalling approach evaluates the usefulness of a number of different variables in predicting a potential crisis. These variables are examined to determine certain threshold levels, which once crossed would 'signal' an impending financial crisis within a pre-specified period of time. The discrete-choice approach uses the Logit/Probit model to estimate the probability of having a currency crisis, given a set of potential factors that are selected upon the recommendations of relevant crisis theories. The structural models are usually built around a particular crisis case, and hence are possibly more suitable for monitoring/early warning purposes.

The fourth approach that is increasingly used in crisis literature is the Markov-switching model. The main advantage of the non-linear changing regime (state) models is that they allow for a continuous dependent variable (compared with the 0-1 binary dependent variable in the Logit/Probit model), and avoid the arbitrary choice of a threshold level to qualify a crisis. Moreover, the non-linear nature of the models is also appealing.

In summary, this chapter surveys the theoretical models and empirical evidence on currency crises, which will provide us with the necessary background for the following studies, and also offer an invaluable reference during our research.
CHAPTER 3

Currency Crises in Latin America, Europe and East Asia

3.1 Introduction

The frequency of currency crises has increased dramatically in the last 30 years (see Bordo et al. (1998)), and the scale and impact of these crises has inspired the crisis literature to grow fast to offer various explanations for these apparently ever-changing phenomena. Are currency crises across the world all the results of similar policy mistakes? Or are they instead the results of myriad unfortunate shocks? The answer is that according to the so-called three generation crisis theoretical models, different factors cause currency crises in different regions.

The initial approach of ‘bad’ fundamentals (e.g. Krugman (1979) and Flood and Garber (1984)) indicating that crises result from expansionary fiscal and/or monetary policies, followed the crisis episodes experienced by many Latin American countries during the late 1970's and 1980's. The Exchange Rate Mechanism (ERM) crashes in 1992/93 when most members of the European Monetary System (EMS) were forced to devalue their currency outside the original fluctuation bands, provided
initiatives for the second generation models (e.g. Obstfeld (1996)). The second
generation models suggest that conflicting goals of government’s policies combined
with investors’ rational expectations can create multiple equilibria, where a crisis
is triggered by a shift in equilibrium even when the fundamentals are sound. More
recently, the third generation models evolve to identify the factors that drove the
East Asian countries, once regarded as the world’s most dramatic economic success,
to the turmoil of 1997/98. They attempt to explain the crises in terms of moral
hazard and bank liquidity crunch problem. Another striking feature of the East
Asia crises is that the crisis originating in one country could quickly spread to other
countries in the region, which is also studied in the context of contagion. Therefore,
it is interesting to see whether the three generation crisis theories can respectively
address the different features of currency crises in Latin America, Europe, and East
Asia where they were originally inspired, as they are supposed to do.

This chapter examines the country experiences in the three regions for the last 30
years. The country-specific Market Pressure Index (MPI) is constructed to gauge the
vulnerability of the exchange rate regime to speculative attacks. These include not
only the actual devaluations, but also the ‘failed’ attacks that are successfully fended
off by selling out reserves and/or raising interest rates. As opposed to imposing an
arbitrary threshold a priori (i.e. 10% or a one standard deviation move in MPI)
as many previous research works did, we employ both the SETAR (Self Exciting
Threshold Autoregression) model and Markov-switching model to identify the crisis
incidents for each country. In this way, MPI is allowed to reveal by itself what move
represents abnormal behaviour, and therefore signals a crisis.

We examine a wide range of selected economic variables as potential factors for
the crashes, based on the recommendations made by the three generation crisis mod-
els. The Limited Dependent Variable Model—Logit Model is adopted to link the
binary crisis variable to the explanatory variables. We obtain our final estimated
results by dropping variables whose coefficients are not statistically significant at
the conventional 10% level. It is shown that to some extent, the three generation
models successfully explain the different causes of currency crises in different regions: a rapidly expanding stock of credit and large deficits in current account play an important role in the Latin American crises; the real economic cycle is highly associated with the European turmoils; and over-borrowing and lack of export competitiveness are the fundamental causes of the crises in East Asia. But the whole picture is far from perfect. In our regressions, some variables are found not to have the predicted signs, and some appear not statistically significant in explaining crises as they should do.

The chapter is organized as follows. The following section briefly surveys the relevant theoretical models and empirical works on currency crises. We then look at the interested variables and data set in section 3. Section 4 discusses the way to establish Market Pressure Index (MPI). The crisis episodes are defined by both the SETAR and Markov-switching model in section 5. Section 6 gives the estimated results of the Logit model of the effects of selected variables on crises. Finally, we conclude.

### 3.2 Related Literature

#### 3.2.1 Three-generation Theoretical Models

The initial approach is first put forward by (Krugman (1979)) and then modified by (Flood and Garber (1984)). In its simplest form, the model assumes that under a pegged exchange rate system, the central bank has to intervene to prop up the fixed rate, whenever investors rebalance their portfolios by buying/selling domestic assets. However, if the government runs expansionary fiscal and/or monetary policies, this is not consistent with the fixed rate and a sudden speculative attack will eventually exhaust the remaining foreign reserves. The Krugman-Flood-Garber model fits the crisis episodes in Latin American economies during the late 1970's and 1980's. At that time, most Latin American countries failed to adopt robust

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1See Chapter 2 Literature Review for more details.
domestic policies when they were pursuing low inflation rates by using pegs as an anchor.

In 1992/93, the attacks on the ERM suggested that factors other than weak economic fundamentals and irresponsible credit expansion could lead to currency crises. This prompted the emergence of the second generation models that emphasize the contingent nature of government policies. Obstfeld (1994) and Obstfeld (1996) have developed a Barro and Gordon (1983) type model in which conflicting goals of economic policies when combined with private agents' perfect foresight, can drive the exchange rate off the peg even when the fundamentals are sound. A circularity is created that leads to the possibility of multiple equilibria—under one equilibrium, the fixed rate is viable, but not under another equilibrium. The crisis could be self-fulfilling as a sudden worsening of expectations can drive shifts in equilibria, which in return validate agents' expectations.

The East Asia crises took many observers by surprise. The essential features that characterize earlier theoretical models—weak fundamentals, excessive monetary expansion and conflicting government policy goals—were not apparent in these countries before the crises. As a result, a whole 'new' generation of models were developed—the third generation models. Amongst them, Krugman (1998) and McKinnon and Pill (1998) suggest that the strong connections between the owners of the local financial institutions and the East Asian governments led the creditors to believe that there were implicit government guarantee for their lendings. There then appears to be a serious problem of moral hazard (international over-borrowing), which is the main force bringing about bubbles in pre-crisis booming asset markets. Change and Velasco (1998) show that a financial (currency) crisis can materialize precisely if many agents lose their confidence in the banks and pull their money out immediately, since their actions will force liquidation of the more profitable long-term investments.
3.2.2 Relevant Empirical Works

The debate on the causes of currency crises has become more fierce than ever, as more economists develop various methods to test the theoretical models.

Frankel and Rose (1996) adopt a Logit model on a panel of annual data for over 100 developing countries from 1971 through 1992 to characterize currency crises. They define a currency crisis as a large depreciation of the nominal exchange rate. They examine the variables mostly suggested by the first generation models, especially budget and debt variables. They find currency crashes tend to occur when the FDI inflows dry up, the domestic credit growth is high, the real exchange rate shows over-valuation, and reserves are low. They also tend to associate crises with recessions, but fail in the end as the causal linkage between the two is not clear. Moreover, surprisingly neither current account nor government budget deficits appear to play an important role in a typical crisis.

Jeanne (1998) studies the respective role of the fundamentals and self-fulfilling speculation in 1992/93 ERM crisis. She first presents a model of a fixed exchange rate system, in which self-fulfilling speculation can arise following a bifurcation in the fundamentals. She then estimates the model in the case of the French Franc in the period 1991-93. She shows that the model, where the relationship between the fundamentals and devaluation expectation is non-linear and may give rise to multiple equilibria, gives a better account of the data than the usual linear models. In other words, she finds some evidence that self-fulfilling speculation is at work.

Kumar et al. (1998) also estimate a Logit model on pooled data for emerging markets from January 1985 through October 1999. The explanatory variables are selected on the basis of an a priori judgement drawn on the three generation crisis theories. Their findings confirm that the most important factors explaining crises are declining reserves and exports according to the first generation models, and weakening real activity for the second generation models. Contagion also plays an important role in the crisis episodes. But surprisingly, they find that the capital flows variables suggested by the third generation models mainly have counter-intuitive
Currency Crises in Latin America, Europe and East Asia

signs, although not statistically significant.

Kamin et al. (2001) use annual data on 26 emerging markets from 1981-1999 to investigate the relative influence of domestic and external factors in bringing about currency crises. They estimate a standard Probit model using real exchange rate changes as measurement of speculative pressure in foreign exchange markets. They find that domestic factors are the primary determinants of vulnerability of the economies to crisis, but it is external factors that often push the stumbling countries over the edge.

We will carry on this task in this chapter by examining the abilities of the three generation crisis models to identify the causes of currency crises by which the models were originally inspired. Moreover, our study tries to make some progress in the area where until now few studies have been conducted; namely to simultaneously examine the different characteristics of currency crises occurring in three major areas—Latin America, Europe and East Asia.

3.3 The Variables and Data Set

3.3.1 The Variables of Interest

Firstly, three variables are needed to establish the country-specific Market Pressure Index (MPI): the nominal exchange rate, the interest rate and international reserves. Based on MPI, a 1-0 crisis variable is created by employing both the SETAR and Markov-switching model. The construction of these variables will be discussed in more detail in the following section.

Nine economic variables are then selected as potential factors explaining currency crises, on the basis of the three generation theoretical models.

1. Variables suggested by the first generation models

As discussed above, the initial approach by Krugman (1979) and Flood and Garber (1984) suggests that several factors could make the economy particularly vulnerable to speculative attacks: namely expansionary fiscal and/or monetary policies,
Currency Crises in Latin America, Europe and East Asia

Piling up of credit stock, current account deficits, and loss in foreign reserves.

Therefore, we include in our model the following variables: (1) Overall Budget Balance, including grants (% of GDP); (2) Domestic Credit provided by Banking Sector (% of GDP); (3) Current Account Balance (% of GDP); (4) International Reserves (months of imports).

2. Variables suggested by the second generation models

The 1992/93 ERM crises prompted the emergence of the second generation crisis models (Obstfeld (1994), Obstfeld (1996)). They emphasize that the conflicting goals of government policies combined with agents' rational expectations could give rise to multiple equilibria, where crises are self-fulfilling. Therefore, factors such as high unemployment or economic recession play a crucial role in triggering currency crises. Here (5) Unemployment and (6) Real GDP growth (annual %) are included in our measurement of the real business cycle.

The assumption of perfect foresight is also important for the second generation models. It is assumed that the depreciation rate the market expects must equal the rate the government finds optimal given market expectations. (7) Market Interest Rate Differential relative to the anchor country (USA for the Latin American and East Asian group, and Germany for European countries) is suggested as a measure of the market expectation of realignment.

3. Variables suggested by the third generation models

Krugman (1998)'s Moral Hazard model also suggests that (2) Domestic Credit provided by Banking Sector (% of GDP) plays an important role in 1997/98 East Asia Financial/Currency Crises, acting rather as a main force in the creation of bubbles in asset markets.

Change and Velasco (1998) apply the Diamond and Dybvig (1983) framework to currency crises and produce the illiquidity model. In their model, the liquidity crunch is suggested to be the main cause of the crises, thus (8) Bank Liquid Reserves to Bank Assets ratio is also examined in our model.

4. Foreign Variables
It is also critical to look at not only individual country’s internal conditions, but the external (global) environment as well. External variables include world economic activities, especially shocks from major markets. We include (9) Foreign Real GDP growth (USA for Latin America and East Asia, Germany for Europe) as a proxy for external demands.

3.3.2 The Data Set

The major part of our data set is extracted from the on-line World Bank’s World Development Indicators (WDI). The data set consists of annual observations from 1970 through 2001 for Argentina, Brazil, Chile, Colombia, Mexico and Peru in the Latin American group, for Belgium, France, Ireland, Italy, Spain and United Kingdom in the European group, and for Indonesia, Korea, Malaysia, Philippines, Singapore and Thailand in the East Asian group as well as Germany and United States. We also use the IMF’s International Financial Statistics (IFS) CD 2004 and ILO (International Labour Organization) website to fill in the missing observations in our data set.  

The quarterly data of the exchange rate, interest rate and reserves to construct the country-specific MPI, is extracted from the IMF’s International Financial Statistics (IFS) CD 2004.

3.4 Market Pressure Index (MPI)

Currency crises can be characterized in many different ways. One approach is to define a ‘currency crisis’ as a nominal depreciation of the domestic currency beyond a certain level. However, economists argue that it is more appropriate to incorporate

\footnote{Effort has been made in order to collect more frequent data (e.g., quarterly), but failed at the end due to unavailability of some variables in some countries. As the purpose of this chapter is to compare the stylized features of currency crises in three regions, it is important to base our analysis on a consistent data set. But one should always bear in mind that movement patterns of some macroeconomic variables change dramatically during months around crises, when interpreting the results of this chapter.}
Currency Crises in Latin America, Europe and East Asia

the behaviour of reserves and interest rates as well, so that we can take into account the successful defences at a cost of sizable loss of reserves and/or high interest rates. Now, the latter has become the standard form to gauge the vulnerability of a country’s exchange rate regime to speculative attacks (see Chui (2000)).

Here, we follow the commonly accepted practice to include in our definition not only ‘successful’ speculative attacks that result in a devaluation, but also ‘failed’ attacks that are warded off by selling off international reserves and/or raising interest rates. This approach adopts the idea of unsuccessful speculative attacks by searching for a sudden fall in reserves and/or increase in interest rate. Thus, the Market Pressure Index is a linear combination of the changes in a country’s exchange rate, interest rates and international reserves:

\[ MPI = w_1 \Delta e + w_2 \Delta i - w_3 \Delta r \]  

(3.1)

where \(e\) is the log of nominal exchange rate against the anchor country (USA in Latin America and East Asia groups, Germany in Europe group); \(i\) is the domestic market interest rate; and \(r\) is the log of international reserves. We multiply all the

\[ \delta_t = E_t[\Delta e_t]/\Delta t = E_t[\Delta x_t]/\Delta t + E_t[\Delta c_t]/\Delta t \]

(\(*\) )

where \(\delta_t\) denotes the differential between the domestic interest rate (for a given maturity \(\Delta t\)) and the base country’s (Deutsche Mark in this case) interest rate. \(E_t\) is the expectation operator, and \(\Delta e_t = e_{t+\Delta t} - e_t\) where \(e_t\) is the spot price of a Deutsche Mark (DM) in domestic currency at time \(t\). Then, the expected rate of change of the central parity \(E_t[\Delta e_t]/\Delta t\) is derived to measure the expected rate of realignment, after subtracting deviation of the spot rate from \(c_t\).

But this method does not incorporate movements in either nominal exchange rate or international reserves, which are predominant phenomena in almost all currency crises. It is also not suitable for the case of Latin America or East Asia where most countries pegged their currencies to US Dollar before crises (in other words, \(E_t[\Delta x_t]/\Delta t = 0\) and the expected nominal depreciation equals expected rate of realignment—no need to adjust for expected exchange rate drift). Moreover, the first equality in (\(*\) ) implies Uncovered Interest Parity (UIP) that is seriously questioned in floating exchange rate regimes (e.g., Froot and Thaler (1990)).

Market Pressure Index, On the other hand, does provide us with a universal measurement of (successful and unsuccessful) severe speculative attacks in exchange markets in Latin America, Europe and East Asia, taking into account changes in exchange rate, interest rates and international reserves. After all, a raw interest differential \(\delta_t\) (an element in MPI) is highly positively correlated with \(E_t[\Delta e_t]/\Delta t\) and therefore a good approximation of it (Rose and Svensson (1994)).
three components by 100. \( \Delta \) denotes the first difference. The minus sign in the equation ensures that a reduction in foreign reserves translates to an increase in the value of the MPI.

Since the volatilities of the exchange rate, reserves, and interest rates are very different, an un-weighted index would be the index that is dominated by the most volatile component(s). Following Kaminsky et al. (1998) and Sachs et al. (1996b)' approach to apply intuitive volatility smoothing weights, we weigh the three components according to their relative precisions so that the resulting components have equal variances. The relative precision is defined as the inverse of each series' standard deviation (s.d.)—the precision of that series—over the sum of the precisions of all three series for the sample period. The formula to calculate the weights is given by:

\[
  w_t = \frac{\frac{1}{s.d.(\Delta r)}}{\frac{1}{s.d.(\Delta r)} + \frac{1}{s.d.(\Delta t)} + \frac{1}{s.d.(\Delta e)}}
\]  

(3.2)

We assume that there is no role for intervention by foreign authorities because, as most economists argue, intervention by foreign countries accounts for only a small proportion of total intervention during currency crises. Note also that, we use the quarterly data rather than annual data to construct the MPI, as the annual data won't provide enough observations that the SETAR model and Markov-switching model use to define currency crises for each country.

The descriptive statistics of the MPis in Latin American, European and East Asian countries are shown in Table 3.1, 3.2 and 3.3 respectively.

It is shown that in general, the countries in Latin America have the largest means of MPis, while the means for MPis in European countries are the smallest. The standard deviations of MPis for some Latin American countries are more than ten times as high as those for European or East Asian countries. This reflects the fact that compared with the other two regions, Latin America has experienced the most volatile periods in the last 30 years. It is also noticed that Latin American countries have most of the highest levels of MPis in our sample, e.g. 310.93 in Argentina in
Table 3.1: Descriptive Statistics for MPIs in Latin American Countries

<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Mean</td>
<td>7.29</td>
<td>1.18</td>
<td>-1.22</td>
</tr>
<tr>
<td>Median</td>
<td>0.93</td>
<td>-2.96</td>
<td>0.25</td>
</tr>
<tr>
<td>Maximum</td>
<td>310.93</td>
<td>158.42</td>
<td>28.90</td>
</tr>
<tr>
<td>Minimum</td>
<td>-260.41</td>
<td>-90.40</td>
<td>-64.14</td>
</tr>
<tr>
<td>Std. Dev.</td>
<td>55.16</td>
<td>28.74</td>
<td>10.04</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Mean</td>
<td>0.39</td>
<td>-0.59</td>
<td>-1.32</td>
</tr>
<tr>
<td>Median</td>
<td>-0.12</td>
<td>0.11</td>
<td>-1.84</td>
</tr>
<tr>
<td>Maximum</td>
<td>28.94</td>
<td>105.98</td>
<td>384.77</td>
</tr>
<tr>
<td>Minimum</td>
<td>-7.42</td>
<td>-72.03</td>
<td>-343.44</td>
</tr>
<tr>
<td>Std. Dev.</td>
<td>4.76</td>
<td>17.54</td>
<td>53.41</td>
</tr>
</tbody>
</table>

a Data for Exchange rate, interest rate and reserves used to construct the MPI is quarterly.
b Period in brackets is the period during which the data is available.

1989Q2.

### 3.5 Identifying Currency Crisis Episodes

A currency crisis is commonly defined as more than a certain percentage or a number of standard deviation move in the index over a period. However, choice of the threshold is arbitrary. This causes problems. A 10% or a one standard deviation move in MPI may be a common phenomenon in one country or country group, but it may be extremely unusual in another. As shown in Table 3.1 and 3.2, the Standard Deviation of MPI in Argentina is 55.16, much higher than 1.72, that for France. As a result, ‘fitting one size to all’ results in characterizing normal moves as crises in some places, and crises as normal in others.

As discussed in Chapter 2 Literature Review, we follow the latest trend in the empirical works to instead adopt the non-linear statistic models to identify crises. Based on the MPI established above, two econometric techniques—the SETAR (Self
Table 3.2: Descriptive Statistics for MPis in European Countries

<table>
<thead>
<tr>
<th>Country</th>
<th>Mean</th>
<th>Median</th>
<th>Maximum</th>
<th>Minimum</th>
<th>Std. Dev.</th>
</tr>
</thead>
<tbody>
<tr>
<td>Belgium (1970Q2-2001Q4)</td>
<td>0.15</td>
<td>-0.29</td>
<td>-0.33</td>
<td>-0.38</td>
<td>16.04</td>
</tr>
<tr>
<td>France (1970Q2-2001Q4)</td>
<td>-0.29</td>
<td>-0.38</td>
<td>-0.30</td>
<td>-0.30</td>
<td>6.38</td>
</tr>
<tr>
<td>Ireland (1973Q2-2001Q4)</td>
<td>-0.33</td>
<td>-0.30</td>
<td>-0.30</td>
<td>-0.30</td>
<td>6.38</td>
</tr>
<tr>
<td>Italy (1971Q2-2001Q4)</td>
<td>0.004</td>
<td>-0.04</td>
<td>-0.15</td>
<td>-0.15</td>
<td>5.65</td>
</tr>
<tr>
<td>Spain (1974Q2-2001Q4)</td>
<td>-0.30</td>
<td>-0.007</td>
<td>-0.15</td>
<td>-0.15</td>
<td>5.65</td>
</tr>
<tr>
<td>United Kingdom (1972Q2-2001Q4)</td>
<td>6.45</td>
<td>1.72</td>
<td>2.04</td>
<td>2.04</td>
<td></td>
</tr>
</tbody>
</table>

Exciting Threshold Autoregression) model and Markov-switching model (see Hamilton (1989) and Hamilton (1994))—are employed to create a 0 – 1 crisis variable for each country. A value of 1 indicates a crisis while a value of 0 indicates a tranquil state.

Both models have been popularly used in the business cycle literature to identify recessions, and have also been increasingly adopted to analyze currency crises. The major advantage of using non-linear state models is that we can avoid the arbitrary choice of the threshold value by allowing the data itself to reveal what move represents abnormal behaviour and therefore signal a crisis. Moreover, the non-linear nature of both models is also appealing for time series such as the MPI.

3.5.1 The SETAR Model

The SETAR model (see Potter (1995)) can be seen as a special case of a non-linear model with a single index restriction. More specifically, the model is defined by
Table 3.3: Descriptive Statistics for MPIs in East Asian Countries

<table>
<thead>
<tr>
<th>Country</th>
<th>Mean</th>
<th>Median</th>
<th>Maximum</th>
<th>Minimum</th>
<th>Std. Dev</th>
</tr>
</thead>
<tbody>
<tr>
<td>Indonesia</td>
<td>0.36</td>
<td>-0.20</td>
<td>41.66</td>
<td>-23.43</td>
<td>6.54</td>
</tr>
<tr>
<td>Korea</td>
<td>0.19</td>
<td>-0.30</td>
<td>11.72</td>
<td>-9.51</td>
<td>2.72</td>
</tr>
<tr>
<td>Malaysia</td>
<td>-0.22</td>
<td>-0.20</td>
<td>4.84</td>
<td>11.72</td>
<td>1.46</td>
</tr>
<tr>
<td>Philippines</td>
<td>0.34</td>
<td>0.42</td>
<td>12.83</td>
<td>-9.02</td>
<td>3.83</td>
</tr>
<tr>
<td>Singapore</td>
<td>-0.60</td>
<td>-0.51</td>
<td>4.36</td>
<td>-4.68</td>
<td>1.31</td>
</tr>
<tr>
<td>Thailand</td>
<td>-0.21</td>
<td>-0.65</td>
<td>13.56</td>
<td>-7.37</td>
<td>2.72</td>
</tr>
</tbody>
</table>

Note: Data for Exchange rate, interest rate and reserves used to construct the MPI is quarterly.

Period in brackets is the period during which the data is available.

\[ MPI_t = \begin{cases} 
\alpha_{01} + \alpha_{11}(L)MPI_t + \mu_t, & MPI_{t-d} \leq c \\
\alpha_{02} + \alpha_{12}(L)MPI_t + \mu_t, & MPI_{t-d} > c 
\end{cases} \]

(3.3)

where \( \alpha_{1i}(L), (i = 1, 2) \) is a vector of associated polynomials in the lag operator. In the model, there are two regimes and the process generating \( MPI_t \) switches between them, depending on whether \( MPI_{t-d} \) is greater or less than the threshold value \( c \), where \( d \) represents the delay in response. The current period value of MPI is chosen as the transition variable in our regression \((d = 0)\). The model in each regime is a linear autoregressive process, while it is non-linear overall caused by the endogenous switching between the two regimes. In other words, the SETAR model involves using the threshold and lagged MPI to explain the MPI itself.

The SETAR model can be rewritten in the alternative form

\[ y_t = (\alpha_{01} + \alpha_{11}(L)y_t)(1 - I(y_{t-d} > c)) + (\alpha_{02} + \alpha_{12}(L)y_t)I(y_{t-d} > c) + \mu_t \]

(3.4)
where \( I(a) \) is an indicator function with \( I(a) = 1 \) when the condition \( a \) is true and \( I(a) = 0 \) when \( a \) is false. The SETAR procedure is designed to estimate a threshold level of the MPI, which once crossed, signals a regime change in the MPI. The value of the threshold, \( c \), is chosen to deliver the lowest standard error in the maximum likelihood estimation.

### 3.5.2 The Markov-switching Model

The Hamilton’s Markov-switching model \(^4\) (see Hamilton (1989) and Hamilton (1994)) allows us to assume that the time series (MPI) is drawn from the different distributions conditional on the state-contingent parameter sets. But since we cannot observe the shifts between regimes (normally two—one tranquil state and the other turmoil state) directly, we must instead draw a probabilistic inference about whether and when changes may occur using the observed behaviour of the MPIs. The Markov-switching model is therefore drawn in the following formula:

\[
MPI_t = v(s_t) + \phi(L)MPI_t + \eta_t
\]  

(3.5)

where \( \phi(L) \) is a vector of associated polynomials in the lag operator. In the model, the intercept term \( v(s_t) \) is assumed to switch between the two states:

\[
v(s_t) = \begin{cases} 
   v_1, & \text{if } s_t = 1 \ ('stable') \\
   v_2, & \text{if } s_t = 2 \ ('crisis')
\end{cases}
\]  

(3.6)

and \( \eta_t \sim \text{NID}(0, \Sigma(s_t)) \), where \( \Sigma(s_t) \) is also assumed to be different between the two regimes. Equation (3.5) is called the ‘measurement equation’ in the Hamilton’s state space model. The state \( s_t \) is presumed to follow a first-order Markov process so that we can write the ‘transition equation’ as

\[
Pr(s_t = i|s_{t-1} = j) = p_{ij}
\]  

(3.7)

\(^4\)See Appendix A for a detailed statistical review of the Markov-switching model.
where \( \xi_{t-1} \) is a vector representing all the information available at time \( t - 1 \), which includes lagged values of \( MPI_t \). As shown in Appendix A, the Hamilton filter (Hamilton (1989)) provides a nonlinear algorithm for generating inference about a discrete-valued unobserved state vector. As a by-product of that algorithm, the log likelihood function for the observed data evaluated at the estimates of parameters can be calculated. Then, the EM method is adopted to obtain maximum likelihood estimators for the Markov-switching model. Moreover, the smoothed estimators of the state vector based on all the information provided by the sample can be calculated by the backward recursion algorithm (Kim (1994)), in order to show the ‘best guess’ of which state the economy is in at period \( t \).

### 3.5.3 Defining Currency Crises by using the SETAR and Markov-switching Model

As opposed to imposing a threshold \textit{a priori} (i.e., a 10% or a one standard deviation move in the MPI), we employ the SETAR and Markov-switching Model to allow the data itself to reveal what move represents abnormal behaviour and therefore signal a crisis. The crisis periods that are identified by the SETAR and Markov-switching model for Latin American, European and East Asian countries are shown in Table 3.4, 3.5 and 3.6 respectively.

As the MPI used in the models is quarterly data, we construct the annual crisis variable in the way that a year of turmoil is defined if there is a crisis (or crises) in any quarter(s) in that year. We adopt a conservative measurement to only grant the ‘crisis’ status to a year when both the SETAR and Markov-switching model identify crises in that year. Also note that no years shown in the table mean that the SETAR model performs poorly to produce too many unnecessary crises. Because of the single index restriction in the SETAR model, this normally happens when there

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5 The estimated results of both models on countries in three regions are reported in Table 3.9–Table 3.14. Figure 4.22–Figure 3.39 show both the filtered and smoothed probabilities of being in volatile state for all countries respectively.
is too frequent fluctuation in the country's MPI, as, for example, in some European countries that adopt floating exchange rates. In this case, we ignore the SETAR model and regard the Markov-switching model as the only way to identify crises.

<table>
<thead>
<tr>
<th>Countries</th>
<th>Currency Crises identified by SETAR Model a</th>
<th>Currency Crises defined by Markov-switching Model b</th>
<th>Currency Crises defined by Both Models c</th>
</tr>
</thead>
</table>

a The estimated Non-linear SETAR model has a regime-dependent intercept and slope parameters.
b The estimated non-linear Markov-switching model has a regime-dependent intercept and allows for heteroscedasticity.
c A currency crisis is considered to occur in a year only if both the SETAR and Markov-switching model define a turmoil quarter (or quarters) during that year.
d Period in parenthesis indicates the period during which the data is available.

The constructed binary crisis variables for the three country groups have the descriptive statistics shown in Table 3.7. It is shown that there are 18 crises in Latin America, 35 crises in Europe and 37 crises in East Asia from 1970 through 2001. The small number of currency crises that Latin American countries experienced in last 30 years does not necessarily mean that Latin America has suffered less than other regions. In contrast, however, Latin American countries experienced the most serious turmoil that ever happened, as indicated in Table 3.1. On the other hand, most of the speculative attacks on European and East Asian currencies picked up by the models, are not so severe and are mostly fended off by the authorities.
Table 3.5: Identification of Currency Crises in Europe

<table>
<thead>
<tr>
<th>Countries</th>
<th>Currency Crises defined by SETAR Model a</th>
<th>Currency Crises defined by Markov-switching Model b</th>
<th>Currency Crises defined by Both Models c</th>
</tr>
</thead>
</table>

a The estimated Non-linear SETAR model has a regime-dependent intercept and slope parameters.
b The estimated non-linear Markov-switching model has a regime-dependent intercept and allows for heteroscedasticity.
c A currency crisis is considered to occur in a year only if both the SETAR and Markov-switching model define a turmoil quarter (or quarters) during that year.
d Period in parenthesis indicates the period during which the data is available.
e No years shown here means that the SETAR model performs badly as it produces too many unnecessary crises. In this case, we ignore the SETAR model and treat turmoil years defined by the Markov-switching model as years of crises.

3.6 The Results

3.6.1 The Methodology

We follow the discrete-choice approach to link the dichotomous crisis variable to the selected variables, on the basis of recommendations made by the three generation crisis models. The limited dependent variable model is also called the qualitative or categorical variable model. Throughout each country group, we pool all the observations across both countries and time periods, and estimate the regression model of the form as

$$Pr(\text{crisis} = 1) = P = f(\beta'X)$$ \hspace{1cm} (3.8)

where $X$ is a vector of explanatory variables and $\beta$ is a vector of parameters. Here,
Table 3.6: Identification of Currency Crises in East Asia

<table>
<thead>
<tr>
<th>Countries</th>
<th>Currency Crises identified by SETAR Model ¹</th>
<th>Currency Crises defined by Markov-switching Model ²</th>
<th>Currency Crises defined by Both Models ³</th>
</tr>
</thead>
<tbody>
<tr>
<td>(1977-2001)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(1977-2001)</td>
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<td></td>
<td></td>
</tr>
<tr>
<td>(1971-2001)</td>
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<tr>
<td>(1977-2001)</td>
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<tr>
<td>(1972-2001)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(1977-2001)</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

¹ The estimated Non-linear SETAR model has a regime-dependent intercept and slope parameters.
² The estimated non-linear Markov-switching model has a regime-dependent intercept and allows for heteroscedasticity.
³ A currency crisis is considered to occur in a year only if both the SETAR and Markov-switching model define a turmoil quarter (or quarters) during that year.
⁴ Period in parenthesis indicates the period during which the data is available.
⁵ No years shown here means that the SETAR model performs badly as it produces too many unnecessary crises. In this case, we ignore the SETAR model and treat turmoil years defined by the Markov-switching model as years of crises.

we assume a Logistic distribution for \( f(\cdot) \) and a standard normal distribution with unit variance for the error term, giving us

\[
Pr(\text{crisis} = 1) = P = \frac{1}{1 + \exp(-X'\beta)} \tag{3.9}
\]

In other words, we estimate a Logit Model. The parameter vector \( \beta \) can be estimated by Maximum Likelihood. In order to interpret the estimated coefficients in the Logit model, we present the effect of a unit change in the \( j \)th explanatory variable on the probability \( P \) as

\[
\frac{\partial Pr(\text{crisis} = 1)}{\partial x_j} = \beta_j P(1 - P) \tag{3.10}
\]

which can be evaluated at the sample mean \( \bar{P} \).
# Table 3.7: Descriptive Statistics for Crisis Variables

<table>
<thead>
<tr>
<th></th>
<th>Latin America</th>
<th>Europe</th>
<th>East Asia</th>
</tr>
</thead>
<tbody>
<tr>
<td>Mean</td>
<td>0.12</td>
<td>0.19</td>
<td>0.23</td>
</tr>
<tr>
<td>Median</td>
<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
</tr>
<tr>
<td>Maximum</td>
<td>1.00</td>
<td>1.00</td>
<td>1.00</td>
</tr>
<tr>
<td>Minimum</td>
<td>0.00</td>
<td>0.00</td>
<td>0.00</td>
</tr>
<tr>
<td>Std. Dev.</td>
<td>0.33</td>
<td>0.40</td>
<td>0.42</td>
</tr>
<tr>
<td>Sum</td>
<td>18.00</td>
<td>35.00</td>
<td>37.00</td>
</tr>
</tbody>
</table>

## 3.6.2 Results for Latin American Countries

Following the conventional method, we include both the current and one-period lagged value of explanatory variables in the Logit model as the data we use is annual. Using the 10% significant level, we drop variables that are not statistically significant. The final estimated results for Latin American Countries are shown in Table 3.8. The fitted value of the crisis variable for Latin American countries during 1970-2001 is graphed on Figure 3.1 with its actual value.

As suggested by the first generation models, our results show that an unsustainable expansion in domestic credit increases the probability that a current fixed exchange rate can be abandoned in Latin America. More specifically, an increase in Domestic Credit provided by the Banking Sector by one percent of GDP raises the likelihood of a current devaluation by 0.58%. However, Overall Budget Balance is not statistically significant in this case.

It is also shown that a deteriorating situation in current accounts of Latin American countries has a both statistically and economically significant effect on the vulnerability of exchange rate regimes to speculative attacks. This too underlines the significant role that weak economic fundamentals play in Latin American turmoil. An increase in Current Account Deficits—a decrease in Current Account Balance—by one percent of GDP pushes the economy closer to a crash by more than 3% in short run. In return, currency crashes in Latin American countries drag the economies further into turmoil and lead to even higher current account deficits. Interestingly, one-period lagged International Reserves appears statistically signif-
Currency Crises in Latin America, Europe and East Asia

Figure 3.1: Actual and Fitted value of Crisis Variable for Latin American Countries

significant but has a negative sign, which contradicts what the Krugman-Flood-Garber model claims—the central bank's foreign reserves would \textit{gradually decrease} to monetize the budget deficits, until an abrupt speculative attack quickly exhausts all the remaining reserves and forces an abandonment of the parity.

Moreover, estimated results show that there exists a positive relationship between the domestic interest rates and the odds of a currency crisis occurring in Latin America. An increase in the domestic Interest Rates by one point (1%) relative to USA has a statistically significant impact on the likelihood of a crash, pushing it up by 0.004%. The relative small magnitude of the effect is a result of the extremely high market interest rates experienced by some Latin American countries during the crisis periods, e.g., 15778.6% in Brazil in 1990. This may offer some evidence that market expectations had played some role in Latin American crises before it was introduced in the second generation models. Other variables do not have statistically significant coefficients.
Figure 3.2: Actual and Fitted value of Crisis Variable for European Countries

It is seen in Figure 3.1 that trends in the short dash line representing the actual crisis variable are well matched by the solid line representing the fitted value for crisis variable. This means that nearly all the crisis incidents occurring in Latin America are explained by the Logit model. The McFadden R-Squared of the Logit model for Latin American countries is 0.39, high enough to show the good fitness of our model for the data.

3.6.3 Results for European Countries

Our estimated results for European countries are also shown in Table 3.8. As before, we drop the explanatory variables with parameters not statistically significant at the conventional level of 10%. Figure 3.2 provides a graph for both the actual and fitted value of crisis variables for European countries during 1970-2001.

As predicted by the second generation models, the real economic circle measured by the Real GDP Growth (annual %) is strongly associated with crisis incidences
in our results; a recession by one percent in real GDP undermines the possibility that the exchange rate regime will be successfully defended in the future by 5.6%. Surprisingly, one-year lagged unemployment does not have a predicted sign.

It is also shown in estimated results that one-period lagged domestic Interest Rate Differential (relative to German market) has a statistically significant effect of $-3.82\%$ on occurrence of a crisis. Together with no effect of its current value, it indicates a negative long-run relationship between interest rate differential and probability of a currency crash. But it is argued that it is the dramatic increases in market interest rate in Germany (the anchor country) rather than the decreases in other European countries’ interest rates that increases the likelihood of a future crisis in Europe. For example, German interest rate goes up from 4.3% to 10.18% in 1972-1973, from 5.87% to 11.26% in 1979-1981, and from 4% to 9.41% in 1988-1992, all of which are the periods of turmoil for most European countries.

Moreover, an abrupt and large loss of International Reserves coincides with the occurrence of a currency crash (14.02%) in European countries. This is in conformity with the first generation models which claim that all the remaining reserves will be eliminated when the speculative attacks force an abandonment of the exchange rate regime. As in the case of Latin America, one-period lagged reserves variable has a counter-intuitive sign, and thus the long-term relationship between international reserves and the likelihood of a crisis could be more complex than what Krugman (1979) and Flood and Garber (1984) predict. Moreover, Overall Budget Balance is significantly associated with currency crashes in Europe; higher budget deficits in European countries by 1% of GDP increase the likelihood of a future currency crisis by 5.32%. Therefore, our results show that factors identified by the first generation models, such as weak economic fundamentals and irresponsible government policies, still play a significant role in later currency crises in Europe.

Interestingly, the variable suggested by the third generation models—ratio of Bank Liquid Reserves to Bank Assets—also has a significant coefficient at 10% in European case. A one percentage point drop in the ratio of liquidity reserves to
assets in the banking sector makes the economy more vulnerable to speculative attacks in the future by 3.57%, while the current value of the variable reverses the effect, but by a smaller magnitude (1.38%). This shows that the problem of liquidity crunch experienced by East Asian countries in 1997/98 had existed in Europe, but the impact that it has undermining credibility of the exchange rate regime seems rather long-term than short-term as suggested by Change and Velasco (1998).

Finally, a strong performance by the German economy coincides with the crisis episodes occurring in other European countries; an increase in German real GDP growth by one percent reduces the probability that the exchange rate regimes in other European economies will be maintained in the future by 10.30%. This reflects the fact that the currency crises experienced by most European countries in the early 1990's are the periods when these economies had to adjust (devalue) their currencies against the Deutsche Mark as a result of their relatively poor economic performances.

Figure 3.2 shows that the actual crisis variable depicted by the short dash line moves nearly simultaneously with its fitted value represented by the solid line. In other words, it is shown that the periods of turmoil predicted by the model mostly match the actual crisis episodes in Europe. The McFadden R-Squared—measurement of Goodness of Fit—is 0.36 in the model for European countries, which again means that our Logit model is a good representation of the experiences of European countries in 1970-2001.

3.6.4 Results for East Asian Countries

We also report estimated results for East Asian countries in Table 3.8. All the explanatory variables with statistically insignificant coefficients at 10% have been dropped from the Logit model. We also present the actual and fitted value of crisis variables for East Asian countries during 1970-2001 in Figure 3.3.

According to the third generation models, the problem of moral hazard (international over-borrowing) plays a crucial role in the crisis episodes in East Asia, acting
as a main force for bubbles in asset markets. The estimated results show that an increase in Domestic Credit provided by Banking Sector by one percent of GDP raises the probability of a current forced abandonment of the parity by 1.75%. But the long-term effect of the credit expansion is more ambiguous, as its one-period lag has an opposite sign and a similar magnitude. Another suggested variable—Ratio of Bank Liquid Reserves to Bank Assets—is not statistically significant in the regression.

But we find evidence that shows factors identified by the first and second generation models still play significant parts in currency crises in East Asia. A sudden significant loss of International Reserves driven by speculative attacks would immediately pushes East Asian economies off the edge (by 24.46%). But as in the previous two cases, the positive sign of one-period lagged reserves variable means that reserves have a more sophisticated long-term relationship with the likelihood of a currency crash. It is also shown that there are signs of deteriorating economic
fundamentals in East Asian before crises. An increase in Current Account Deficits by one percent of GDP undermines the likelihood of a parity will be successfully defended in the future by 2.88%. But different from the case of Latin America, immediate improvements in all East Asian countries’ current accounts are under way just after the crises—for example, the 1997/98 crises saw that the current account balance (% of GDP) changed from $-2.27\%$ to $4.29\%$ in Indonesia, from $-4.42\%$ to $12.73\%$ in Korea, from $-5.92\%$ to $13.20\%$ in Malaysia, from $-5.28\%$ to $2.37\%$ in Philippines, from $13.82\%$ to $19.18\%$ in Singapore and from $-2.15\%$ to $12.29\%$ in Thailand. Moreover, the Overall Budget Balance has a counter-intuitive sign in the regression, which may reflect the fact that before crises most East Asian governments, unlike their counterparts in Latin America, are not engaged in irresponsibly expansionary fiscal policies.

A slowdown in East Asian economies also has a significant impact on the credibility of exchange rate regimes; an economic recession by one percent in Real GDP Growth in the previous and current year is associated with an increase in the probability of a currency crisis in East Asia by $3.29\%$ and $3.55\%$ respectively. A higher Interest Rate Differential (relative to USA market), which is a measure of inflationary expectations, pushes up the odds of a currency crash in East Asian markets (by $4.34\%$) as well. A high speculation of an impending depreciation may result in higher unemployment through wage inflation and also make serving debts more costly to the government who may find it optimal to devalue the domestic currency.

As in the case of Europe, one-period lagged Interest Rate Differential does not have a predicted sign in the regression for East Asia. It is also argued that a strong performance of the USA economy (although USA Real GDP Growth does not have a statistically significant coefficient) and the resultant high market interest rate in the second half of 1990's are major factors behind the reverse of capital flows in East Asia. The USA interest rate increases from $3.02\%$ in 1993 significantly to $5.84\%$ in 1995, and then gradually to $6.24\%$ in 2000.

It is seen in Figure 3.3 that the short dash line for the actual crisis variable
almost coincides with the solid line for the fitted crisis variable. Thus, it means that most of the currency crises in East Asian region can be explained by the Logit model. The McFadden R-squared in the model is 0.32 for East Asia, a promising figure showing that our model fits well the East Asian episodes.

3.7 The Conclusion

The initial approach to modeling currency crisis emphasizes the important role that unsustainable credit expansion and unsound economic fundamentals play in the late 1970s and 1980s' crises in Latin American countries. The second generation models suggest elements other than irresponsible policies and weak fundamentals—conflicting goals of government policies (the contingent nature of the government) combined with private sectors' perfect foresight—are the main causes of the 1992/93 ERM crises. The latest crisis models are inspired by the 1997/98 East Asian Currency/Financial Crises, identifying factors like moral hazard and illiquidity as potential factors that trigger the crises. In this chapter, however, we search for the stylized facts associated with currency crises occurring in these three regions—Latin America, Europe and East Asia, in order to see how good the three generation models are in explaining these crisis episodes that originally inspired them.

We construct a country-specific Market Pressure Index (MPI), which is the weighted average of the percentage change in nominal exchange rate, the difference in interest rate and the negative percentage change in international reserves. The aim in doing so is to capture not only the actual depreciations of currencies but also the ‘failed’ attacks successfully fended off by raising interest rates and/or selling out international reserves. The weights are chosen so that the three components have the same conditional standard deviations. The higher the index, the higher the speculative pressure the country is facing.

Based on the MPI, a 0–1 crisis variable is created by employing the SETAR (Self Exciting Threshold Autoregression) and Markov-switching model for each country. The advantage of adopting these non-linear state models is that we can avoid the
arbitrary choice of a threshold level a priori, as both models allow the data to reveal what move represents abnormal behaviour and therefore signal a crisis. Our results stem from the Limited Dependent Variable Model—Logit Model, which links the binary crisis variable to a wide range of selected variables based on the three generation models. We only keep the variables that are statistically significant at 10% conventional level.

We find that currency crises in Latin American countries seem to take place when (1) the domestic credit is expanding at an unsustainable pace, (2) the current account balance is deteriorating, and (3) there is currently a high expectation of realignment in markets.

Crises in European countries tend to occur when (1) the economy has slid into recession, (2) the international reserves are drying out, (3) the government has been engaged in irresponsibly expansionary fiscal policies, (4) the banking sector is facing a liquidity crunch, and (5) the German economy has performed relatively strongly and its interest rate has been high.

Finally, currency crashes in East Asian countries seem to happen when (1) there is a current unsustainable credit expansion as a result of moral hazard (international over-borrowing), (2) the government is running out international reserves, (3) the situation in current account has deteriorated, (4) the economy has been slowing down, (5) there is a high speculation of an impending devaluation, and (6) USA market interest rate has been high.

In summary, the estimated results from our research provide some evidence that the three generation models address the different features of currency crises occurring in three regions. For example, irresponsible credit expansion and unsound economic fundamentals identified by the first generation models are predominating factors in Latin America before crises, the close relationship between the economic cycle and occurrence of crises in Europe is pointed out in the second generation models, and moral hazard (international over-borrowing) defined by the third generation models does act as the main force behind the bubbles in East Asian asset markets.
But the overall picture is far from perfect. It is also shown in our results that some variables appear to have counter-intuitive signs and some are not statistically significant at the conventional level as they should be. This raises enquiries for further research in introducing new theories to explain the nature/causes of currency crises, which is discussed in the following Chapter.
Table 3.8: Logit Model Results for Latin America, Europe and East Asia

<table>
<thead>
<tr>
<th>Variable</th>
<th>Coefficients</th>
<th>Coefficients</th>
<th>Coefficients</th>
</tr>
</thead>
<tbody>
<tr>
<td>BANKLIQUID</td>
<td>-</td>
<td>-</td>
<td>-</td>
</tr>
<tr>
<td>BANKLIQUID(-1)</td>
<td>-3.57%**</td>
<td>-</td>
<td>-</td>
</tr>
<tr>
<td>CREDITBYBANK</td>
<td>1.75%***</td>
<td>-3.82%**</td>
<td>-3.39%**</td>
</tr>
<tr>
<td>CREDITBYBANK(-1)</td>
<td>-1.88%***</td>
<td>-24.46%***</td>
<td>-18.67%***</td>
</tr>
<tr>
<td>RESERVES</td>
<td>3.84%**</td>
<td>18.80%***</td>
<td>18.67%***</td>
</tr>
<tr>
<td>RESERVES(-1)</td>
<td>0.004%*</td>
<td>-3.82%**</td>
<td>-4.34%**</td>
</tr>
<tr>
<td>INTDIFF</td>
<td>0.004%*</td>
<td>0.004%*</td>
<td></td>
</tr>
<tr>
<td>INTDIFF(-1)</td>
<td>-3.82%**</td>
<td>4.34%**</td>
<td></td>
</tr>
<tr>
<td>UNEMPLOYMENT</td>
<td>-7.53%***</td>
<td></td>
<td></td>
</tr>
<tr>
<td>UNEMPLOYMENT(-1)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>CURACCOUNT</td>
<td>-3.23%***</td>
<td>3.26%**</td>
<td></td>
</tr>
<tr>
<td>CURACCOUNT(-1)</td>
<td></td>
<td>-3.26%**</td>
<td>-2.88%**</td>
</tr>
<tr>
<td>BUDGET</td>
<td></td>
<td>-5.32%**</td>
<td>2.02%*</td>
</tr>
<tr>
<td>BUDGET(-1)</td>
<td>-5.32%**</td>
<td>-3.55%**</td>
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</tr>
<tr>
<td>REALGDP</td>
<td></td>
<td>-5.60%*</td>
<td>-3.29%*</td>
</tr>
<tr>
<td>REALGDP(-1)</td>
<td>-5.60%*</td>
<td>-3.29%*</td>
<td></td>
</tr>
<tr>
<td>USA(GM)GDP</td>
<td>10.30%***</td>
<td></td>
<td></td>
</tr>
<tr>
<td>USA(GM)GDP(-1)</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

McFadden R-squared: 0.39 0.36 0.32

* Estimate Method: Maximum Likelihood (Quadratic hill climbing) for the Binary Logit model. Covariance matrix computed using second derivatives.

** Latin American countries include Argentina, Brazil, Chile, Columbia, Mexico and Peru.

*** European countries include Belgium, France, Ireland, Italy, Spain and UK.

**** East Asian countries include Indonesia, Korea, Malaysia, Philippines, Singapore and Thailand.

Estimation of effect of one-unit change in explanatory variable on the probability of a currency crisis (expressed in percentage) is evaluated at the sample mean.

Parameters with ** are significant at 5% level. Parameters with *** are significant at 1% level.

BANKLIQUID—Bank Liquid Reserves to Bank Assets ratio
CREDITBYBANK—Domestic Credit provided by Banking Sector (% of GDP)
RESERVES—International Reserves (months of imports)
INTDIFF—Interest Rate Differential relative to USA (Germany for Europe)
UNEMPLOYMENT—Unemployment (%)
CURACCOUNT—Current Account Balance (% of GDP)
BUDGET—Overall Budget Balance, including grants (% of GDP)
REALGDP—Real GDP Growth (annual %)
USA(GM)GDP—Annual Growth Rate of Real GDP in USA (Germany for Europe)
Table 3.9: SETAR Model\textsuperscript{a} Estimated Results\textsuperscript{b} in Latin America

<table>
<thead>
<tr>
<th></th>
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<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Intercept (Tranquil)</td>
<td>-1.53 \textsuperscript{d}</td>
<td>-6.00 \textsuperscript{d}</td>
<td>-1.66 \textsuperscript{d}</td>
<td>-0.78 \textsuperscript{d}</td>
<td>-3.85 \textsuperscript{d}</td>
<td>-9.89 \textsuperscript{d}</td>
</tr>
<tr>
<td>Intercept (Volatile)</td>
<td>63.74 \textsuperscript{e}</td>
<td>54.69 \textsuperscript{e}</td>
<td>13.27 \textsuperscript{e}</td>
<td>11.17 \textsuperscript{e}</td>
<td>35.51 \textsuperscript{e}</td>
<td>72.22 \textsuperscript{e}</td>
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<tr>
<td>Coefficients on AR(1) (Tranquil)</td>
<td>0.26 \textsuperscript{f}</td>
<td>0.10 \textsuperscript{f}</td>
<td>-0.08 \textsuperscript{f}</td>
<td>-0.02 \textsuperscript{f}</td>
<td>-0.12 \textsuperscript{f}</td>
<td>-0.13 \textsuperscript{f}</td>
</tr>
<tr>
<td>(Volatile)</td>
<td>-4.38 \textsuperscript{f}</td>
<td>1.80 \textsuperscript{f}</td>
<td>-1.09 \textsuperscript{f}</td>
<td>-0.18 \textsuperscript{f}</td>
<td>-1.53 \textsuperscript{f}</td>
<td>-2.13 \textsuperscript{f}</td>
</tr>
<tr>
<td>Coefficients on AR(2) (Tranquil)</td>
<td>2.77 \textsuperscript{f}</td>
<td>0.02 \textsuperscript{f}</td>
<td>0.25 \textsuperscript{f}</td>
<td>-0.17 \textsuperscript{f}</td>
<td>-2.13 \textsuperscript{f}</td>
<td>0.58 \textsuperscript{f}</td>
</tr>
<tr>
<td>(Volatile)</td>
<td>2.40 \textsuperscript{f}</td>
<td>0.05 \textsuperscript{f}</td>
<td>0.90 \textsuperscript{f}</td>
<td>-0.50 \textsuperscript{f}</td>
<td>-1.63 \textsuperscript{f}</td>
<td>0.67 \textsuperscript{f}</td>
</tr>
<tr>
<td>Coefficients on AR(3) (Tranquil)</td>
<td>-0.07 \textsuperscript{f}</td>
<td>(-0.97) \textsuperscript{f}</td>
<td>-0.16 \textsuperscript{f}</td>
<td>(-0.50) \textsuperscript{f}</td>
<td>0.003 \textsuperscript{f}</td>
<td>(0.05) \textsuperscript{f}</td>
</tr>
<tr>
<td>(Volatile)</td>
<td>0.17 \textsuperscript{f}</td>
<td>(0.34) \textsuperscript{f}</td>
<td>0.13 \textsuperscript{f}</td>
<td>(2.50) \textsuperscript{f}</td>
<td>0.02 \textsuperscript{f}</td>
<td>(0.03) \textsuperscript{f}</td>
</tr>
<tr>
<td>Coefficients on AR(4) (Tranquil)</td>
<td>0.13 \textsuperscript{f}</td>
<td>(2.50) \textsuperscript{f}</td>
<td>0.02 \textsuperscript{f}</td>
<td>(0.03) \textsuperscript{f}</td>
<td>0.003 \textsuperscript{f}</td>
<td>(0.00) \textsuperscript{f}</td>
</tr>
<tr>
<td>(Volatile)</td>
<td>0.003 \textsuperscript{f}</td>
<td>(0.00) \textsuperscript{f}</td>
<td>0.003 \textsuperscript{f}</td>
<td>(0.00) \textsuperscript{f}</td>
<td>0.003 \textsuperscript{f}</td>
<td>(0.00) \textsuperscript{f}</td>
</tr>
<tr>
<td>Std. Dev. (Tranquil)</td>
<td>29.68</td>
<td>16.17</td>
<td>4.55</td>
<td>1.92</td>
<td>12.29</td>
<td>35.85</td>
</tr>
<tr>
<td>Std. Dev. (Volatile)</td>
<td>80.10</td>
<td>46.96</td>
<td>7.89</td>
<td>8.22</td>
<td>25.66</td>
<td>95.72</td>
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<tr>
<td>Threshold</td>
<td>36.76</td>
<td>23.92</td>
<td>4.34</td>
<td>2.78</td>
<td>10.61</td>
<td>25.83</td>
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<td>SC criterion</td>
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<td>8.00</td>
<td>6.61</td>
<td>4.94</td>
<td>8.40</td>
<td>10.47</td>
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<td>LR Linearity test \textsuperscript{e}</td>
<td>89.77</td>
<td>111.80</td>
<td>60.95</td>
<td>84.85</td>
<td>43.77</td>
<td>73.44</td>
</tr>
</tbody>
</table>

\textsuperscript{a} Estimated Model: Non-linear SETAR (Self Exciting Threshold Autoregression) with a regime-dependent intercept and slope parameter.

\textsuperscript{b} All results are obtained using H.M. Krolzig's MSVAR package for Ox.

\textsuperscript{c} Period in parenthesis indicates the period during which the data is available.

\textsuperscript{d} Number in parenthesis below regression coefficient is t-value.

\textsuperscript{e} Number in square bracket below test statistic is p-values.
Table 3.10: Markov-switching Model Estimated Results in Latin America

<table>
<thead>
<tr>
<th></th>
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<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>Intercept (Tranquil)</td>
<td>2.64 (\pm) 1.26</td>
<td>-2.68 (\pm) 1.64</td>
<td>0.12 (\pm) 0.30</td>
<td>-0.11 (\pm) 0.56</td>
<td>-0.70 (\pm) 0.76</td>
<td>-1.72 (\pm) 0.97</td>
</tr>
<tr>
<td>Intercept (Volatile)</td>
<td>69.91 (\pm) 0.50</td>
<td>13.21 (\pm) 0.31</td>
<td>-0.74 (\pm) 0.32</td>
<td>2.82 (\pm) 0.39</td>
<td>6.53 (\pm) 0.19</td>
<td>12.76 (\pm) 0.30</td>
</tr>
<tr>
<td>AR(1)</td>
<td>0.06 (\pm) 1.32</td>
<td>0.06 (\pm) 1.13</td>
<td>-0.003 (\pm) 2.44</td>
<td>0.16 (\pm) 0.97</td>
<td>0.04 (\pm) 0.56</td>
<td>0.03 (\pm) 0.77</td>
</tr>
<tr>
<td>AR(2)</td>
<td>0.02 (\pm) 2.21</td>
<td>0.02 (\pm) 0.81</td>
<td>0.02 (\pm) 0.35</td>
<td>0.03 (\pm) 0.37</td>
<td>-0.06 (\pm) 1.34</td>
<td>-0.02 (\pm) 0.62</td>
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<tr>
<td>AR(3)</td>
<td>0.06 (\pm) 0.56</td>
<td>0.06 (\pm) 0.16</td>
<td>-0.04 (\pm) 2.44</td>
<td>-0.17 (\pm) 0.34</td>
<td>-0.02 (\pm) 1.65</td>
<td>0.05 (\pm) 0.34</td>
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<tr>
<td>AR(4)</td>
<td>0.05 (\pm) 3.17</td>
<td>0.08 (\pm) 1.57</td>
<td>0.06 (\pm) 2.94</td>
<td>0.11 (\pm) 1.61</td>
<td>0.05 (\pm) 0.74</td>
<td>0.05 (\pm) 1.51</td>
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<tr>
<td>Std. Dev. (Tranquil)</td>
<td>16.25 (\pm) 0.25</td>
<td>16.40 (\pm) 1.96</td>
<td>2.84 (\pm) 1.22</td>
<td>1.22 (\pm) 7.28</td>
<td>17.93 (\pm) 17.93</td>
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<tr>
<td>Std. Dev. (Volatile)</td>
<td>263.67 (\pm) 5.87</td>
<td>121.96 (\pm) 121.96</td>
<td>18.61 (\pm) 18.61</td>
<td>19.53 (\pm) 19.53</td>
<td>68.82 (\pm) 68.82</td>
<td>191.97 (\pm) 191.97</td>
</tr>
<tr>
<td>Transition Prob. (%)</td>
<td>0.97 (\pm) 0.97</td>
<td>0.95 (\pm) 0.95</td>
<td>0.86 (\pm) 0.86</td>
<td>0.92 (\pm) 0.92</td>
<td>0.95 (\pm) 0.95</td>
<td>0.96 (\pm) 0.96</td>
</tr>
<tr>
<td>SC criterion</td>
<td>9.63 (\pm) 0.33</td>
<td>9.36 (\pm) 0.33</td>
<td>9.54 (\pm) 0.42</td>
<td>5.21 (\pm) 0.60</td>
<td>7.89 (\pm) 0.03</td>
<td>9.63 (\pm) 0.40</td>
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<tr>
<td>LR Linearity test *</td>
<td>144.76 (\pm) 0.005</td>
<td>115.05 (\pm) 0.000</td>
<td>53.52 (\pm) 0.000</td>
<td>77.55 (\pm) 0.000</td>
<td>88.24 (\pm) 0.000</td>
<td>190.33 (\pm) 0.000</td>
</tr>
</tbody>
</table>

* Estimated Model: Non-linear Markov-switching Autoregressive (MS-AR) model with a regime-dependent intercept and allowing for heteroscedasticity in the disturbance term.

b All results are obtained using H.-M. Krolzig's MSVAR package for Ox.

c Period in parenthesis indicates the period during which the data is available.

d Number in parenthesis below regression coefficient is t-value.

e Number in square bracket below test statistic is p-values.
<table>
<thead>
<tr>
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</thead>
<tbody>
<tr>
<td>Intercept (Tranquil)</td>
<td>4.60 (d)</td>
<td>-1.53</td>
<td>-2.29</td>
<td>-1.03</td>
<td>-2.48</td>
<td>-0.95</td>
</tr>
<tr>
<td>Intercept (Volatile)</td>
<td>5.76</td>
<td>0.99</td>
<td>1.35</td>
<td>1.45</td>
<td>2.07</td>
<td>1.11</td>
</tr>
<tr>
<td>Coefficients on AR(1) (Tranquil)</td>
<td>0.05</td>
<td>0.06</td>
<td>0.10</td>
<td>0.03</td>
<td>0.41</td>
<td>0.21</td>
</tr>
<tr>
<td>(Volatile)</td>
<td>-0.23</td>
<td>0.30</td>
<td>0.33</td>
<td>0.52</td>
<td>0.17</td>
<td>0.03</td>
</tr>
<tr>
<td>Coefficients on AR(2) (Tranquil)</td>
<td>0.03</td>
<td>0.12</td>
<td>0.06</td>
<td>(-0.14)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>(Volatile)</td>
<td>-0.02</td>
<td>-0.06</td>
<td>0.12</td>
<td>0.01</td>
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</tr>
<tr>
<td>Coefficients on AR(3) (Tranquil)</td>
<td>0.10</td>
<td>0.04</td>
<td>0.06</td>
<td>-0.14</td>
<td></td>
<td></td>
</tr>
<tr>
<td>(Volatile)</td>
<td>0.03</td>
<td>-0.05</td>
<td>0.06</td>
<td>0.07</td>
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<td></td>
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<tr>
<td>Coefficients on AR(4) (Tranquil)</td>
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<tr>
<td>(Volatile)</td>
<td>0.07</td>
<td>0.02</td>
<td>-0.23</td>
<td>-0.07</td>
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</tr>
<tr>
<td>Std. Dev. (Tranquil)</td>
<td>4.50</td>
<td>0.94</td>
<td>2.06</td>
<td>1.07</td>
<td>1.76</td>
<td>1.38</td>
</tr>
<tr>
<td>Std. Dev. (Volatile)</td>
<td>2.83</td>
<td>1.16</td>
<td>1.31</td>
<td>1.03</td>
<td>1.92</td>
<td>1.65</td>
</tr>
<tr>
<td>Threshold</td>
<td>0.99</td>
<td>-0.06</td>
<td>-0.35</td>
<td>0.48</td>
<td>-0.39</td>
<td>0.10</td>
</tr>
<tr>
<td>SC criterion</td>
<td>6.00</td>
<td>3.46</td>
<td>4.14</td>
<td>3.51</td>
<td>4.69</td>
<td>3.30</td>
</tr>
<tr>
<td>LR Linearity test *</td>
<td>124.30</td>
<td>86.41</td>
<td>84.02</td>
<td>98.87</td>
<td>89.74</td>
<td>72.82</td>
</tr>
</tbody>
</table>

* Estimated Model: Non-linear SETAR (Self Exciting Threshold Autoregression) with a regime-dependent intercept and slope parameter.

b All results are obtained using H.-M. Krolzig's MSVAR package for Ox.

a Estimated Model: Non-linear SETAR (Self Exciting Threshold Autoregression) with a regime-dependent intercept and slope parameter.

b Period in parenthesis indicates the period during which the data is available.

c Number in parenthesis below regression coefficient is t-value.

d Number in square bracket below test statistic is p-value.
### Table 3.12: Markov-switching Model Estimated Results in Europe

<table>
<thead>
<tr>
<th></th>
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</tr>
</thead>
<tbody>
<tr>
<td>Intercept (Tranquil)</td>
<td>-5.60 (4.26)</td>
<td>-1.58 (3.03)</td>
<td>-6.63 (3.96)</td>
<td>0.05 (0.36)</td>
<td>-2.38 (2.18)</td>
<td>-2.25 (1.05)</td>
</tr>
<tr>
<td>Intercept (Volatile)</td>
<td>6.48 (12.66)</td>
<td>0.82 (1.34)</td>
<td>3.25 (0.70)</td>
<td>-0.16 (0.18)</td>
<td>2.91 (2.29)</td>
<td>0.89 (1.07)</td>
</tr>
<tr>
<td>AR(1)</td>
<td>-0.26 (4.25)</td>
<td>0.27 (2.59)</td>
<td>0.04 (0.57)</td>
<td>0.43 (4.64)</td>
<td>0.10 (0.82)</td>
<td>0.27 (3.41)</td>
</tr>
<tr>
<td>AR(2)</td>
<td>-0.99 (12.66)</td>
<td>-0.00 (1.00)</td>
<td>0.05 (0.57)</td>
<td>-0.05 (0.54)</td>
<td>-0.05 (0.54)</td>
<td>-0.05 (0.54)</td>
</tr>
<tr>
<td>AR(3)</td>
<td>-0.09 (1.34)</td>
<td>-0.09 (1.34)</td>
<td>-0.09 (1.34)</td>
<td>-0.09 (1.34)</td>
<td>-0.09 (1.34)</td>
<td>-0.09 (1.34)</td>
</tr>
<tr>
<td>AR(4)</td>
<td>-0.02 (0.00)</td>
<td>-0.02 (0.00)</td>
<td>0.005 (0.00)</td>
<td>-0.10 (0.00)</td>
<td>-0.10 (0.00)</td>
<td>-0.10 (0.00)</td>
</tr>
<tr>
<td>Std. Dev. (Tranquil)</td>
<td>5.77 (12.55)</td>
<td>1.35 (2.50)</td>
<td>1.81 (3.50)</td>
<td>1.50 (5.00)</td>
<td>3.66 (8.00)</td>
<td>2.27 (2.27)</td>
</tr>
<tr>
<td>Std. Dev. (Volatile)</td>
<td>6.13 (12.55)</td>
<td>2.00 (2.50)</td>
<td>2.20 (3.50)</td>
<td>2.60 (5.00)</td>
<td>2.54 (8.00)</td>
<td>2.26 (2.26)</td>
</tr>
<tr>
<td>Transition Prob. (%)</td>
<td>0.79 (0.06)</td>
<td>0.67 (0.06)</td>
<td>0.96 (0.06)</td>
<td>0.98 (0.06)</td>
<td>0.88 (0.06)</td>
<td>0.87 (0.06)</td>
</tr>
<tr>
<td>(Tranquil → Tranquil)</td>
<td>0.40 (0.10)</td>
<td>0.74 (0.10)</td>
<td>0.32 (0.10)</td>
<td>0.90 (0.10)</td>
<td>0.87 (0.10)</td>
<td>0.83 (0.10)</td>
</tr>
<tr>
<td>(Volatile → Volatile)</td>
<td>6.96 (0.00)</td>
<td>4.12 (0.00)</td>
<td>4.53 (0.00)</td>
<td>4.09 (0.00)</td>
<td>5.62 (0.00)</td>
<td>4.01 (0.00)</td>
</tr>
</tbody>
</table>

- Estimated Model: Non-linear Markov-switching Autoregressive (MS-AR) model with a regime-dependent intercept and allowing for heteroscedasticity in the disturbance term.
- All results are obtained using H.M. Krolzig's MSVAR package for Ox.
- Period in parenthesis indicates the period during which the data is available.
- Number in parenthesis below regression coefficient is t-value.
- Number in square bracket below test statistic is p-value.
Table 3.13: SETAR Model \textsuperscript{a} Estimated Results \textsuperscript{b} in East Asia

<table>
<thead>
<tr>
<th></th>
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</thead>
<tbody>
<tr>
<td>Intercept (Tranquil)</td>
<td>(-1.02 \textsuperscript{d})</td>
<td>(-0.76)</td>
<td>(-1.18)</td>
<td>(-2.94)</td>
<td>(-2.04)</td>
<td>(-1.77)</td>
</tr>
<tr>
<td>Intercept (Volatile)</td>
<td>(-2.43)</td>
<td>(-3.60)</td>
<td>(-8.70)</td>
<td>(-7.77)</td>
<td>(-12.18)</td>
<td>(-9.36)</td>
</tr>
<tr>
<td>Coefficients on AR(1) (Tranquil)</td>
<td>0.09</td>
<td>0.04</td>
<td>0.17</td>
<td>0.11</td>
<td>-0.02</td>
<td>-0.12</td>
</tr>
<tr>
<td>(Volatile)</td>
<td>0.06</td>
<td>0.44</td>
<td>2.14</td>
<td>1.19</td>
<td>-0.20</td>
<td>-1.58</td>
</tr>
<tr>
<td>Coefficients on AR(2) (Tranquil)</td>
<td>-0.01</td>
<td>0.40</td>
<td>0.46</td>
<td>0.31</td>
<td>0.10</td>
<td>0.21</td>
</tr>
<tr>
<td>(Volatile)</td>
<td>(-0.02)</td>
<td>1.02</td>
<td>3.86</td>
<td>(3.86)</td>
<td>1.40</td>
<td>(1.45)</td>
</tr>
<tr>
<td>Coefficients on AR(3) (Tranquil)</td>
<td>-0.12</td>
<td>-0.07</td>
<td>-1.31</td>
<td>-0.98</td>
<td>0.07</td>
<td>0.98</td>
</tr>
<tr>
<td>(Volatile)</td>
<td>-0.01</td>
<td>0.46</td>
<td>0.46</td>
<td>0.31</td>
<td>0.10</td>
<td>0.21</td>
</tr>
<tr>
<td>Coefficients on AR(4) (Tranquil)</td>
<td>-0.19</td>
<td>-0.20</td>
<td>-2.51</td>
<td>-1.70</td>
<td>0.07</td>
<td>0.23</td>
</tr>
<tr>
<td>(Volatile)</td>
<td>-0.20</td>
<td>(1.70)</td>
<td>-2.51</td>
<td>(-1.70)</td>
<td>0.07</td>
<td>0.23</td>
</tr>
<tr>
<td>Std. Dev. (Tranquil)</td>
<td>3.83</td>
<td>1.86</td>
<td>1.01</td>
<td>2.08</td>
<td>0.85</td>
<td>1.41</td>
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<tr>
<td>Std. Dev. (Volatile)</td>
<td>10.80</td>
<td>3.48</td>
<td>0.81</td>
<td>2.41</td>
<td>0.87</td>
<td>2.37</td>
</tr>
<tr>
<td>Threshold</td>
<td>4.54</td>
<td>1.65</td>
<td>-0.16</td>
<td>-0.18</td>
<td>-0.96</td>
<td>0.21</td>
</tr>
<tr>
<td>SC criterion</td>
<td>6.08</td>
<td>4.54</td>
<td>2.02</td>
<td>5.10</td>
<td>2.83</td>
<td>4.32</td>
</tr>
<tr>
<td>LR Linearity test \textsuperscript{a}</td>
<td>73.81</td>
<td>54.36</td>
<td>94.68</td>
<td>83.69</td>
<td>93.78</td>
<td>84.23</td>
</tr>
</tbody>
</table>

\textsuperscript{a} Estimated Model: Non-linear SETAR (Self Exciting Threshold Autoregression) with a regime-dependent intercept and slope parameter.

\textsuperscript{b} All results are obtained using H-M. Krolzig's MSVAR package for Ox.

\textsuperscript{c} Period in parenthesis indicates the period during which the data is available.

\textsuperscript{d} Number in parenthesis below regression coefficient is t-value.

\textsuperscript{e} Number in square bracket below test statistic is p-values.
Table 3.14: Markov-switching Model \(^a\) Estimated Results \(^b\) in East Asia

<table>
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</thead>
<tbody>
<tr>
<td>Intercept (Tranquil)</td>
<td>-0.35 (^d)</td>
<td>(-0.60)</td>
<td>-1.07</td>
<td>-3.43</td>
<td>-0.66</td>
<td>0.45</td>
</tr>
<tr>
<td></td>
<td>((-0.59))</td>
<td>((-1.84))</td>
<td>((-2.66))</td>
<td>((-1.64))</td>
<td>((1.71))</td>
<td>((1.42))</td>
</tr>
<tr>
<td>Intercept (Volatile)</td>
<td>4.10</td>
<td>(-0.36)</td>
<td>0.66</td>
<td>4.54</td>
<td>0.66</td>
<td>1.42</td>
</tr>
<tr>
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<td>((-0.15))</td>
<td>((0.75))</td>
<td>((3.05))</td>
<td>((0.97))</td>
<td>((1.08))</td>
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<tr>
<td>AR(1)</td>
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<td>0.37</td>
<td>-0.08</td>
<td>0.13</td>
<td>-0.11</td>
</tr>
<tr>
<td></td>
<td>((0.97))</td>
<td>((2.67))</td>
<td>((3.07))</td>
<td>((-0.86))</td>
<td>((0.96))</td>
<td>((-1.01))</td>
</tr>
<tr>
<td>AR(2)</td>
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<td>((-0.54))</td>
<td>0.13</td>
<td>-0.12</td>
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<tr>
<td></td>
<td>((3.43))</td>
<td>((1.71))</td>
<td>((-0.54))</td>
<td>((-1.32))</td>
<td>((-1.01))</td>
<td>((-1.32))</td>
</tr>
<tr>
<td>AR(3)</td>
<td>0.06</td>
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<td>((-0.01))</td>
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<tr>
<td></td>
<td>((0.83))</td>
<td>((0.41))</td>
<td>((-0.51))</td>
<td>((-0.80))</td>
<td>((-0.01))</td>
<td>((-0.80))</td>
</tr>
<tr>
<td>AR(4)</td>
<td>-0.18</td>
<td>-0.02</td>
<td>0.03</td>
<td>((-0.33))</td>
<td>((-0.33))</td>
<td>((-0.33))</td>
</tr>
<tr>
<td>Std. Dev. (Tranquil)</td>
<td>2.48</td>
<td>1.16</td>
<td>1.69</td>
<td>2.60</td>
<td>1.41</td>
<td>1.17</td>
</tr>
<tr>
<td>Std. Dev. (Volatile)</td>
<td>15.19</td>
<td>7.68</td>
<td>1.97</td>
<td>4.44</td>
<td>1.82</td>
<td>3.21</td>
</tr>
<tr>
<td>Transition Prob. (%)</td>
<td>0.50</td>
<td>0.93</td>
<td>0.67</td>
<td>0.46</td>
<td>0.74</td>
<td>0.86</td>
</tr>
<tr>
<td>(Tranquil → Tranquil)</td>
<td>((0.93))</td>
<td>((0.93))</td>
<td>((0.67))</td>
<td>((0.46))</td>
<td>((0.74))</td>
<td>((0.86))</td>
</tr>
<tr>
<td>(Volatile → Volatile)</td>
<td>0.30</td>
<td>0.44</td>
<td>0.61</td>
<td>0.70</td>
<td>0.01</td>
<td>0.80</td>
</tr>
<tr>
<td>SC criterion</td>
<td>0.02</td>
<td>4.41</td>
<td>3.84</td>
<td>5.66</td>
<td>3.68</td>
<td>4.84</td>
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<tr>
<td>LR Linearity test (^e)</td>
<td>66.35</td>
<td>72.97</td>
<td>21.43</td>
<td>22.78</td>
<td>23.31</td>
<td>57.27</td>
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</tbody>
</table>

\(^a\) Estimated Model: Non-linear Markov-switching Autoregressive (MS-AR) model with a regime-dependent intercept and allowing for heteroscedasticity in the disturbance term.

\(^b\) All results are obtained using H-M. Krolzig's MSVAR package for Ox.

\(^c\) Period in parenthesis indicates the period during which the data is available.

\(^d\) Number in parenthesis below regression coefficient is t-value.

\(^e\) Number in square bracket below test statistic is p-values.
Currency Crises in Latin America, Europe and East Asia

Figure 3.4: Argentina: Estimated Probabilities of being in Volatile State by SETAR Model 1979-2001

Figure 3.5: Belgium: Estimated Probabilities of being in Volatile State by SETAR Model 1970-2001

Figure 3.6: Brazil: Estimated Probabilities of being in Volatile State by SETAR Model 1970-2001
Figure 3.7: Chile: Estimated Probabilities of being in Volatile State by SETAR Model 1977-2001

Figure 3.8: Columbia: Estimated Probabilities of being in Volatile State by SETAR Model 1986-2001

Figure 3.9: France: Estimated Probabilities of being in Volatile State by SETAR Model 1970-2001
Figure 3.10: Indonesia: Estimated Probabilities of being in Volatile State by SETAR Model 1978-2001

Figure 3.11: Ireland: Estimated Probabilities of being in Volatile State by SETAR Model 1973-2001

Figure 3.12: Italy: Estimated Probabilities of being in Volatile State by SETAR Model 1971-2001
Figure 3.13: South Korea: Estimated Probabilities of being in Volatile State by SETAR Model 1977-2001

Figure 3.14: Malaysia: Estimated Probabilities of being in Volatile State by SETAR Model 1971-2001

Figure 3.15: Mexico: Estimated Probabilities of being in Volatile State by SETAR Model 1981-2001
Figure 3.16: Peru: Estimated Probabilities of being in Volatile State by SETAR Model 1970-2001

Figure 3.17: Philippine: Estimated Probabilities of being in Volatile State by SETAR Model 1977-2001

Figure 3.18: Singapore: Estimated Probabilities of being in Volatile State by SETAR Model 1972-2001
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4.1 Introduction

The mainstream literature on the credibility issue of an exchange rate regime (see Backus and Driffield (1985a) and Backus and Driffield (1985b)) argues that a defence of the fixed parity can signal the monetary authority's 'toughness' towards inflation, and thus improve the credibility of the regime. More recently, literature has re-examined this issue in the context of a dynamic model where there are persistence effects from failing to devalue following an adverse supply effect. Thus in Drazen and Masson (1994) there is persistence in unemployment, and defending the fixed parity leads to higher future unemployment undermining the possibility that the currency peg will be maintained. In Benigno and Missale (2001) -henceforth BM- the fundamentals are driven not by unemployment, but by public debt. In both models if there is little uncertainty about the government's preferences, or if the difference between a perceived 'tough' and a 'weak' government is small, resisting a currency crisis may in fact reduce the credibility of the exchange rate regime.
Following BM’s approach, we adopt a three-period stochastic version of the Barro and Gordon (1983) model, where the probability of devaluation in each period is derived from the monetary authority’s optimization problem. Monetary policy is conducted in terms of an escape clause that specifies the threshold value of a negative supply-side shock above which devaluation will occur. We present a more general framework than BM does for studying the policymaker’s optimization problem, a framework which distinguishes between the commitment and discretionary optimal escape clauses. We also extend BM’s model to allow for non-deflatable debt, that is in cases where the government also issues securities whose real returns cannot be eroded through an unexpected devaluation.

First, we examine the case where the central bank can commit itself to a 2-period escape clause rule that is publicly known. Minimization of the bank’s loss function delivers the optimum solution with respect to the policymaker’s utility, which we regard as a benchmark. Then, we move to the complete information game, where the monetary authority cannot commit, but its type is known to the public. As in BM, it is shown that defending the fixed parity and refraining from inflationary financing increases the debt burden and thus the likelihood of having to resort to future devaluation. In other words, with a high level of nominal debt a current devaluation always increases the probability that a future defence of the new parity will succeed. This is the ‘debt-burden’ effect. We also confirm the BM result that the probability of a first-period devaluation also increases with the level of public debt and, more interestingly, with the share of short-term debt. Comparison with the commitment case shows that the discretionary two-period escape clause rules are sub-optimal and involve an expected inflation bias. This can be lessened by the delegation of monetary policy to a tough banker, but this relocates the problem to one of establishing the credibility of such toughness.

This leads to an asymmetric information game where the monetary authority’s preferences are not known to the private sector. The decision to devalue might reveal a weak authority, thus leading to inflationary expectations which, in turn, increase
the likelihood of a future devaluation. This is the 'signalling' effect. This effect is important when there is substantial uncertainty about the cost of devaluation to the monetary authority and when the debt is small. In this case, reputational motive, especially when a monetary authority is facing a high level of debt, could increase the probability that the fixed parity will be maintained in the first period, as the tough authority wants to signal its type and the weak authority pretends to be tough in order to exploit lower interest rates. Whether the debt-burden or the signalling effect prevails depends on the importance of debt fundamentals relative to the monetary authority’s credibility.

The probability of a second-period devaluation does not depend on either debt maturity or reputational considerations. Therefore, the theoretical model predicts that for any given level of uncertainty, in countries where the debt-burden effect dominates, a higher ratio of short-term debt implies higher short-term interest rates (and a higher probability of current devaluation) and a flatter yield curve. In contrast, in countries where the signalling effect prevails, a larger proportion of short-term debt is associated with lower current interest rates and a steeper yield curve.

We use the Reinhart and Rogoff (2004) index to define the actual changes in the exchange rate regime for 11 emerging economies from 1993Q2 to 2001Q4. We adopt the Self Exciting Threshold Autoregression (SETAR) model and the Markov Regime Switching (MRS) model to identify periods with high pressure in exchange markets. We identify observations where foreign exchange market pressure is accommodated (i.e. there is a shift in the regime) and where it is not (i.e. the monetary authority successfully defends the peg by selling international reserves and/or increasing interest rates). Countries where the debt-burden effect dominates are subsequently distinguished from others where the signalling effect prevails, based on the behaviour of market expectations following successful defences and/or actual crises (regime shifts). The theory’s two predictions about the impact of debt maturity on the likelihood of crisis and the yield curve are then tested using GMM (General Method of Moments) estimation for dynamic panel data.
This chapter is organized as follows. Section 2 sets out an open-economy stochastic Barro-Gordon model with short-term and long-term debt. Section 3 examines three cases: commitment, and non-commitment, first with symmetric information and then with asymmetric information. The results are demonstrated by numerical simulation in Section 4. Section 5 presents the interested variables and data set needed for the empirical analysis. In Section 6, 11 emerging economies in our sample are split into two groups—'signalling' countries and 'debt-burden' countries. We then test the two predictions made by the theoretical model in Section 7. Section 8 concludes.

4.2 The Model

We consider a three-period open economy of the Barro and Gordon (1983) model where the unexpected inflation that follows a devaluation increases output both through a standard price-output effect and through a tax-reduction effect. As in Obstfeld (1994) and Velasco (1996), we assume that taxes are only levied in the last period and that taxes involve an output cost. These tax distortions increase the incentive to devalue. Following the escape clauses approach of Obstfeld (1997), we assume that output is stochastic to take into account that even a tough policymaker, with a high aversion to inflation and therefore to devaluation, will devalue if the economy is hit by unusually large shocks. We also take the size of a devaluation as given, i.e. not depending on the magnitude of the shock. In other words, the monetary authority's choice is between maintaining a fixed parity and the alternative of a devaluation of a fixed size.

Because of Purchasing Power Parity, we can make the simplifying assumption that the inflation rate $\pi_t$, in periods $t = 1, 2$ equals the rate of devaluation. Therefore, the output in period 1 is given by

$$y_1 = y^* + \alpha(\pi_1 - E_0\pi_1) - k - u_1$$  \hspace{1cm} (4.1)
where \( y^* \) is the target output, \( \pi_1 - E_0 \pi_1 \) is inflation surprise and \( k \) is a goods or labour market distortion. \( \pi_t = 0 \) when the monetary authority maintains the fixed parity and \( \pi_t = d \) when the monetary authority devalues. Thus the equilibrium level of output under perfect foresight (no shocks) is \( y^* - k \), which is considered to be too low by the authority. The output in period 2 is given by

\[
y_2 = y^* + \alpha(\pi_2 - E_1 \pi_2) - k - T - u_2
\]  

(4.2)

where \( T \) is the output cost of distortionary taxation required to repay debt \( B = B_{10} + B_{20} \), where \( B_{10} \) = one-period (short-term) real debt and \( B_{20} \) = two-period (long-term) real debt, both issued in period 0. Let \( \mu \in [0, 1] \) be the proportion of inflation-sensitive debt.

Consider accumulated short-term and long-term debt, \( B_{12} \) and \( B_{22} \) respectively, at the end of periods \( t = 1, 2 \). Let \( r_t \) = real interest rate in period \( t \) given by \( r_t = \pi_t - \pi_t \) to a linear approximation. For nominal inflation-sensitive short-term debt, \( \pi_t = E_{t-1} \pi_1 + E_{t-1} \pi_t \), is the nominal interest rate at the beginning of period \( t \) demanded by the private sector to achieve an expected real interest rate of \( E_{t-1} r_t \). Therefore \( r_t = E_{t-1} r_t - (\pi_t - E_{t-1} \pi_t) = r^* - (\pi_t - E_{t-1} \pi_t) \) where we assume \( E_{t-1} r_t = r^* \), the fixed foreign real interest rate. Similarly for long-term debt \( r_t = r^* - (\pi_t - E_0 \pi_t) \). Assuming \( r^* = 0 \) for convenience, we then have accumulated short-term and long-term debt at the end of period 2 given by

\[
B_{12} = [1 - \mu(\pi_1 - E_0 \pi_1) - \mu(\pi_2 - E_1 \pi_2)]B_{10}
\]

\[
B_{22} = [1 - \mu(\pi_1 - E_0 \pi_1) - \mu(\pi_2 - E_0 \pi_2)]B_{20}
\]  

(4.3)

At the end of period 2, taxes \( T \) must be levied to repay the accumulated debt, \([B_{12} + B_{22}]\). Then assuming the effect on output to be proportional to debt we then have from (4.3)

\[
T = \tau[1 - \mu(\pi_1 - E_0 \pi_1) - \mu(\pi_2 - E_1 \pi_2)]B_{10} + \tau[1 - \mu(\pi_1 - E_0 \pi_1) - \mu(\pi_2 - E_0 \pi_2)]B_{20}
\]
Normalizing $\tau = 1^1$ and substituting (4.2) into the output equation (4.2) then yields

$$y_2 - y^* = (\pi_2 - E_0 \pi_1)m + (\pi_1 - E_0 \pi_1)\mu B - (E_1 \pi_2 - E_0 \pi_2)S - K - u_2$$  \hspace{1cm} (4.5)

where $B = B_{10} + B_{20}$ as before, is the real value of total debt at period 0, $m = \alpha + \mu B$, $S = \alpha + \mu B_{10}$ and $K = k + B$ is the expected deviation of output from target. Our basic model is then given by (4.1) and (4.5). The output effect of a revision in expectations depends not only on total debt $B$, but also on $S = \alpha + \mu B_{10}$ which increases with short-term inflation-sensitive debt, $\mu B_{10}$. If a devaluation leads to an upward revision in the interest rate (a first-period devaluation produces expectations of a further depreciation), then the authority is worse off the shorter the maturity of its debt (the higher $S$), because short-term debt is refinanced at higher-than-expected interest rates. On the other hand, however, if the inflation created by devaluation in the first period eases the real debt burden and thus diminishes the need for further devaluation, then the benefits for the monetary authority increase with short-term debt (the higher $S$) because interest rates turn out to be lower than expected.

At $t = 0$, the monetary authority of type $i = W$(Weak), $T$(Tough) is assumed to minimize its loss function:

$$L_0 = E_0 \left[ \sum_{t=1}^{2} \beta^{t-1} L_t^i \right] = E_0[L_1^i + \beta L_2^i]$$  \hspace{1cm} (4.6)

where $E_0$ denotes expectations conditional on the information in period 0, $0 < \beta < 1$ is the discount factor and $L_t^i$ is a single-period loss function in which the monetary authority weighs the cost of devaluation against the output deviation from the target.

---

1This implies that debt itself is measured in terms of output loss.
\[ y^*: \]
\[ L_t^i = \theta^i \pi_t^2 + (y_t - y^*)_t; \quad t = 1, 2 \]  
\[(4.7)\]

where \( \theta^i \) measures the cost of devaluation relative to output for type \( i \). Our model demonstrates that \( L_1^i \) and \( L_2^i \) are functions:

\[ L_1^i = L_1^i(\pi_1, E_0 \pi_1, u_1); \]
\[ L_2^i = L_2^i(\pi_1, E_0 \pi_1, \pi_2, E_0 \pi_2, E_1 \pi_2, u_2) \]

(in addition to fixed parameters \( k, y^*, B_{10} \) and \( B_{20} \)) since in period 2 taxes \( T \) are repaid.

### 4.3 Three Possible Equilibria

#### 4.3.1 Commitment under Complete Information

This section assumes complete information and presents a more general framework than BM in that it considers the case where the monetary authority can commit itself to some predefined rules. Firstly, we set out the general framework.

The two-period expected loss function in period 0 can be rewritten as

\[ \Lambda_0 = E_0(L_1) + \beta [\rho_1 E_0 L_2(D) + (1 - \rho_1) E_0 L_2(F)] \]  
\[(4.8)\]

where \( \rho_1 \) = probability of a first-period devaluation and \( L_2(h) \) is the welfare loss in period 2 following a history \( h = D(\text{Devaluation}), F(\text{Maintaining the fixed parity}) \). Here we focus on \( \rho_1 \) and \( \rho_2(h) \) as instruments for the monetary authority to minimize its loss. As we shall see, this is equivalent to choosing a threshold \( \hat{u}_1 \) in the first period for the magnitude of negative supply-side shock at which the authority devalues, and a state-contingent shock \( \hat{u}_2 \) which depends on the realization of \( \hat{u}_1 \) in period 1 in BM. Under complete information, the type of the monetary authority (the cost of devaluation, \( \theta \)) is known to the private sector so we suppress the type...
superscript in this and the next section to ease the notation.

The sequence of events under commitment is as follows. At period 0, the monetary authority issues fixed-rate one-period (short-term) and two-period (long-term) bonds, including debts denominated in a foreign currency—or the monetary authority is obliged to pay off all the debts issued by the government. The monetary authority commits itself to the two-period rule consisting of probability of devaluation in period 1, \( \rho_1 \) and in period 2, \( \rho_2(h), h = D, F \) following a history \( h = D, F \) in period 1. As mentioned above, the rule is equivalent to the shock thresholds \( \tilde{u}_1 \) and \( \tilde{u}_2(h) \) at which devaluation occurs in period 1 and in period 2 respectively.

The interest rates of the debts are determined by (rational) expectation of inflation rates (rates of devaluation). Then in periods 1 and 2 shocks occur and the monetary authority implements this commitment rule to devalue or maintain the parity. At the end of period 1 the one-period debt is rolled over and at the end of period 2 the total debt service is repaid by levying distortionary taxes.

Now let \( f(u) \) be the probability density function for the disturbance \( u_t, t = 1, 2 \) in (4.1) and (4.5). Then the probability, as of period 1, of a second-period devaluation is given by

\[
\rho_2(h) = \Pr[u_2 > \tilde{u}_2(h)] = \int_{\tilde{u}_2(h)}^{\infty} f(u)du
\]

where \( \tilde{u}_2(h) \) is a threshold value such that if \( u_2 > \tilde{u}_2(h) \) devaluation occurs in period 2. As before, the index \( h = D, F \) indicates that the monetary authority devalued or maintained the fixed parity in period 1. We assume that the shock \( u_2 \) is uniformly distributed over the interval \([-v, v]\), then \( f(u) = \frac{1}{2v} \). So we can rewrite (4.9) as

\[
\rho_2(h) = \frac{v - \tilde{u}_2(h)}{2v}
\]

(4.10)

where \( \tilde{u}_2(h) \in [-v, v] \). Similarly, we can evaluate the probability of a devaluation in period 1 as

\[
\rho_1 = \frac{v - \tilde{u}_1}{2v}
\]

(4.11)
If the bounds for the shocks are large then $\rho_1, \rho_2(h) \in [0,1]$ is ensured. In fact in what follows we assume that

$$2v > dm$$

(4.12)

which (recalling $m = \alpha + \mu B$ and $S = \alpha + \mu B_{10}$) implies that $2v > dS$.

The expected inflation in period 2 formed in period 1 is now

$$E_1\pi_2(h) = dp_2(h) + 0 \times (1 - \rho_2(h)) = dp_2(h)$$

(4.13)

Using backward induction, we can then obtain the expectation of inflation in period 2 formed in period 0:

$$E_0\pi_2 = \rho_1 E_1\pi_2(D) + (1 - \rho_1) E_1\pi_2(F) = [\rho_1 p_2(D) + (1 - \rho_1) p_2(F)]d$$

(4.14)

and the expected inflation in period 1:

$$E_0\pi_1 = \rho_1 d$$

(4.15)

The expected inflation rates ($E_0\pi_1$, $E_0\pi_2$) are then a measure of the yield curve over the 2 periods. Integrating $L_2(h) = \theta \pi_2^2 + [y_2(h) - y^*]^2$ over the interval $[-v, v]$ and using equation (4.11), we can evaluate the expected loss $E_1L_2(h)$ in period 2, $h = D, F$, as

$$E_1L_2(h) = \frac{1}{2v} \left\{ \int_{-v}^{\pi_2(h)} (y_2 - y^*)^2 du_2 + \int_{\pi_2(h)}^{v} [\theta d^2 + (y_2 - y^*)^2] du_2 \right\}$$

(4.16)

because $\pi_1 = 0$ if $h = F$, and $\pi_1 = \delta$ if $h = D$. This expectation in period 1 is formed knowing the history $h = D, F$, but not yet knowing the realization of the shock $u_2$.

After considerable algebra we obtain from (4.16):

$$E_1L_2(h) = \frac{v^2}{3} + [\Pi(h)]^2 + dp_2(h)[\theta d + 2m\Pi(h) + 2mu(\rho_2(h) - 1) + dm^2]$$

(4.17)
where

$$\Pi(h) = -mE_0\pi_2 + (\pi_1(h) - E_0\pi_1)B - (E_1\pi_2(h) - E_0\pi_2)S - K$$  \hspace{1cm} (4.18)

which depends on the history $h = D,F$ at the end of period 1. We can now rewrite (4.5) as

$$y_2 - y^* = m\pi_2 + \Pi(h) - u_2$$  \hspace{1cm} (4.19)

In period 0, the expected welfare loss for the first period is

$$E_0L_1 = \frac{1}{2\theta} \left\{ \int_{-\infty}^{\infty} (y_1 - y^*)^2 du_1 + \int_{-\delta}^{\delta} [\theta\sigma^2 + (y_1 - y^*)^2] du_1 \right\}$$  \hspace{1cm} (4.20)

Let $\Theta = -\alpha E_0\pi_1 - k = -\alpha d_0 - k$. Then

$$y_1 = y^* + \alpha(\pi_1 - E_0\pi_1) - k - u_1 = y^* + \alpha \pi_1 + \Theta - u_1.$$  

Then using (4.17), we can rewrite $E_0L_1$ as

$$E_0L_1 = \frac{y^3}{3} + \Theta^2 + d_0[\theta \sigma^2 + 2\alpha(\rho_1 - 1) + 2\alpha \Theta + \alpha^2 d]$$  \hspace{1cm} (4.21)

Note that $E_0L_2(h) = E_0(E_1L_2(h)) = E_1L_2(h)$, since $L_2(h)$ is independent of the first-period shock $u_1$ that is the only random variable in period 1.

Substituting (4.13) - (4.15) into (4.18) and noting that $m = \alpha + \mu B = \alpha + \mu(B_2 + B_1)$, $S = \alpha + \mu B_1$, we arrive at

\begin{align*}
\Pi(F) &= -\mu B_20(\rho_1\rho_2(D) + (1 - \rho_1)\rho_2(F))d - \rho_1d\mu B - \rho_2(F)dS - K  \\
\Pi(D) &= -\mu B_20(\rho_1\rho_2(D) + (1 - \rho_1)\rho_2(F))d + (1 - \rho_1)d\mu B - \rho_2(D)dS - K
\end{align*}

It is convenient to treat these expressions as constraints that the monetary authority faces when it minimizes its loss function $A^*_0$ with respect to its instruments.
\( \rho_1, \rho_2(h) \) in period 0. Therefore, we form a Lagrangian:

\[
\mathcal{L}_0 = \lambda_0 + \lambda_F(\Pi(F) + \mu B_{20}(\rho_1 \rho_2(D) + (1 - \rho_1) \rho_2(F))d + \rho_1 d\mu B + \rho_2(F)dS + K)
\]
\[
+ \lambda_D(\Pi(D) + \mu B_{20}(\rho_1 \rho_2(D) + (1 - \rho_1) \rho_2(F))d - (1 - \rho_1) d\mu B + \rho_2(D)dS + K)
\]

(4.24)

Minimizing \( \mathcal{L}_0 \) with respect to \( \rho_1, \rho_2(h), \Pi(h) \), we get the five first-order conditions (f.o.c.) as shown below:

\[
\frac{\partial \mathcal{L}_0}{\partial \rho_1} = d[(\theta + \alpha^2)d + 2\alpha \rho_1(2v - \alpha d) - 2\alpha v)] + \beta(E_0 L_2(D) - E_0 L_2(F))
\]
\[
+ (\lambda_F + \lambda_D)(\mu B_{20}(\rho_2(D) - \rho_2(F)) + \mu B)d = 0
\]

(4.25)

\[
\frac{\partial \mathcal{L}_0}{\partial \rho_2(D)} = \beta \rho_1 d[\theta d + 2m \Pi(D) + 2mv(2\rho_2(D) - 1) + dm^2]
\]
\[
+ (\lambda_F + \lambda_D)\mu B_{20}\rho_1 + \lambda_D dS = 0
\]

(4.26)

\[
\frac{\partial \mathcal{L}_0}{\partial \rho_2(F)} = \beta (1 - \rho_1) d[\theta d + 2m \Pi(F) + 2mv(2\rho_2(F) - 1) + dm^2]
\]
\[
+ (\lambda_F + \lambda_D)\mu B_{20}(1 - \rho_1) + \lambda_F dS = 0
\]

(4.27)

\[
\frac{\partial \mathcal{L}_0}{\partial \Pi(D)} = 2\beta \rho_1 (\Pi(D) + md \rho_2(D)) + \lambda_D = 0
\]

(4.28)

\[
\frac{\partial \mathcal{L}_0}{\partial \Pi(F)} = 2\beta (1 - \rho_1)(\Pi(F) + md \rho_2(F)) + \lambda_D = 0
\]

(4.29)

These 5 equations plus (4.22) and (4.23) can now be solved for the seven variables \( \rho_1, \rho_2(D), \rho_2(F), \Pi(D), \Pi(F), \lambda_D \) and \( \lambda_F \). This is the commitment solution for the monetary authority; in other words, in period 0 the monetary authority commits to the 'policy rules' in the form of escape clauses \( \rho_1, \rho_2(h) \) which are carried out in period 1 and 2.

It is of particular interest to examine the effect of devaluation or not in period 1 on the probability of devaluation in period 2. Eliminating the Lagrangian multipliers from (4.28) and (4.29) in (4.26) and (4.27) and after a little algebra we arrive at

\[
\rho_2(F) - \rho_2(D) = \frac{d\mu^2 B_{20}}{(\alpha + \mu B)(2v - dS) - \mu B_{20}dS}
\]

(4.30)
Condition (4.12) ensures that $(\alpha + \mu B)(2v - dS) - \mu B_{20}dS > 0$. Thus under commitment $\rho_2(F) - \rho_2(D) \geq 0$, so the probability of a devaluation following a fixed parity in period 1 is at least as big as that following a devaluation. If there is no long-term debt $B_{20} = 0$ then debt has no effect on $\rho_2$ and $\rho_2(F) = \rho_2(D)$, and as the long-term composition of debt increases to the point where $B = B_{20}$ and $S = \alpha + \mu B_{10} = 1$ then $\rho_2(F) - \rho_2(D)$ increases to $\frac{2d_0^2}{2v(\alpha + \mu B) - d(\alpha + \mu B)}$. Thus we have:

**Proposition 1.** Under commitment, the probability of a devaluation in period 2 following a fixed parity in period 1, $\rho_2(F)$, is the same as that following a devaluation, $\rho_2(D)$ if there is no long-term debt. As the composition of long-term debt increases, then $\rho_2(F) - \rho_2(D) > 0$ and increases. If there is any long-term debt then $\rho_2(F) - \rho_2(D) > 0$ also increases with total debt.

The intuition behind this result is as follows. The central bank when committing to $\rho_1$, $\rho_2(D)$, $\rho_2(F)$ implies a commitment to the first-period nominal interest rate (equal to $\rho_1 d$) and the second-period nominal interest rate following first-period devaluation (D) or not (F), equal to $\rho_2(D)d$ and $\rho_2(F)d$ respectively. This is also confirmed by our prerequisites that the realization of the shocks and devaluation or not in the first period by the monetary authority, are publicly observed under complete information. If there is only short-term debt, then any actions taken by the authority in period 1 (devalue or not) do not have any impact on how the short-term debt is rolled over at end of that period, as everything has been defined in the commitment. Therefore, $\rho_2(D) = \rho_2(F)$. However, if there is any long-term debt then a first period devaluation that follows a sufficiently large negative supply-side shock, will erode some of the debt and thereby reduce the need for a further devaluation in the second period. Thus the optimal second-period choice with commitment sees $\rho_2(D) < \rho_2(F)$.

It is useful to examine a special case where there is no debt and $B = T = 0$ and

---

2This follows because (4.12)$\Rightarrow 2v > d(\alpha + \mu B) > d \left(S + \frac{S}{(\alpha + \mu B)\mu B_{20}}\right)$. 

---
$m = S = \alpha$. In this case we have

\[ y_t = y^* + \alpha(\pi_t - E_{t-1}\pi_t) - \kappa - u_t; \quad t = 1, 2 \]  \hspace{1cm} (4.31)

and there is no structural dynamics in the model. The policy-problem in each period is now identical and is an escape-clause version of the Barro-Gordon stabilization problem. It is straightforward to show that this special case of the above problem leads to

\[ \rho_1 = \rho_2(D) = \rho_2(F) = \frac{1}{2} - \frac{\theta d}{2\alpha(2\nu - \alpha\delta)} \]  \hspace{1cm} (4.32)

If we strengthen (4.12) to

\[ 2\nu \geq dm + \theta d \]  \hspace{1cm} (4.33)

then expected inflation in each period is non-negative, even when there is no debt.

### 4.3.2 Discretion under Complete Information

As private expectations of inflation in period 1 are formed in period 0, \( \pi_1 = \rho_1 d \) where \( \rho_1 \) is found from the f.o.c. above, will no longer be optimal once the monetary authority can choose its own monetary policies (without commitment). Similarly in period 1, the authority will also find \( \pi_2(h) = \rho_2(h)d \) no longer an optimal policy after the private sector revises its expectation of the second-period inflation where again \( \rho_2(h) \) is the commitment rule in period 2. With discretion the monetary authority will minimize its loss function taking private expectations of inflation in all periods as given, i.e., taking \( \Theta \) and \( \Pi(h) \) as given. This means that the constraints (4.22) and (4.23) do not bind; i.e., \( \lambda_F = \lambda_D = 0 \) in our original Lagrangian function (4.24).

The sequence of events under discretion is now as follows. At period 0, the private sector forms expectations about inflation in periods 1 and 2, which determine the fixed interest rates of the one-period (short-term) and two-period (long-term) bond issues respectively. In period 1, after the realization of the output shock, the monetary authority decides whether to devalue or maintain the fixed parity. At the end of period 1, the one-period debt is rolled over at the interest rate that
is determined by the revision of the private sector's expectation of second-period inflation. In period 2, after the realization of a second shock, the authority decides whether to devalue and finally repays the debt by levying distortional taxes. The game tree for this sequence of events is given in Figure 4.1.

Figure 4.1: The Game Tree under Discretion. The dotted lines refer to the case of asymmetric information examined in Section 4.3.3.
To solve the discretionary case, we first examine the impact of a first-period devaluation on the probability of a second-period devaluation and then go back to the first-period problem. Putting $\lambda_F = \lambda_D = 0$ in our f.o.c (4.26) and (4.27) above, we arrive at

$$\rho_2(h) = \frac{1}{2} - \frac{1}{2v} \left[ \frac{\theta d}{2m} + \frac{dm}{2} - mE_0\pi_2 + (\pi_1 - E_0\pi_1)B - (E_1\pi_2(h) - E_0\pi_2)S - K \right]$$

(4.34)

The likelihood of a devaluation in period 2, $\rho_2(h)$, depends on whether the authority has devalued in period 1, both directly through, $\pi_1$, and through a revision in expectations, $E_1\pi_2(h) - E_0\pi_2$. A devaluation in period 1, i.e. $\pi_1 = d$, reduces the likelihood of a second-period devaluation as unexpected inflation reduces the real debt burden. This is also because a downward revision in expected inflation, $E_1\pi_2$, and thus in the interest rate, decreases the debt burden in the second period to the extent that the debt is short-term.

Using equation (4.34) and (4.13), we have

$$\rho_2(F) - \rho_2(D) = \frac{d\mu B}{2v - dS} > 0$$

(4.37)

Thus under discretion, the probability of a second-period crisis after maintaining the parity, $\rho_2(F)$, is greater than that after a devaluation, $\rho_2(D)$. Once again, it is shown that—given complete information—a first-period devaluation always at least improves the likelihood that the new parity will be maintained. But in the discretion case, we can now see that $\rho_2(F)$ is always greater than $\rho_2(D)$ no matter whatever the composition of debt is. Let $\Delta \rho^D_2$ and $\Delta \rho^F_2$ be the difference in probabilities under
commitment and discretion respectively. Then from (4.30) and (4.37) we have

$$\Delta \rho_2^D - \Delta \rho_2^C = \frac{d \mu BS(2v - dm)}{(2v - dS)[(\alpha + \mu B)(2v - dS) - \mu B_{20} dS]} > 0$$  (4.38)

since by the large shock condition (4.12), $2v - dm > 0$ and the numerator has been shown to be positive. To summarize:

**Proposition 2.** Under complete information and discretion, the probability of a devaluation in period 2 following a fixed parity in period 1, $\rho_2(F)$, is greater than that following a devaluation, $\rho_2(D)$ irrespective of the composition of debt. The difference $\rho_2(F) - \rho_2(D)$ is greater under discretion than under commitment.

A first-period devaluation causes unexpected inflation, which reduces the real debt burden and thus the expected deviation of second-period output from target. Devaluation also affects the probability of a devaluation in period 2 through a downside revision in expectations and hence a lower-than-expected interest rate at which the short-term debt is rolled over. Thus, a devaluation in period 1 reduces the likelihood of a second-period crisis.

As shown in proposition 1, in the commitment case with complete information, short-term debt cannot act as a channel whereby the first-period devaluation decreases the likelihood of a second-period crisis. In other words, whether the authority devalues in period 1 or not does not affect the probability of a devaluation in period 2 if $B = B_{10}$. But as we have already found in the case of discretion under complete information, the current devaluation can ease the debt burden for the future, and the short-term debt is crucial to this effect since the benefits of first-period devaluation are magnified by the amount of debt that is rolled over. On the one hand $\Delta \rho_2^C$ does not depend on $B_{10}$ in the commitment case, while on the other hand a rise in $B_{10}$ increases the difference $\rho_2(F) - \rho_2(D)$ in the discretion case. Therefore, $\Delta \rho_2^D$ is greater than $\Delta \rho_2^C$ as long as there exists some short-term debt.

Now consider the authority’s choice in period 1. We can rewrite equation (4.17)
Substituting $\lambda_D = \lambda_F = 0$ into (4.25) and noting that $E_0L_2(h) = E_1L_2(h)$, we can get the likelihood of a first-period devaluation as follows:

$$\rho_1 = \frac{1}{2\alpha(2v - \alpha d)/d} \{d(2\alpha(k + v) - (\theta + \alpha^2)d) - \beta[E_1L_2(D) - E_1L_1(F)]\}$$

which is equivalent to

$$\hat{\alpha}_1 = \frac{1}{2\alpha d} \{\theta^2d^2 + \alpha^2d - 2\alpha d(E_0\pi_1 + k) - \beta[E_1L_1^2(F) - E_1L_2^2(D)]\}

the threshold value of the first-period shock for a crisis to occur.

From (4.14), which still applies, (4.35) and (4.36) we have

$$E_0\pi_2 = \frac{d}{2v - dm}[v + K - \theta d - \frac{dm}{2}]$$

which shows that expected inflation and thus the probability of a devaluation in period 2 does not depend on either the term structure of debt or $\rho_1$, the probability of the first-period devaluation. This is because the maturity of the debt affects both the probability of a devaluation in period 1 and (both short- and long-term) interest rates. For given interest rates, a shorter maturity, which increases the probability of a devaluation in period 1, tends to reduce the likelihood of a second devaluation. However, a shorter maturity also increases interest rates and tax distortions with offsetting effects on the probability of a second-period devaluation.

Using (4.39), (4.40) and (4.41), the probability of a first-period devaluation can be derived as follows

$$\rho_1 = \frac{1}{2} + \frac{(2\alpha k - \theta d)(2v - dS)^2 + 4v\beta(2v - dS)\mu B(k + B)}{2\alpha(2v - \alpha d)(2v - dS)^2 - 4v\beta(2v - dm)d(\mu B)^2}$$

The probability of a devaluation decreases with its cost to the authority, $\theta$, and...
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increases with distortions, chiefly, with the debt burden, $B$. By differentiating $\rho_1$ with respect to $B_{10}$ while holding the total debt constant, we can get

$$\frac{\partial \rho_1}{\partial B_{10}} \propto k + \mu B + \frac{2\nu - dm}{2\nu - dS} (2E\pi_1 - d) \mu B \geq 0$$

(4.43)

Hence we arrive at

**Proposition 3.** Under complete information and discretion, the probability of a first-period crisis increases with the share of short-term debt $B_{10}$ and with the volume of inflation-sensitive debt, $\mu B$.

The effect of a history of devaluation or maintaining the fixed rate in period 1 on the probability of a crisis in period 2 depends on the term $S = \alpha + \mu B_{10}$, which increases with short-term debt $B_{10}$. The bigger proportion of short-term debt makes maintaining the fixed parity in the first period more costly leading to a greater chance that the monetary authority might choose to devalue in period 1.

Intuitively, it would appear that more short-term debt makes the inflation rate which is expected after maintaining the parity higher while making inflation expected after a devaluation lower. In other words, long-term debt can minimize the probability of an exchange-rate devaluation in the first period when the monetary authority's preferences are known to the market.

As we did for commitment, it is useful to examine a special case where there is no debt and $B = T = 0$ and $m = S = 1$ for discretion as well. It is straightforward to show that this special case of the above problem leads to

$$\rho_1 = \rho_2(D) = \rho_2(F) = \frac{1}{2} - \frac{\theta d - 2\alpha k}{2\alpha(2v - ad)}$$

(4.44)

Comparing (4.44) and (4.32) we see that monetary policy with discretion (and no debt) leads to a familiar expected inflationary bias of $\frac{k}{(2v - ad)}$ which arises because the central bank has an ambitious output target above the natural rate of $k > 0$. One way of eliminating this inflationary bias is for a government with preference $\theta = \theta^G$ to delegate to an independent 'conservative' banker in the Rogoff-sense with $\theta^D > \theta^G$.
such that \((\theta^D - \theta^G)d = 2ak\). However the toughness of the new authority may lack credibility, which leads us to the next section where information is asymmetric.

Our results yield testable predictions regarding the yield curve \((E_0\pi_1, E_0\pi_2)\). As shown from (4.41), the expected second-period inflation rate increases with the total volume of debt\(^3\), \(B\), and with the proportion of inflation-sensitive debt \(\mu\), but is independent of its maturity structure. Combined with the previous proposition we then have:

**Proposition 4.** The expected second-period inflation rate increases with the total volume of debt and the ratio of inflation-sensitive debt, but is independent of its maturity structure. Together with proposition 3 this implies that in the case of discretion under complete information, the yield curve becomes flatter as the proportion of short-term debt increases.

### 4.3.3 Discretion under Asymmetric Information

We now consider the case where monetary policy is conducted by a monetary authority with preferences unknown to the public. We assume there are two possible types of authorities: a ‘tough’ authority with a relatively high weight on inflation in (4.7) \(\theta = \theta^T\), and a ‘weak’ authority with a preference parameter \(\theta^W < \theta^T\). As discussed above, one interpretation of this asymmetric information is that the authority with preference \(\theta = \theta^G\) delegate to an independent ‘conservative’ banker in the Rogoff-sense with \(\theta^D > \theta^G\), but the credibility of this central bank independence from the authority needs to be tested by the public in a process of Bayesian learning.

The solution to the case of discretion under asymmetric information follows the complete information case, proceeding by backwards induction starting in period 2 but with the following changes: the probabilities of devaluation in periods 1 and 2 now are \(\rho_1, \rho_2(D), \rho_2(F)\), \(i = T, W\) are now type-dependent; in periods \(t = 0, 1, 2\), the private sector attaches a probability \(q_t\) that the authority is tough (i.e., \(\theta = \theta^T\))

\(^3\)This follows from the fact that \(\frac{d}{dn} \left[ v + K - \frac{dn}{2} \right] > 0\).
and a probability $1 - q_t$ that it is weak ($\theta = \theta^W$); in period 0 the private sector has a prior $q_0$ which is up-dated at the end of each of the following two periods observing devaluation $D$ or a maintenance of the fixed exchange rate $F$ and possibly the realization of output shocks (or an estimate of shocks) $u_t, t = 1, 2$ occurring during the period. It is also assumed that the debt maturity is the same for both types of authorities, as are their interest rates, $E_0\pi_1$ and $E_0\pi_2$. Then following the first-period shock each type of monetary authority will decide whether to devalue or not, taking into account the impact of this decision on the beliefs of the private sector entering period 2.

Private sector beliefs now become

$$E_0\pi_1 = [q_0\rho^T + (1 - q_0)\rho^W]d$$

$$E_0\pi_2 = q_0d[\rho^T_1 \rho^T_2 (D) + (1 - \rho^T_1) \rho^T_2 (F)] + (1 - q_0)d[\rho^W_1 \rho^W_2 (D) + (1 - \rho^W_1) \rho^W_2 (F)]$$

$$E_1\pi_2(h) = (1 - q_1(h))d\rho^W_2(h) + q_1(h)d\rho^T_2(h)$$

where the history $h = [j, u_1], j = D, F$ consists of two observations by the public, whether there is a change in exchange rate or not and the shock in period 1. From the Appendix B we can now show

$$E_1\pi_2(F \cap u_1) - E_1\pi_2(D \cap u_1) = \frac{d^2}{2v - dS} \left[ \mu B - q_1(F \cap u_1) \frac{(\theta^F - \theta^W)}{2m} \right]$$

Then using (4.35) and (4.36), we have

$$\rho_2^T(F \cap u_1) - \rho_2^T(D \cap u_1) = \frac{d\mu B}{2v} + \frac{dE_1\pi_2(F \cap u_1) - E_1\pi_2(D \cap u_1)}{2v} \frac{S}{2v}$$

$$= \frac{d\mu B}{2v} + \frac{d^2 S}{2v(2v - dS)} \left[ \mu B - q_1(F \cap u_1) \frac{(\theta^F - \theta^W)}{2m} \right]$$

(4.49)
The right-hand-side of (4.49) can be either positive or negative. On the one hand, a first-period devaluation may diminish the likelihood of a second-period devaluation by easing the debt burden. If a devaluation in period 1 leads to a lower interest rate than a successful defence of the parity, then the authority gets a second-period gain. That is because the short-term debt, $\mu B_{10}$, is rolled over at a lower-than-expected interest rate. This is exactly what we see in the case of complete information. Intuitively, for the 'debt-burden' effect to dominate, the level of the inflation-sensitive debt must be high relative to uncertainty about the government type.

On the other hand, a devaluation in period 1 could send a signal of a weak authority which is heading for further devaluation and thus may lead to higher-than-expected inflation and interest rates in period 2. In the case that the interest rate rises following a devaluation turn the (4.49) negative, the authority expects a second-period loss from abandoning fixed parity. This case is relevant when the uncertainty over monetary authority's preferences is great (or the difference between preferences, $\theta^p - \theta^w$, is large) relative to the level of deflatable debt, $\mu B$, and successful defence of the current exchange rate regime sends a strong signal of the authority's determination not to devalue; that is, when the 'signaling' effect prevails over the debt-burden effect.

Interestingly, (4.49) shows that whether the exchange rate regime gains or loses credibility does not depend on the maturity of the debt; instead, the short-term debt, $B_{10}$, increases the difference in the probabilities of a second-period devaluation, since either benefits or costs of first-period devaluation are magnified by the amount of debt that is rolled over. In addition, comparing (4.37) with (4.48) we arrive at the result:

**Proposition 5.** Under discretion and asymmetric information, the credibility of the exchange regime is increased by a successful defence if and only if the difference between preferences, $\theta^p - \theta^w$, is large relative to the level of deflatable debt, $\mu B$. Moreover, the difference in probabilities $\rho_2(F \cap u_1) - \rho_2(D \cap u_1)$ is less than that under
complete information owing to the signalling effect which depends on the degree of uncertainty \( \theta^T - \theta^W \).

This is because reputation considerations provide both type of authorities with an incentive to defend the exchange rate—for a tough authority to signal its type, and for a weak authority to pretend to be tough in order to exploit lower interest rates. The greater the difference in the authority’s preferences \( \theta^T - \theta^W \) is, the stronger the incentive is.

If we denote \( \Delta \rho_2^D = \rho_2(F \cap u_1) - \rho_2(D \cap u_1) \) under discretion with asymmetric information, and \( \Delta \rho_2^C \) and \( \Delta \rho_2^F \) under discretion with complete information, and commitment with complete information respectively, as before, we now see that if \( dB \geq q_1(F \cap u_1)(\theta^T - \theta^W) > 0 \) then

\[
\Delta \rho_2^C \leq \Delta \rho_2^D < \Delta \rho_2^F \tag{4.50}
\]

In fact given the level and composition of debt, and \( q_1(F \cap u_1) \) under discretion, there exists an optimal degree of ‘ambiguity’ \( \theta^T - \theta^W \) that will achieve the first-best difference in probabilities,

\[
\theta^T - \theta^W = \frac{2mBS(2v - dm)}{[(1 + B)(2v - dS) - B20dS]q_1(F \cap u_1)} \tag{4.51}
\]

Interestingly, this is another an example where if commitment is impossible, complete transparency (complete information) is counterproductive.

The First-period Solutions for the Separating Equilibrium

The characterization of the possible equilibria when the private sector can make inferences after observing both whether there is a devaluation or not and the shock in the first period, is rather complicated involving pooling or separating equilibria depending on the realization of \( u_1 \). Here we follow the simplifying assumption of Drazen and Masson (1994) to focus only on such separating equilibrium that action would fully reveal, with a weak authority finding it optimal to devalue and
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a tough authority not to devalue, at the realization of shocks. In other words, a monetary authority is believed to be weak if it chooses to devalue, \( q_t(D \cap u_t) = 0, t = 1, 2 \); while it is expected to continue defending the parity if it maintains the fixed rate in the first period, \( q_t(F \cap u_t) = 1, t = 1, 2 \). It should be noted that this assumption does not imply that the tough authority will never devalue or the weak authority always devalues. Actually, both types of authorities will choose to devalue if the economy is hit by shocks that exceed their respective threshold levels. The separating equilibrium can be summarized by the following equations:

\[
\rho^i_t = \frac{1}{2\alpha(2\nu - \alpha d)} \left\{ d(2\alpha(k + \nu) - (\theta^i + \alpha^2)d) - \beta[E_t^d I_2^d(D) - E_t^d I_1^d(F)] \right\}; \ i = W, T \tag{4.52}
\]

where \( E_t^i[\cdot] \) signifies the private expectations of the monetary authority of type \( i \).

\[
E_t^d I_2^d(h) = \frac{\nu^2}{3} + [\Pi(h)]^2 + d\rho^d_2(h)[\theta^d d + 2m\Pi(h) + 2mv(\rho^d_2(h) - 1) + dm^2]; \ h = D, F \tag{4.53}
\]

where, as before,

\[
\Pi(h) = -mE_0\pi_2 + (\pi_1(h) - E_0\pi_1)\mu_B - (E_1\pi_2(h) - E_0\pi_2)S - K \tag{4.54}
\]

where \( \pi_1(D) = d \) and \( \pi_1(F) = 0 \).

\[
\rho^d_2(h) = \frac{1}{2} - \frac{1}{2\nu} \left( \frac{\theta^d d}{2m} + \frac{dm}{2} + \Pi(h) \right) \tag{4.55}
\]

Substituting (4.55) into (4.53), we can obtain

\[
E_t^d I_2^d(h) = \frac{\nu^2}{3} + [\Pi(h)]^2 - 2dmv(\rho^d_2(h))^2 \tag{4.56}
\]

which replaces (4.53). From (4.46) and (4.55) we now have

\[
E_0\pi_2 = \frac{d}{2\nu - dm} \left[ v + K - \frac{(q_0\theta^d + (1 - q_0)\theta^w)d}{2m} - \frac{dm}{2} \right] \tag{4.57}
\]
The second-period devaluation expected by the private sector in period 0 is the same as in the case of complete information, except for the cost of devaluation, \( \theta \), which is now replaced by its expectation under asymmetric information. As before, the debt maturity does not affect \( E_0 \pi_2 \). We also have

\[
E_0 \pi_1 = [q_0 \rho^T_1 + (1 - q_0) \rho^W_1] d \quad (4.58)
\]

\[
E_1 \pi_2(h) = [(1 - q_1(h)) \rho^W_2(h) + q_1(h) \rho^T_2(h)] d \quad (4.59)
\]

The equilibrium is completed with the updating equations

\[
q_1(F) = \frac{(1 - \rho^T_1)q_0}{(1 - \rho^T_1)q_0 + (1 - \rho^W_1)(1 - q_0)} = 1 \quad (4.60)
\]

\[
q_1(D) = \frac{\rho^T_1 q_0}{\rho^T_1 q_0 + \rho^W_1 (1 - q_0)} = 0 \quad (4.61)
\]

and \( q_0 = \frac{1}{2} \) (uniform distribution for the prior belief that the authority is tough by the private sector). Using Equations (4.52) to (4.61), we can solve for the likelihood of a first-period devaluation, \( \rho^i_1, i = W, T \) as follows:

\[
\rho^T_1 = 1 - \frac{S + \phi \omega_T + k + \lambda - \phi \omega_T \phi}{4v}
\]

\[
\rho^W_1 = 1 - \frac{S + \phi \omega^W + k - \phi \omega_T \phi + \lambda}{4v}
\]

where

\[
\lambda = \frac{d}{4v} (2v - \beta \phi)(\theta^T - \theta^W) \geq 0 \quad (\therefore \theta^T \geq \theta^W)
\]

\[
\phi = 2\mu B v d g - Z
\]

\[
Z = \frac{s d^2 (\theta^T - \theta^W)}{2m(2v - dS)}
\]

\[
w^i = k - \frac{d}{4v} (\theta^T + \theta^W - 2\theta^i)
\]

\[
g = \frac{1}{2v - dS}
\]

\[
\eta = \frac{2v - dS}{2v - dS} \leq 1 \quad (\therefore m = s + B_{20})
\]
The expected devaluation in period 1 is then given by

\[ E_0 \pi_1 = \frac{d}{2} + \frac{kd - \theta T d^2 + \beta \phi w^T + \phi^T}{2v - d - \beta \mu B \eta \phi} \] 

(4.64)

As in Benigno and Missale (2001), it is shown in (4.64) that apart from the first two terms in the numerator capturing the first-period effects, the sign of \( E_0 \pi_1 \) depends on \( \phi = 2 \mu B \eta d - Z \). The term \( 2 \mu B \eta d \) captures the 'debt-burden' effect as \( \mu B \) is the level of inflation-sensitive debt, while the term \( Z \) represents the impact of the uncertainty over the authority's preferences which is determined by the difference, \( \theta T - \theta W \). Hence, the expectation of a first-period devaluation depends on whether the 'debt-burden' or the 'signalling' effect prevails: it is smaller when there is substantial uncertainty as to which type of authority is involved whereas it is greater when the debt level is high. Furthermore, we can show

\[
E_2^1 L_2(D) - E_2^1 L_2(F) = 2v[\rho_1^1(D) - \rho_2^1(F)][[d_m(\rho_1^2(D) + \rho_1^2(F)) + \Pi(D) + \Pi(F)]
\]

\[
= 2v[\rho_1^1(D) - \rho_2^1(F)][E_2^1 y_2(D) + E_2^1 y_2(F) - 2y^*] 
\]

(4.65)

where \( y_2(h) \) is second-period output following \( h = D, F \). Thus, the sign of the difference between the expected second-period loss from devaluation and that from the parity maintenance in period 1 depends on \( \rho_1^1(D) - \rho_2^1(F) \), which in turn, depends on whether the debt-burden effect or the signaling effect dominates.

**The Impact of Short-Term Debt**

As in Section 4.3.2, it is interesting to examine the role of debt maturity in determining the current and forward interest rates. Differentiating \( E_0 \pi_1 \) with respect to the short-term debt \( B_{10} \) while holding the total debt constant gives:

\[
\frac{\partial E_0 \pi_1}{\partial B_{10}} = \frac{\beta k g \phi d' + \beta B d \eta (\phi + \phi')(E_0 \pi_1 - \frac{d}{2})}{2v - d - \beta B \eta \phi} 
\]

(4.66)
where

\[
\phi' = 2Bvdg - Z' < \phi \\
Z' = \frac{vd(\theta^T - \theta^W)}{m(2v - dS)} > Z \quad (\because 2v > dS)
\]

As the sign of this derivative is determined by \(\phi = 2Bvdg - Z (\phi' = 2Bvdg - Z')\), the effect of debt maturity on the expected devaluation in the first period depends on the relative importance of 'debt-burden' effect to 'signalling' effect. When there is little uncertainty about the authority's type, so that \(Z = \frac{vd(\theta^T - \theta^W)}{m(2v - dS)} \) tends to zero, the short-term debt increases the probability, as perceived by the private sector, of a first-period devaluation, as it does under complete information. On the other hand, if the authority's resolve is uncertain—i.e., when \(\phi\) and \(\phi'\) are negative—the probability of a devaluation in period 1 decreases with short-term debt, as defending the exchange rate in adverse circumstances (larger short-term debt) sends a stronger signal of anti-inflation stance.

**Proposition 6.** In the case of discretion under asymmetric information, the debt maturity increases the probability of a first-period devaluation if the 'debt-burden' effect dominates, while when the 'signalling' effect prevails short-term debt enhances the credibility of the exchange rate regime. Both effects are magnified by the volume of inflation-sensitive debt, \(\mu B\).

The reason for a short maturity of debt to play a role when the 'signalling' effect prevails is that it may boost market confidence in the parity if it is maintained. In this case, an increase in short-maturity debt—i.e. a larger \(\mu B_{10}\)—would force the monetary authority to resist a possible crisis in order to avoid rolling over the debt at a higher-than-expected interest rate. This is because when the signaling effect dominates, a first-period devaluation would increase the expected second-period devaluation and thus the interest rate. However, when there is little uncertainty about the authority’s preferences, the probability of a first-period devaluation increases with short-term debt, as it does under complete information.
Hence, together with (4.57) we have

**Proposition 7.** The ‘signalling’ effect is crucial in determining the slope of the yield curve. If the uncertainty about the authority’s type is substantial, a higher ratio of short-term debt implies a lower short-term interest rate and a steeper yield curve as the forward rate keeps constant. On the other hand, when there is little uncertainty about authority’s preferences, a short maturity leads to a higher current interest rate and thus a flatter yield curve, as in the case of complete information.

### 4.4 Numerical Simulation

Simulation is conducted using Matlab to demonstrate the numerical results from our theoretical model for the three cases: commitment, discretion under complete information, and discretion under asymmetric information.

In our design, we set $v = 0.1$ so that output shock is uniformly distributed over the interval $[-0.1, 0.1]$. As we assume a fixed-size depreciation if monetary authority decides to devalue, the size $d$ is chosen as 0.2. The real value of total debt at period 0 $B = 0.075$, consisting of half one-period debt and half two-period debt. We choose $k = 0$ to eliminate other sources of credibility problems apart from those associated with debt accumulation. The proportion of inflation-sensitive debt is set at $\mu = 1$ so that all debt is deflatable. $\theta = 0.044$ is chosen to give $E_0(\pi_2) = 12.5\%$ on an annual basis. The discount factor is set at $\beta = (\frac{1}{1.05})^5$ based on a ‘period’ of 5 years.

The benefits gained from commitment is equivalent to the permanent percentage increases in output.

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4A variety of different set of parameters were also chosen to test sensitivity of our model, including $v = 0.5, d = 0.4, B = 0.25, k = 0, \mu = 0.8, \theta = 0.45$ so that $E_0(\pi_2) = 5.0\%, \beta = (\frac{1}{1.05})^5$. The simulation results confirm our findings about relationships between interested variables, although the magnitudes of intercepts and slopes may change from graph to graph.
4.4.1 Commitment versus Discretion under Complete Information

In Figure 4.2, the relationship between the ratio of short-term debt (in public debt) and the probability of a first-period crisis under commitment is compared with that under discretion and complete information. As shown in the theoretical model, the likelihood of having a crisis in the first period when the authority fully commits itself to a 2-period rule, is generally smaller than that when it pursues a discretionary policy under complete information. That is also known as ‘inflation bias’ when combining the contingent nature of economic policies with private sector’s perfect foresight.

When the authority’s preferences are publicly known, the probability of a first-period crisis increases with the ratio of short-term debt, as the authority tries to exploit the lower-than-expected interest rates following a devaluation. A first-period unexpected devaluation also reduce the real value of debt.

Figure 4.3 shows that in the case of commitment, the probability of a devalua-
tion in period 2 after maintaining the parity in the first period is higher than that following a devaluation, as long as there is some long-term debt. But the difference between the two probabilities decreases gradually as the composition of short-term debt increases, and could eventually diminish when all the debt is short-term.

Correspondingly, Figure 4.4 shows the situation when the monetary authority lacks a commitment mechanism but instead pursues discretionary policies under complete information. In contrast, the likelihood of a devaluation in period 2 after a successful defence is higher than that following an earlier devaluation, regardless of the composition of debt. If only the fundamentals matter, a current devaluation always eases the authority’s debt burden and thus increases the probability that a new parity will be successfully defended.

It is also shown from comparing Figure 4.3 with 4.4 that the difference of the two probabilities is greater in the case of discretion under complete information than in the case of commitment.

Yield curves under both commitment and discretion with complete information are depicted in Figure 4.5. In the case of discretion under complete information,
the expected inflation rate in period 1—formed in period 0—increases with the proportion of short-term debt. But the expectation of the second-period inflation rate does not depend on the maturity of debt. In other words, as $B_0$ increases, $E_0(\pi_1)$ goes up but $E_0(\pi_2)$ keeps constant, so $\frac{E_0(\pi_2)}{E_0(\pi_1)}$ goes down. Therefore, we would expect a flatter yield curve when the authority pursues a discretionary policy that is known to private sectors.

The substantial gain that the authority acquires when committing itself to a certain rule is shown in Figure 4.6. It is equivalent to a permanent output increase of around 7%. This is an example that the commitment rule is the first-best choice for the monetary authority when designing macroeconomic policies.

### 4.4.2 Discretion under Asymmetric Information

Figures 4.7 to 4.9 show us the numeric results—in particularly about the yield curves—when the authority pursues a discretionary policy that is not publicly known. It is known from our theory that the expectation of a second-period de-
Figure 4.5: Yield Curves with Commitment and Discretion

Figure 4.6: The Gain from Commitment
valuation that is formed in period 0 (hence the forward rate), does not depend on either the debt maturity or the authority's reputation. The credibility of the exchange rate regime in the long term only depends on fundamentals.

However in the short run, both tough and weak authorities have the incentive to defend the parity, in order to enhance their reputation and thus exploit the low interest rates. The larger the difference between the two types of authorities $\theta^T - \theta^W$, the stronger the incentive is for any given expected cost, $q_0 \theta^T + (1-q_0) \theta^W$. Therefore, an increase in uncertainty over the authority's preferences $(\theta^T - \theta^W)$ would lower the market expectation of an imminent alignment of exchange rate, $E_0 r_1$, which implies a steeper yield curve. Figure 4.7 shows this effect with the short-term debt fixed at 50% of total debt.

The effect of a shorter debt maturity on the expectation of a first-period devaluation is ambiguous. If the 'debt-burden' effect prevails over the 'signalling' effect, the likelihood of a current devaluation increases with the ratio of short-term debt, as it does under complete information. Therefore, we would expect a flatter yield
Figure 4.8: Yield Curves with Discretion and Asymmetric Information: High Debt Case ($B = 0.075, \theta^T - \theta^W = 0.044$); the ‘Debt-Burden’ Effect Dominates.

However, if the uncertainty over the authority’s type is substantial, a short debt maturity provides the authority with an incentive to resist a current crisis, in order to roll over the debt at lower-than-expected interest rates. As a result, the probability of a first-period devaluation decreases with the proportion of short-term debt. In this case, a steeper yield curve is shown in Figure 4.9.

4.5 The Variables and Data Set for Empirical Works

4.5.1 The Variables of Interest

We adopt the difference in interest rates between domestic market and US market to represent the probability of a current crisis as perceived by private sector. This measurement follows Drazen and Masson (1994) and BM. A wide interest rate spread implies an increased likelihood of an impending currency crisis, whereas a
narrow interest rate differential highlights confidence in the market that the parity will be maintained. An approximation of the yield curve is obtained by dividing the government long-term bond yield (a proxy for forward interest rates) by the market interest rate. The proportion of short-term debt is calculated by dividing the domestic debt with remaining maturity up to one year by the total domestic public debt.

A Market Pressure Index (MPI) is adopted to capture both large depreciations of local currency and failed attacks that are fended off by raising interest rates and/or selling international reserves. The continuous MPI is a weighted sum of the percentage change in the nominal exchange rate, the difference in the market interest rate, and the negative percentage change in the foreign exchange reserves\(^5\), according to their relative precisions.

As suggested by the relevant crisis theories and empirical works, three economic

\(^5\)As in Chapter 3, the minus sign ensures that a reduction in foreign reserves translates to an increase in the value of the MPI.
variables are also included in our model: (1) the Current Account Balance (% of GDP) that is the sign of the country's export competitiveness; (2) the Overall Budget Balance (% of GDP) to capture the authority's fiscal position; and (3) the International Reserves (minus gold, months of imports) always found significant to explain currency crises in many empirical works.

4.5.2 The Data Set

We obtain quarterly observations from 1993Q3 to 2003Q4 for 11 developing/emerging economies: Argentina, Brazil, the Czech Republic, Hungary, India, Malaysia, Mexico, Russia, South Africa, South Korea and Thailand.

The data for domestic public debt as well as the share of short-term debt comes from the Bank for International Settlements (BIS). Data of the remaining variables (budget balance, current account balance, etc.) come from the IMF's International Financial Statistics (IFS). The monthly exchange rate regime index (henceforth RR) is from Reinhart and Rogoff (2004). We convert this to a quarterly index. We use Datastream to extract quarterly data on long-term government bond yields and when it is not available, we interpolate yearly data to obtain quarterly observations. We could not obtain information on long-term government bond yields for Argentina, Brazil and India.

4.6 Differentiating Debt-Burden and Signalling Countries

In order to test the predictions made by our theory, we need to distinguish countries where the debt-burden effect prevails from those where the signalling effect dominates. We know from our theoretic model that a strenuous defence of the parity can signal the government's commitment and thus enhance credibility of the exchange rate regime, when there is substantial uncertainty about the authority's type; but with the debt-burden effect prevailing, this policy worsens fundamentals by increas-
ing the debt burden, making the economy more vulnerable to adverse shocks in the future.

As indicated in Proposition 5, countries where the debt burden is more important than signalling will find that a current devaluation increases the probability that a future defence of the new parity will succeed. In other words, we should observe a drop in the interest rate differential following an actual crisis (and/or a rise following a successful defence) in those countries. On the other hand, in the countries where there is substantial uncertainty over the authority’s resolve not to devalue, a depreciation in the current period would reveal a weak authority heading towards further devaluations. This generates inflationary expectations and a higher interest rate differential. Meanwhile, resisting a crisis enhances the reputation of the authority and thus the expectation that the parity will be maintained. Therefore, in signalling countries the interest rate spread rises after a crisis while it drops after maintaining the fixed rate.

Here, we figure out the type of each emerging economy, by looking at the patterns shown by its interest rate differentials following an actual devaluation and/or successful defence.

4.6.1 Identifying Actual Crises and Successful Defences

Before distinguishing between the two types of countries, it is necessary to identify the actual crises and successful defences experienced by them during the sample period.

We use the continuous Market Pressure Index (MPI) to gauge foreign exchange market pressures. The SETAR and MRS model are employed to define periods of severe attacks on exchange rate regimes. The estimated results of both models on 11 emerging economies are reported in Table 4.5–Table 4.8. Figure 4.22–Figure

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6Refer to Section 3.4 for more details.
7Refer to Section 3.5 for more details.
8Refer to Section 3.5 and Appendix A for more details.
9Note that we only regard a period as a period with severe pressure if both SETAR and Markov-switching model define it as such a period.
4.43 show both the filtered and smoothed probabilities of being in volatile state for all countries respectively.

The advantages of the non-linear state models is that they avoid imposing an arbitrary threshold levels \textit{a priori}—by allowing the data itself to reveal what move shows abnormal behaviour and hence signal a crisis. Moreover, the non-linear nature of both models is also appealing for the time series variables such as the MPI. Because of its single index restriction, the SETAR model seems to be slightly less effective in identifying periods of high pressure in foreign exchange markets, especially where there are frequent fluctuations in exchange markets.

The actual crisis episodes in the 11 emerging economies are identified by the regime changes in the Reinhart-Rogoff index. We then compare the actual crises (produced by the RR index) with the periods of high market pressure (defined by the SETAR and MRS model) for each country. It is shown in Table 4.1 that nearly all the actual devaluations are accommodated by periods of severe pressures in foreign exchange markets. This places confidence in the usage of the SETAR and MRS models.

A successful defence is then defined as a period when the high pressure in foreign exchange markets does not lead to a regime change in the RR index. In other words, the parity is maintained by the authority facing serious attacks at a great cost of loss of foreign reserves and high interest rates. Therefore, we can differentiate between the actual crises and the successful defences for the 11 selected countries\textsuperscript{10}.

Table 4.1: Actual Crises and Successful Defences

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\textsuperscript{10}Note that if a long period of more than one quarters, which is defined by both SETAR and MRS model as a high pressure period, includes one quarter that is identified by the RR index as an actual crisis, we regard the whole period as just one crisis incident. Furthermore, We only consider the sample size ranging from 1999Q3 to 2001Q4 as data for the RR index after 2001 is not available.
Table 4.1: Actual Crises and Successful Defences

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Table 4.1: Actual Crises and Successful Defences

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Notes: The actual crisis periods are defined by the Reinhart and Rogoff (1994) index. The estimated non-linear SETAR model has a regime-dependent intercept and slope parameters. The estimated non-linear Markov-switching model has a regime-dependent intercept and allows for heteroscedasticity. A period of severe market pressure on exchange rate is a period that is identified by both SETAR and Markov-switching models as such a period. A period of successful defence is therefore a period when there is serious market pressure but not a 'softening' of the exchange rate regime.

4.6.2 11 Emerging Economies

After making the distinction between the actual devaluations and the successful defences, we can now look at the situation facing each country following crises and/or successful defences, in order to identify its type as either debt-burden or signalling. The differentials of domestic market interest rates relative to USA is adopted to measure the probability of a currency crisis as perceived by the market.

1. Argentina

The interest rate spread in Argentina (INTDIF) is shown in Figure 4.10 alongside with the actual crises (identified by the RR index, XRINDEX) and periods of severe attacks (defined by both the SETAR and MRS model—CRISISSETAR and CRISISMRS respectively).

It is shown in Table 4.1 that a currency/debt crisis breaks out in Argentina from 2001Q3 to 2002Q2, which sees the abandonment of its currency board at the end of 2001. Figure 4.10 shows that the interest rate differential drops dramatically
after the crisis, indicating that the actual devaluation eases the authority’s debt burden and lowers the market expectation of a future crisis. Therefore, Argentina is categorized as a country where the "debt-burden" effect prevails over the "signalling" effect.

A successful defence of Argentina’s peso can also be seen in 1995Q1. Surprisingly, the interest rate spread drops to its normal level after an increase under the pressure. An explanation for that is the ‘currency board’ that Argentina adopts following the periods of turmoil in the early 1990’s, works well, boosting the authority’s reputation after it is successfully defended. However, when there is little uncertainty about the authority’s type later on, the ever-growing public debt becomes a more dominating issue.

2. Brazil

The peak in the interest rate differential in Brazil appears in the early days of the sample (Figure 4.11), corresponding to an actual crisis in 1994Q3 (Table 4.1). The interest differential falls drastically afterwards, showing signs of improvement.
in the fundamentals that reduces the probability of a future forced devaluation.

Because the peak is so high that the rest of the graph shows little movements, we produce Figure 4.12 to only focus on the remainder of the period (1994Q3-2003Q4). It is shown that before running into a full-scale crisis in 1999Q1, Brazil manages to defend its exchange rate regime in 1998Q3. However, this successful defence does not enhance the credibility of the regime, but instead by refraining the authority from inflationary financing its debts, leads to a forced devaluation several months later. It is then the actual devaluation that finally pulls the interest rate differential back to its normal level.

The parity is successfully defended again in 2000Q2, but the results are the same: the interest rate spread soon goes up to a new height in 2001Q4 and even higher in 2003Q2. Overall, Brazil is considered as a debt-burden country, where the public debt plays a prevailing role.

3. Czech Republic

From Table 4.1, it is seen that there are two major crises in Czech Republic
Figure 4.12: Brazil's Interest Rate Differential relative to USA 1994Q3-2003Q4
Figure 4.13: Czech Republic's Interest Rate Differential relative to USA
Figure 4.14: Hungary’s Interest Rate Differential relative to USA

in the last decade. Different from Argentina and Brazil, the Czech Republic’s first devaluation at the beginning of 1996 leads to the second, more full-grown crisis in 1997Q2. In other words, the interest rate spread keeps rising after 1996Q1’s crisis and reaches its peak in 1997Q2, as shown in Figure 4.13. This shows that Czech Republic is a ‘signalling’ country where a devaluation reveals a weak authority and thus increases the expectation of a further crisis.

4. Hungary

The interest rate differential in Hungary shows the same pattern as in Czech Republic (Figure 4.14).

An actual crisis occurs in Hungary in the early days of our sample (1994Q2), which pushes the interest spread to an even higher point—over 25% in 1995Q2 up from 20%. On the other hand, a series of successful defences conducted by the authority—starting from 1994Q4-1995Q2, 1996Q1, 1996Q4-1997Q3 to 2000Q2—
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dampens market speculations concerning a possible realignment, and eventually pushes the interest rate differential to its lowest level since the start of our sample. Thus, Hungary is regarded as a signalling country, where resisting a currency crisis and bearing the cost of servicing public debt sends a strong signal of the authority’s resolve against devaluation (inflation), and hence enhances the credibility of the exchange rate regime.

5. India

Shown in Figure 4.15, India’s interest rate differential has a general shape of the letter ‘V’. The big drop on the left hand side starts months after a crisis in 1995Q4. On the right hand side of the graph, a successful defence pushes the differential up from the bottom. They are signs that an early crisis eases the debt burden for India and leads to a fall in market speculation, while the country’s debt condition deteriorates after a successful defence and the market expects a future forced devaluation. Therefore, India is a debt-burden country according to our theory.\(^\text{11}\)

6. Malaysia

In Table 4.1, we can see that Malaysia experiences a period of turmoil during 1997Q2-1998Q1, when it decides to drop the crawling band and adopt a freely floating exchange rate regime (1997Q3). Figure 4.16 shows that this decision pushes the interest rate differential to an even higher level until its peak in 1998Q2. In contrast, a successful defence in 1995Q4 keeps the interest spread low throughout the whole year of 1996. It is therefore shown that Malaysia is a country where the ‘signalling’ effect prevails over the ‘debt-burden’ effect.

7. Mexico

Mexico’s situation in the last ten years is simple. In Figure 4.17, we see that the interest rate spread drops dramatically after it reaches its highest point—when Mexico abandons its crawling peg and the Mexican peso sharply depreciates against

\[^{11}\text{There are another two successful defences carried out by the Indian authority: one in 1998Q1 that leads to an initial small drop in the differential, but together with another one in 1998Q3 which raises the spread to its new height in 1998Q4.}\]
Figure 4.15: India’s Interest Rate Differential relative to USA
Figure 4.16: Malaysia's Interest Rate Differential relative to USA

Figure 4.17: Mexico's Interest Rate Differential relative to USA
US Dollar in 1996Q1 (4.1). Mexico is thus considered as a country where a current devaluation improves the fundamentals and makes a future crisis less likely.

8. Russia

Russia's interest rate differential relative to USA is shown in Figure 4.18.

We can see that the spread initially goes down until 1997Q3-1997Q4 when there is a successful defence against speculative attacks. The defence, on the other hand, starts to push the spread up and to its peak one year later when a full-scale currency crisis occurs (1998Q3-Q4). The Russian market seems to settle down after the 1998 Crisis as the interest rate differential falls dramatically to its normal level in 1999Q2. That is exactly what we would expect from a 'debt-burden' country—resisting a crisis worsens the fundamentals and thus increases the likelihood of a future devaluation, whereas a devaluation dampens market speculations concerning a further devaluation. Hence, Russia is a country with a significant 'debt-burden' effect. 

\[\text{Figure 4.18: Russia's Interest Rate Differential relative to USA}\]

\[\text{Another successful defence is recorded in 1999Q3 when the interest rate differential randomizes}\]
9. South Africa

As the SETAR model produces too many unnecessary high pressure periods for South Africa, we here only consider the MRS model that identifies 1996Q2, 1998Q3 and 2001Q4 as periods when the country's currency is facing serious attacks (Table 4.1).

It is shown in Figure 4.19 that the initial successful defence of the exchange rate in 1996Q2, although initially keeping the interest rate differential in the rising trend in the first two quarters, draws the spread down to a low level in 1998Q1. This pattern is more clear when we analyze the second successful defence in 1998Q3, when the interest rate differential drops drastically from its peak of nearly 15% to around 3%. It seems that the market takes the reputation of the authority seriously into its account\(^\text{13}\). Therefore, South Africa is categorized as a country where the

---

\(^{13}\) Another successful defence is identified in 2001Q4-2002Q1, when the interest spread continues
'signalling' effect dominates.

10. South Korea

It is shown in Figure 4.20 that in 1994Q3, a speculative attack is successfully warded off by South Korean authority, which draws the interest rate differential down from its peak.

The figure also sees another two separate attacks that are fended off in 1996Q3 and 1997Q1. The result of the first successful defence is a lower interest spread, and the second defence also holds the market expectation of a realignment down. On the other hand, failing to resist the pressure in exchange markets and choosing to devalue its currency in 1997Q4 drags South Korea into deep chaos—the interest rate differential shoots to a record high of 18.26 after the reputation of the authority is rising. This could be seen as an early sign that the market starts to lose its confidence on the authority, since South Africa's economy has been struggling to thrive in the early 2000's. But it does not affect the main picture of South Africa.
ruined. Thus, there is little doubt that South Korea is a 'signalling' country.

11. Thailand

As shown in Figure 4.21, the interest rate differential in the Thai market is
pushed to 15% in 1998Q2, even higher than 13.6% during the 1997 crises—rather
than drawn down to its normal level as expected by many people. In other words,
the Thai crisis in 1997Q2-Q3 is not the end of the story, but instead the beginning
of it—a devaluation reveals a weak authority, generates inflationary expectation and
thus even higher interest rates.

Overall, this procedure places Argentina, Brazil, India, Mexico and Russia in
the debt-burden group and the Czech Republic, Hungary, Malaysia, South Africa,
South Korea and Thailand in the signalling group.

Figure 4.21: Thailand's Interest Rate Differential relative to USA
4.7 Tests of Two Predictions

4.7.1 Two Predictions

It is shown in our theoretical model that if the domestic currency is pegged to a foreign currency, defending the peg with higher interest rates may be too expensive, especially when there is a large share of short-term debt needed to be rolled over. Instead, the authority may decide to devalue in order to improve the fundamentals (through reducing the real debt burden), and thus increase the likelihood that a future exchange rate regime will be maintained.

This logic, however, ignores the 'signalling' effect; in other words, it neglects to take into account the effects that the decision to devalue may reveal a weak authority and hence lead to a further devaluation—if there is substantial uncertainty about the authority’s preferences in market.

This forms our first testable prediction. When there is little uncertainty over the authority’s preferences and the level of debt is high, a devaluation leads to a lower interest rate. The greater the amount of short-term debt that is to be rolled over, the greater the incentive to devalue and exploit the lower rate. Thus the likelihood of a first-period devaluation increases with the share of short-term debt. On the other hand, when the uncertainty about the authority type is substantial, it is the no-devaluation decision that produces a lower interest rate. In this case, the probability of a first-period devaluation decreases with the proportion of short-term debt, because defending the exchange rate sends a strong signal of a tough authority, especially in adverse circumstances such as a high ratio of short-term debt. Therefore, the effect of the debt maturity on the probability of a current devaluation depends on whether the signaling or the debt-burden effect prevails (see Proposition 6).

Equation (4.57) shows that the expected second-period devaluation formed in period 0 is independent of the debt maturity. In other words, the short-term debt does not affect the forward interest rate, which only depends on fundamentals.
Together with the first prediction, we can draw the term structure of interest rates for the debt-burden countries as well as the signalling countries (Proposition 7) as our second prediction. The reputational incentive lowers the short-term interest rate but has no impact on the forward rate. Therefore, when there is substantial uncertainty about the authority's preferences, a higher ratio of short-term debt implies lower current interest rates and a steeper yield curve. In contrast, if the fundamentals outweigh the reputational considerations, a higher proportion of short-term debt is associated with higher current interest rates and a flatter (or more downward-sloping) yield curve.

4.7.2 The Methodology

The methodology we adopt to test the above two predictions is the generalised method of moments (GMM)\(^\text{14}\) for a dynamic panel data model (see Arellano and Bond (1991)). The model we estimate has the form

\[
y_{it} = \alpha y_{i(t-1)} + \beta'(L)x_{it} + \lambda_i + \eta_i + v_{it}, \quad t = q + 1, \ldots, T_i; \quad i = 1, \ldots, N. \tag{4.67}
\]

where \(y_{it}\) is the dependent variable that is either the market interest rate differential relative to the USA or the yield curve, \(x_{it}\) is a vector of explanatory variables, \(\beta'(L)\) is a vector of associated polynomials in the lag operator, and \(\eta_i\) and \(\lambda_i\) are the individual- and time-specific effect respectively. The number of time periods available on the \(i\)th individual is \(T_i\) and the number of individuals is \(N\).

We include four lags of the independent variables in the original regression as the data we use is quarterly. One period lagged dependent variable is also included in the RHS of the model to capture all the other effects that are not explicitly considered. As Arellano and Bond (1991) argue, when the \(v_{it}\) are heteroscedastic, simulations suggest that the asymptotic standard errors for the two-step estimators can be a poor guide for testing hypothesis in a typical sample size. In our case, both the

\(^{14}\)Refer to the Appendix C for model details.
interest rate differential and yield curve appear to be more volatile during currency crises than in relatively stable periods. Therefore, we report on the one-step GMM estimators with the asymptotic heteroscedasticity-consistent standard errors, based on which the inference tests are more reliable.15

The assumption of no serial correlation (or only moving average) in the disturbance term $v_{it}$ is essential for the consistency of our one-step GMM first-difference estimators, as the lagged dependent variable in the model is instrumented by itself with further lags. The Sargan test of over-identifying restrictions is performed, and the test statistics that are based on the two-step GMM estimators and are heteroscedasticity-consistent are also reported16.

### 4.7.3 The Results

**More short-term debt—a heavy burden or a good incentive for the authority to defend the parity in short run?**

In each of the two groups of countries, we pool all the observations across countries and time periods, and regress the interest rate differential on its one-period lag and four explanatory variables—short-term debt ratio, current account balance, budget balance and international reserves. The results for both the debt-burden and signalling countries are shown in Table 4.2.

The estimated results show that in countries where the debt issue plays a dominating role over uncertainty on the authority’s preferences, the probability of a current crisis increases with the ratio of short-term debt in total public debt. The 2-, 3-, 4-period lagged short-term debt ratio variables are statistically significant at the 10% level with positive and negative signs; the sum of the three significant parameters ($62.44, -48.24$ and $42.41$) is $56.61$ indicating an overall long-term positive relationship. For the signalling countries, the relationship between the probability

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15See Blundell and Bond (1998) for more details.
16Again, see Arellano and Bond (1991) for comprehensive discussions on several different test procedures.
Table 4.2: Effect of Short-term Debt on Likelihood of Current Devaluation I

<table>
<thead>
<tr>
<th>Variable</th>
<th>Debt burden</th>
<th>Signalling</th>
</tr>
</thead>
<tbody>
<tr>
<td>DINTDIF(-1)</td>
<td>0.30***</td>
<td>0.50***</td>
</tr>
<tr>
<td>DSTDEBT</td>
<td>-1.17</td>
<td>-5.35*</td>
</tr>
<tr>
<td>DSTDEBT(-1)</td>
<td>-0.88</td>
<td>-6.93**</td>
</tr>
<tr>
<td>DSTDEBT(-2)</td>
<td>62.44***</td>
<td>-1.60</td>
</tr>
<tr>
<td>DSTDEBT(-3)</td>
<td>-48.24**</td>
<td>1.37</td>
</tr>
<tr>
<td>DSTDEBT(-4)</td>
<td>42.41**</td>
<td>-1.53</td>
</tr>
<tr>
<td>DCACC</td>
<td>0.82*</td>
<td>0.17**</td>
</tr>
<tr>
<td>DCACC(-1)</td>
<td>-1.34*</td>
<td>-0.002</td>
</tr>
<tr>
<td>DCACC(-2)</td>
<td>2.37***</td>
<td>-0.12</td>
</tr>
<tr>
<td>DCACC(-3)</td>
<td>-2.57***</td>
<td>-0.11</td>
</tr>
<tr>
<td>DCACC(-4)</td>
<td>1.06**</td>
<td>0.06</td>
</tr>
<tr>
<td>DBUDG</td>
<td>1.62</td>
<td>0.04</td>
</tr>
<tr>
<td>DBUDG(-1)</td>
<td>-2.37***</td>
<td>-0.03</td>
</tr>
<tr>
<td>DBUDG(-2)</td>
<td>2.17***</td>
<td>-0.02</td>
</tr>
<tr>
<td>DBUDG(-3)</td>
<td>0.70*</td>
<td>-0.04</td>
</tr>
<tr>
<td>DBUDG(-4)</td>
<td>-2.08***</td>
<td>-0.0003</td>
</tr>
<tr>
<td>DRES</td>
<td>-2.48</td>
<td>-3.91***</td>
</tr>
<tr>
<td>DRES(-1)</td>
<td>5.36</td>
<td>1.71</td>
</tr>
<tr>
<td>DRES(-2)</td>
<td>-8.19*</td>
<td>0.15</td>
</tr>
<tr>
<td>DRES(-3)</td>
<td>7.19***</td>
<td>-0.90</td>
</tr>
<tr>
<td>DRES(-4)</td>
<td>-2.40*</td>
<td>1.61*</td>
</tr>
<tr>
<td>Constant</td>
<td>-30.37**</td>
<td>-0.45</td>
</tr>
</tbody>
</table>

Wald test (joint) 1.008e + 011***
Sargan test -1.395e - 013 2.875e - 016

Estimation Method: GMM for dynamic panel data. All results are obtained using the DPD package by Doornik, Arellano and Bond for Ox. The transformation we adopt is the first-difference of the level equations. The time dummies are also included in both models. Debt-burden countries include Argentina, Brazil, India, Mexico, Russia. Signaling countries include the Czech Republic, Hungary, Malaysia, South Africa, South Korea and Thailand. INTDIF, the interest rate differential with the USA is the dependent variable. 'D' indicates first-difference. The number in bracket is the lag of the variable. Parameters with *** are significant at 1% level, parameters with ** are significant at 5% level, and parameters with * are significant at 1% level. Estimators are one-step GMM estimators, with p-values based on their asymptotic heteroscedasticity-consistent standard errors. The Sargan test statistics are based on the two-step GMM estimators and are heteroscedasticity-consistent. STDEBT -Ratio of Short-term Domestic Public Debt to Total Domestic Public Debt, CACC -Current Account (% of GDP), BUDG -Overall Budget Balance (% of GDP), RES -International Reserves (months of imports).
of an impending crisis and the ratio of short-term debt is negative. In Table 4.2, the current and 1-period lagged short-term debt ratio variables have statistically significant and negative coefficients (−5.35 and −6.93 respectively).

Therefore, the first prediction by our theoretical model is supported: in countries where the ‘debt-burden’ effect dominates, a bigger proportion of short-term debt leads to a higher market expectation of a devaluation, as the incentive to devalue and improve fundamentals is greater; while in countries where the ‘signalling’ effect prevails, a bigger short-term debt ratio lowers the probability of a crisis, as resisting great pressure not to devalue strengthens the authority’s stance.

The one-period lagged dependent variables in both cases are statistically significant even at 1% level, which provides support for our usage of a dynamic setting. All other explanatory variables also have at least one significant coefficient in our regressions. Wald tests show that all the explanatory variables are jointly significant at even 1% level in the regression for debt-burden dominating countries, although not for signalling countries. Sargan tests show that the one-step GMM-estimators in our results are asymptotically consistent in the sense that it is valid to instrument the lagged dependent variable with its further lags.

In order to test the robustness of our estimated results, we drop the variables whose coefficients are not statistically significant at 10% level from our original regressions. The new results are shown in Table 4.3.

The results are very similar. The 2-, 3-, and 4-period lagged value of short-term debt ratio are statistically significant and have a sum of 52.20 in the ‘debt-burden’ group, which means that the probability of a current devaluation increases with the proportion of short-term debt in those countries. However in ‘signalling’ countries, the statistically significant variables of the short-term debt ratio are its current value and one-period lag (with coefficients of −4.33 and −5.07 respectively). This shows that an increase in the short-term debt by 1% of total public debt lowers the interest rate differential (likelihood of a current crisis) by 9.40%.

The one-period lagged dependent variables are still statistically significant at the
### Table 4.3: Effect of Short-term Debt on Likelihood of Current Devaluation II

<table>
<thead>
<tr>
<th>Variable</th>
<th>Debt burden</th>
<th>Signalling</th>
</tr>
</thead>
<tbody>
<tr>
<td>DINTDIF(-1)</td>
<td>0.27***</td>
<td>0.63***</td>
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<tr>
<td>DSTDEBT</td>
<td></td>
<td>-4.33**</td>
</tr>
<tr>
<td>DSTDEBT(-1)</td>
<td></td>
<td>-5.07**</td>
</tr>
<tr>
<td>DSTDEBT(-2)</td>
<td>68.38***</td>
<td></td>
</tr>
<tr>
<td>DSTDEBT(-3)</td>
<td>-57.96***</td>
<td></td>
</tr>
<tr>
<td>DSTDEBT(-4)</td>
<td>41.78**</td>
<td></td>
</tr>
<tr>
<td>DCACC</td>
<td></td>
<td>0.06*</td>
</tr>
<tr>
<td>DCACC(-1)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>DCACC(-2)</td>
<td>1.81***</td>
<td></td>
</tr>
<tr>
<td>DCACC(-3)</td>
<td>-2.38***</td>
<td></td>
</tr>
<tr>
<td>DCACC(-4)</td>
<td>1.20***</td>
<td></td>
</tr>
<tr>
<td>DBUDG</td>
<td></td>
<td></td>
</tr>
<tr>
<td>DBUDG(-1)</td>
<td>-1.67***</td>
<td></td>
</tr>
<tr>
<td>DBUDG(-2)</td>
<td>2.39***</td>
<td></td>
</tr>
<tr>
<td>DBUDG(-3)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>DBUDG(-4)</td>
<td>-1.43***</td>
<td></td>
</tr>
<tr>
<td>DRES</td>
<td></td>
<td>-2.01***</td>
</tr>
<tr>
<td>DRES(-1)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>DRES(-2)</td>
<td>-4.76***</td>
<td></td>
</tr>
<tr>
<td>DRES(-3)</td>
<td>6.20***</td>
<td></td>
</tr>
<tr>
<td>DRES(-4)</td>
<td>-2.78*</td>
<td>1.35***</td>
</tr>
<tr>
<td>Constant</td>
<td>-30.83**</td>
<td>-0.39</td>
</tr>
<tr>
<td>Wald test(joint)</td>
<td>9.888e + 009***</td>
<td>139.7***</td>
</tr>
<tr>
<td>Sargan test</td>
<td>-1.637e − 013</td>
<td>-4.580e − 016</td>
</tr>
</tbody>
</table>

*Notes: Same as in Table 4.2.*
1% level in both cases. The Wald test for joint significance of all explanatory variables turns out to be statistically significant even at 1% level for both the signalling group and the debt-burden group. Sargan tests show no signs of serial correlation in the disturbance terms in both regressions.

**Will the emerging economies all face a favourable term structure of interest rates—a steep yield curve?**

We have shown that developing countries facing substantial uncertainty about their authorities’ types, should reap lower short-term interest rates as they strengthen their anti-inflation stances with a shorter debt maturity. Because the long-term rate is only determined by fundamentals, these countries face a steeper yield curve as a reward for their willing to fight against speculative attacks in adverse circumstances. It is also shown that in countries with a high level of debt, a larger amount of short-term debt always pushes short-term interest rates higher as the incentive to roll it over at a lower-than-expected rate is greater. As a result of the fact that the debt maturity does not affect the forward rate, these countries face a flatter yield curve.

As already mentioned, it has not been possible to find data for Argentina, Brazil and India and we can only run regressions for the signalling group. As before, we pool all the available data across both countries and time periods, and estimate a dynamic panel data model by GMM linking the yield curve (calculated by dividing the government long term bond yield by current market interest rate) to the short-term debt ratio and other 3 explanatory variables. The estimate results are shown in Table 4.4.

It is shown that coefficients of the current period and 1-, 3-, 4-period lagged short-term debt ratio variable are statistically significant at the 10% level. But the positive parameters (2.74 for 1st lag and 1.09 for 3rd lag) outweigh their negative counterparts (−0.83 for current value and −1.87 for 4th lag). In other words, a one percentage point increase in the proportion of short-term debt to total public debt raises the long-term to short-term interest rate ratio by more than one percent.
### Table 4.4: Effect of Short-term Debt on Yield Curve

<table>
<thead>
<tr>
<th>Variable</th>
<th>Signalling I</th>
<th>Signalling II</th>
</tr>
</thead>
<tbody>
<tr>
<td>DYC(-1)</td>
<td>0.78***</td>
<td>0.78***</td>
</tr>
<tr>
<td>DSTDEBT</td>
<td>-0.83*</td>
<td>-0.84**</td>
</tr>
<tr>
<td>DSTDEBT(-1)</td>
<td>2.74***</td>
<td>2.45***</td>
</tr>
<tr>
<td>DSTDEBT(-2)</td>
<td>-0.49</td>
<td>-</td>
</tr>
<tr>
<td>DSTDEBT(-3)</td>
<td>1.06***</td>
<td>1.15***</td>
</tr>
<tr>
<td>DSTDEBT(-4)</td>
<td>-1.87***</td>
<td>-2.04***</td>
</tr>
<tr>
<td>DCACC</td>
<td>-0.01**</td>
<td>-</td>
</tr>
<tr>
<td>DCACC(-1)</td>
<td>0.008</td>
<td>-</td>
</tr>
<tr>
<td>DCACC(-2)</td>
<td>0.02***</td>
<td>0.02***</td>
</tr>
<tr>
<td>DCACC(-3)</td>
<td>0.02**</td>
<td>-</td>
</tr>
<tr>
<td>DCACC(-4)</td>
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<td>-</td>
</tr>
<tr>
<td>DBUDG</td>
<td>0.005</td>
<td>-</td>
</tr>
<tr>
<td>DBUDG(-1)</td>
<td>0.005</td>
<td>-</td>
</tr>
<tr>
<td>DBUDG(-2)</td>
<td>0.008*</td>
<td>0.005*</td>
</tr>
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<td>DBUDG(-3)</td>
<td>0.002***</td>
<td>-</td>
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<td>DBUDG(-4)</td>
<td>-0.002</td>
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</tr>
<tr>
<td>DRES</td>
<td>0.19*</td>
<td>-</td>
</tr>
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<td>DRES(-1)</td>
<td>0.01</td>
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<td>DRES(-2)</td>
<td>-0.36</td>
<td>-</td>
</tr>
<tr>
<td>DRES(-3)</td>
<td>0.35**</td>
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<tr>
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<tr>
<td>Constant</td>
<td>0.25</td>
<td>0.29</td>
</tr>
</tbody>
</table>

Wald test (joint) 13.39 835.4***
Sargan test 4.320e - 017 2.651e - 016

Notes: The second column shows primary results, the third column shows results after dropping variables whose coefficients are statistically insignificant. YC – the government long-term yield over the market interest rate.
Therefore, the yield curve in countries where the 'signalling' effect prevails over the 'debt-burden' effect, becomes steeper when the short-term debt ratio increases. The second prediction made by our theory is proven.

Dropping the variables whose coefficients are not statistically significant at the 10% level does not alter the results. The current value and 1-, 3-, 4-period lag of the short-term debt ratio are still significant with a sum of the respective parameters equal 0.72. This shows the positive relationship between the yield curve and the proportion of short-term debt in 'signalling' countries, as we predict.

The one-period lagged dependent variable are both significant even at the 1% level in two regressions. This justifies our adoption of the dynamic panel data model. The Wald test statistic, although not significant in the initial regression, becomes significant after we drop the insignificant variables. The Sargan test based on two-step GMM estimators also shows that there is no serial correlation in the error term in either regression. In other words, our one-step GMM-estimators are asymptotically consistent.

4.8 Conclusions

It is shown in this chapter that a successful defence of the parity may signal the authority's anti-inflation resolve and thus enhance the credibility of an exchange rate regime, or on the other hand, it may increase the debt burden for the authority and make the economy more vulnerable to adverse shocks in the future. Which effect a successful defence against speculative attacks may have on the credibility of exchange rate regime is therefore uncertain—depends on the importance of public debt relative to authority's reputation incentive.

We adopt a three-period stochastic version of the Barro and Gordon (1983) model, where the decision to devalue the currency creates unexpected inflation, which increases output both through a standard price-output effect and through reduction of the distortionary taxes. The probability of a devaluation in each period is derived by solving the optimal problem facing the authority.
Three cases are studied. The first case where the authority can commit itself to a public-known 2-period rule (fixed $\rho_1$ and $\rho_2$) produces the socially optimal solution regarded as the benchmark. The second case where the authority pursues discretionary policies under complete information, shows that resisting a currency crisis and refraining from inflationary financing could worsen the fundamentals and thus increase the need for future devaluation. But a devaluation, on the other hand, reduces the real debt burden and hence improves the prospects that the new exchange rate regime will be maintained.

However, the 'signalling' effect can not be ignored when determining the impact of current actions on future policy, if the uncertainty over the authority's preferences is substantial. As a result, defending the parity could send a strong signal of the authority's type and thus enhance the credibility of the regime, while a devaluation reveals a weak authority and leads to higher expectations of a further devaluation. Which effect prevails depends on the relative importance of the fundamentals and the monetary authority's reputation. The results from our theoretical model are demonstrated by the numerical simulation using Matlab.

A central focus of our study is on the effect of debt, especially debt maturity on the term structure of interest rates in the two types of countries. In countries where the debt-burden effect dominates, a larger proportion of short-term debt strengthens the incentive to roll it over at a lower-than-expected interest rate, and thus increases the probability of a current devaluation and leads to a higher short-term interest rate. As the forward rate is independent of debt maturity, we then expect a flatter yield curve. But in countries where there is substantial uncertainty about the authority's type, the intention to send a strong signal of authority's anti-inflation stance when facing a high ratio of short-term debt, may imply a lower current interest rate. Together with a forward rate that only depends on fundamentals, this leads to a steeper yield curve.

In our empirical work, the 'debt-burden' countries are first distinguished from 'signalling' countries by analyzing the different patterns shown by their interest
rates after actual devaluations and/or successful defences. The actual crises are defined by the regime changes in the Reinhart-Rogoff index. Both the SETAR (Self Exciting Threshold Autoregression) model and the Markov Regime Switching model are adopted to define the periods when there is severe pressure in the exchange markets. The periods of successful defences are then identified as periods when the serious market pressure does not lead to actual devaluations.

The two predictions made by our theoretical model have then been tested in 11 emerging countries by using GMM (General Method of Moments) for the dynamic panel data model, except for the yield curve in the debt-burden countries where data is not available. The likelihood of a current crisis is represented by the current market interest rate differential relative to USA, while the yield curve is calculated by dividing the government long term bond yield (measurement of forward rate) by the market interest rate. Besides the short-term debt ratio, another three explanatory variables are selected to include in our regressions. The estimated results confirm the two predictions made by our theoretical model.

There are still some factors in currency crises which we have not examined under the issue of credibility. For example alongside unemployment and public debt, domestic bank credit may be another channel through which current actions would affect future policies. The experience of East Asian countries in the last two decades shows that defending the fixed rate while enhancing authority's anti-inflation stance, may send a wrong signal to the financial market to encourage aggressive and risky lending and thus making the economy more vulnerable to adverse shocks in the future. This is a task left for future research.


Table 4.5: SETAR Model \(^a\) Estimated Results \(^b\) for 11 Emerging Economies (1)

<table>
<thead>
<tr>
<th>Countries</th>
<th>Argentina</th>
<th>Brazil</th>
<th>Czech Republic</th>
<th>Hungary</th>
<th>India</th>
<th>Malaysia</th>
</tr>
</thead>
<tbody>
<tr>
<td>Intercept (Tranquil)</td>
<td>-2.55 (^c)</td>
<td>-5.10</td>
<td>-2.63</td>
<td>-1.06</td>
<td>-0.55</td>
<td>-0.39</td>
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<tr>
<td></td>
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<td>(-2.68)</td>
<td>(-6.33)</td>
<td>(-3.26)</td>
<td>(-5.41)</td>
<td>(-2.25)</td>
</tr>
<tr>
<td>Intercept (Volatile)</td>
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<td>24.15</td>
<td>2.45</td>
<td>1.30</td>
<td>0.66</td>
<td>1.35</td>
</tr>
<tr>
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<td>(12.14)</td>
<td>(3.51)</td>
<td>(10.06)</td>
<td>(4.03)</td>
<td>(9.53)</td>
</tr>
<tr>
<td>Coefficients on AR(1) (Tranquil)</td>
<td>-0.06</td>
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<td>0.27</td>
<td>0.25</td>
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<td>0.23</td>
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<td>(0.86)</td>
<td>(1.70)</td>
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<tr>
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<td></td>
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</tr>
<tr>
<td>Coefficients on AR(3) (Tranquil)</td>
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<td></td>
<td>(1.48)</td>
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<td>(0.85)</td>
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<td>(-0.93)</td>
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<tr>
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<tr>
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<td>3.04</td>
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<tr>
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<td>31.51</td>
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<td>31.04</td>
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</tr>
</tbody>
</table>

\(^a\) Estimated Model: Non-linear SETAR (Self Exciting Threshold Autoregression) with a regime-dependent intercept and slope parameter.

\(^b\) All results are obtained using H.-M. Krolzig’s MSVAR package for Ox.

\(^c\) Number in parenthesis below regression coefficient is t-value.

\(^d\) Number in square bracket below test statistic is p-value.
Table 4.6: SETAR Model\(^a\) Estimated Results\(^b\) for 11 Emerging Economies (2)

<table>
<thead>
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<th>Countries</th>
<th>Mexico</th>
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<th>South Africa</th>
<th>South Korea</th>
<th>Thailand</th>
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<td>(-4.00)</td>
</tr>
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<td>(3.45)</td>
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<td>(3.22)</td>
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<td>(1.95)</td>
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<tr>
<td>(Tranquil)</td>
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</tr>
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<td>(2.69)</td>
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<td>(1.91)</td>
</tr>
<tr>
<td>(Volatile)</td>
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<td>-0.97</td>
<td>-0.37</td>
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<td>-1.20</td>
</tr>
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<td>(-1.20)</td>
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<td>(-1.70)</td>
</tr>
<tr>
<td>(Volatile)</td>
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<td>0.66</td>
<td></td>
<td>0.66</td>
<td>0.66</td>
</tr>
<tr>
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<td>(1.53)</td>
<td>(1.53)</td>
</tr>
<tr>
<td>Coefficients on AR(4)</td>
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<td></td>
<td></td>
</tr>
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<td></td>
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</tr>
<tr>
<td>(Volatile)</td>
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<td></td>
</tr>
<tr>
<td>Std. Dev. (Tranquil)</td>
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<td>1.03</td>
<td>1.87</td>
<td>1.78</td>
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<td>Std. Dev. (Volatile)</td>
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<td>9.68</td>
<td>1.74</td>
<td>3.07</td>
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</tr>
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<td>0.87</td>
<td>0.75</td>
<td>0.95</td>
</tr>
<tr>
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<td>7.63</td>
<td>4.25</td>
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<td>LR Linearity test(^d)</td>
<td>46.99</td>
<td>43.34</td>
<td>37.76</td>
<td>28.40</td>
<td>28.75</td>
</tr>
</tbody>
</table>

\(^a\) Estimated Model: Non-linear SETAR (Self Exciting Threshold Autoregression) with a regime-dependent intercept and slope parameter.

\(^b\) All results are obtained using R-M. Krolzig's MSVAR package for Ox.

\(^c\) Number in parenthesis below regression coefficient is t-value.

\(^d\) Number in square bracket below test statistic is p-value.
Table 4.7: Markov-switching Model \(^{a}\) Estimated Results \(^{b}\) for 11 Emerging Economies (1)

<table>
<thead>
<tr>
<th>Countries</th>
<th>Argentina</th>
<th>Brazil</th>
<th>Czech Republic</th>
<th>Hungary</th>
<th>India</th>
<th>Malaysia</th>
</tr>
</thead>
<tbody>
<tr>
<td>Intercept (Tranquil)</td>
<td>-0.14  (^{(c)})</td>
<td>-17.71</td>
<td>-0.67</td>
<td>-4.77</td>
<td>-0.24</td>
<td>-0.62</td>
</tr>
<tr>
<td></td>
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<td>(-8.48)</td>
<td>(-1.50)</td>
<td>(-5.03)</td>
<td>(-2.87)</td>
<td>(-2.15)</td>
</tr>
<tr>
<td>Intercept (Volatile)</td>
<td>1.15  (^{(c)})</td>
<td>22.54</td>
<td>-0.52</td>
<td>1.60</td>
<td>0.41</td>
<td>1.66</td>
</tr>
<tr>
<td></td>
<td>(0.18)</td>
<td>(2.42)</td>
<td>(0.17)</td>
<td>(10.76)</td>
<td>(0.75)</td>
<td>(15.41)</td>
</tr>
<tr>
<td>AR(1)</td>
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<td>-0.53</td>
<td>0.08</td>
<td>0.25</td>
<td>0.19</td>
<td>0.51</td>
</tr>
<tr>
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<td>(-0.28)</td>
<td>(-0.28)</td>
<td>(0.06)</td>
<td>(4.07)</td>
<td>(1.98)</td>
<td>(6.62)</td>
</tr>
<tr>
<td>AR(2)</td>
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<td>(0.44)</td>
<td>(-2.35)</td>
<td>(2.44)</td>
<td>(2.44)</td>
</tr>
<tr>
<td>AR(3)</td>
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<td>-0.44</td>
<td>-0.44</td>
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<td>(-1.50)</td>
<td>(0.39)</td>
<td>(1.44)</td>
<td>(-3.96)</td>
<td>(-3.96)</td>
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<tr>
<td>AR(4)</td>
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<td>-0.09</td>
<td>-0.09</td>
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<tr>
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<td>(0.11)</td>
<td>(-0.61)</td>
<td>(-0.58)</td>
<td>(-0.58)</td>
</tr>
<tr>
<td>Std. Dev. (Tranquil)</td>
<td>2.26</td>
<td>4.41</td>
<td>2.13</td>
<td>1.25</td>
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<td>0.55</td>
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<tr>
<td>Std. Dev. (Volatile)</td>
<td>18.80</td>
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<td>(Tranquil → Tranquil)</td>
<td>[0.000]</td>
<td>[0.004]</td>
<td>[0.013]</td>
<td>[0.000]</td>
<td>[0.002]</td>
<td>[0.000]</td>
</tr>
<tr>
<td>(Volatile → Tranquil)</td>
<td>0.77  (^{(d)})</td>
<td>0.44</td>
<td>0.01</td>
<td>0.60</td>
<td>0.05</td>
<td>0.85</td>
</tr>
<tr>
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<td>[0.013]</td>
<td>[0.000]</td>
<td>[0.000]</td>
<td>[0.002]</td>
<td>[0.000]</td>
</tr>
<tr>
<td>SC criterion</td>
<td>7.09</td>
<td>8.96</td>
<td>5.97</td>
<td>4.55</td>
<td>2.90</td>
<td>3.50</td>
</tr>
<tr>
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<td>43.48  (^{(d)})</td>
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<td>8.94</td>
<td>24.94</td>
<td>12.57</td>
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<td>[0.013]</td>
<td>[0.000]</td>
<td>[0.002]</td>
<td>[0.000]</td>
</tr>
</tbody>
</table>

\(^{a}\) Estimated Model: Non-linear Markov-switching Autoregressive (MS-AR) model with a regime-dependent intercept and allowing for heteroscedasticity in the disturbance term.

\(^{b}\) All results are obtained using H-M. Krolzig's MSVAR package for Ox.

\(^{c}\) Number in parenthesis below regression coefficient is \(t\)-value.

\(^{d}\) Number in square bracket below test statistic is \(p\)-values.
Table 4.8: Markov-switching Model a Estimated Results b for 11 Emerging Economies (2)

<table>
<thead>
<tr>
<th>Countries</th>
<th>Mexico</th>
<th>Russia</th>
<th>South Africa</th>
<th>South Korea</th>
<th>Thailand</th>
</tr>
</thead>
<tbody>
<tr>
<td>Intercept (Tranquil)</td>
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<td>-3.33</td>
<td>-1.93</td>
<td>-3.36</td>
<td>-6.44</td>
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<tr>
<td></td>
<td>(-3.19)</td>
<td>(-3.29)</td>
<td>(-6.06)</td>
<td>(-11.02)</td>
<td>(-13.38)</td>
</tr>
<tr>
<td>Intercept (Volatile)</td>
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<td>4.74</td>
<td>1.66</td>
<td>1.63</td>
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<tr>
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<td>(3.30)</td>
<td>(0.88)</td>
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<td>-0.06</td>
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<td>(4.08)</td>
<td>(-0.92)</td>
<td>(15.71)</td>
<td>(3.03)</td>
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<td>Std. Dev. (Tranquil)</td>
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<td>4.19</td>
<td>1.10</td>
<td>0.43</td>
<td>1.62</td>
</tr>
<tr>
<td>Std. Dev. (Volatile)</td>
<td>11.70</td>
<td>17.81</td>
<td>2.22</td>
<td>7.10</td>
<td>5.38</td>
</tr>
<tr>
<td>Transition Prob. (%)</td>
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<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>(Tranquil → Tranquil)</td>
<td>1.00</td>
<td>0.86</td>
<td>0.62</td>
<td>0.96</td>
<td>0.99</td>
</tr>
<tr>
<td>(Volatile → Volatile)</td>
<td>0.62</td>
<td>0.52</td>
<td>0.23</td>
<td>0.16</td>
<td>0.56</td>
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<td>SC criterion</td>
<td>6.35</td>
<td>5.01</td>
<td>5.43</td>
<td>5.02</td>
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<tr>
<td>LR Linearity test d</td>
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<td>[0.001]</td>
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</tbody>
</table>

a Estimated Model: Non-linear Markov-switching Autoregressive (MS-AR) model with a regime-dependent intercept and allowing for heteroscedasticity in the disturbance term.
b All results are obtained using H-M. Krohig’s MSVAR package for Ox.
c Number in parenthesis below regression coefficient is t-value.
d Number in square bracket below test statistic is p-values.
Figure 4.22: Argentina: Estimated Probabilities of being in Volatile State by SETAR Model 1993Q3–2003Q4

Figure 4.23: Brazil: Estimated Probabilities of being in Volatile State by SETAR Model 1993Q3–2003Q4

Figure 4.24: Czech Republic: Estimated Probabilities of being in Volatile State by SETAR Model 1993Q3–2003Q4
Figure 4.25: Hungary: Estimated Probabilities of being in Volatile State by SETAR Model 1993Q3–2003Q4

Figure 4.26: India: Estimated Probabilities of being in Volatile State by SETAR Model 1993Q3–2003Q4

Figure 4.27: Malaysia: Estimated Probabilities of being in Volatile State by SETAR Model 1993Q3–2003Q4
Figure 4.28: Mexico: Estimated Probabilities of being in Volatile State by SETAR Model 1993Q3–2003Q4

Figure 4.29: Russia: Estimated Probabilities of being in Volatile State by SETAR Model 1993Q3–2003Q4

Figure 4.30: South Africa: Estimated Probabilities of being in Volatile State by SETAR Model 1993Q3–2003Q4
Figure 4.31: South Korea: Estimated Probabilities of being in Volatile State by SETAR Model 1993Q3–2003Q4

Figure 4.32: Thailand: Estimated Probabilities of being in Volatile State by SETAR Model 1993Q3–2003Q4

Figure 4.33: Argentina: Filtered and Smoothed Probabilities of being in Volatile State by Markov-switching Model 1993Q3–2003Q4
Figure 4.34: Brazil: Filtered and Smoothed Probabilities of being in Volatile State by Markov-switching Model 1993Q3–2003Q4

Figure 4.35: Czech Republic: Filtered and Smoothed Probabilities of being in Volatile State by Markov-switching Model 1993Q3–2003Q4

Figure 4.36: Hungary: Filtered and Smoothed Probabilities of being in Volatile State by Markov-switching Model 1993Q3–2003Q4
Figure 4.37: India: Filtered and Smoothed Probabilities of being in Volatile State by Markov-switching Model 1993Q3–2003Q4

Figure 4.38: Malaysia: Filtered and Smoothed Probabilities of being in Volatile State by Markov-switching Model 1993Q3–2003Q4

Figure 4.39: Mexico: Filtered and Smoothed Probabilities of being in Volatile State by Markov-switching Model 1993Q3–2003Q4
Figure 4.40: Russia: Filtered and Smoothed Probabilities of being in Volatile State by Markov-switching Model 1993Q3–2003Q4

Figure 4.41: South Africa: Filtered and Smoothed Probabilities of being in Volatile State by Markov-switching Model 1993Q3–2003Q4

Figure 4.42: South Korea: Filtered and Smoothed Probabilities of being in Volatile State by Markov-switching Model 1993Q3–2003Q4

Figure 4.43: Thailand: Filtered and Smoothed Probabilities of being in Volatile State by Markov-switching Model 1993Q3–2003Q4
Empirical Tests of Contagion in 1997/98 East Asia Crises

5.1 Introduction

The newly industrializing countries in East Asia, once regarded as the world's most dramatic economic success, suddenly slipped into financial/currency turmoil in September 1997. The most striking aspect of the crises was that the crisis quickly spread from Thailand where it originated, to the other countries in the East Asian region as well as to Brazil and Russia in 1998. In other words, the crises demonstrate that a sharp depreciation in one country's currency can have a strong, negative impact on the vulnerability of exchange rate regimes in its neighboring countries—which may or may not share a similar economic structure and development. Do these episodes of highly correlated co-movements amongst different exchange markets provide any evidence of contagion?

Before answering the question, it is necessary to define contagion. The mainstream literature suggests that contagion occurs if the cross-market correlation amongst countries increases significantly during periods of turmoil, compared with
periods of stability. Based on this concept, some economists find evidence of contagion effects in several crisis episodes, including the 1997/98 East Asia Crises. But as Forbes and Rigobon (2002) argue, these traditional tests for contagion that are based on conditional correlation coefficients, are biased due to heteroscedasticity. We will follow their approach to adjust the correlation coefficients for this bias, and base our contagion tests on the resultant unconditional cross-market correlation coefficients.

The main contribution of this chapter lies in the adoption of the non-linear Markov-switching VAR model to describe the crisis transmission mechanism for the first time. We focus our research on recent experiences in East Asia—from July 1991 to December 1998, while Thailand is chosen as the country where the crisis originally takes place. We also use the Market Pressure Index (MPI) to capture both the successful and unsuccessful attacks on the East Asian currencies.

The estimated results justify our usage of the Markov-switching VAR model in depicting crisis episodes in East Asia during the 1990's. It is also shown that tests based on the conditional correlation coefficients find evidence of contagion between Thailand, Indonesia, Malaysia and Philippines. However, the adjustment for heteroscedasticity shows a significant impact on the test results, as in Forbes and Rigobon (2002). The unconditional (adjusted) correlation coefficients fall down during the crises and are smaller than ones in more stable periods in all cases, indicating that all evidence of contagion disappears after adjustment. Instead, the high level of market co-movements between Thailand, Malaysia, Philippines and Indonesia shows the strong economic linkage—interdependence—amongst these East Asian economies, which exists in all states of the world.

This chapter is organized as follows. A definition of contagion is given in section 2 where we also briefly survey the relevant empirical literature. Section 3 discusses the disadvantage of the conventional technique of using conditional cross-market correlation coefficients to test for contagion. One adjustment method suggested by Forbes and Rigobon (2002) is given to correct the bias due to heteroscedasticity. Section 4 gives the interested variables and data set. The Markov-switching VAR
Empirical Tests of Contagion in 1997/98 East Asia Crises

model is then employed to test for contagion in East Asian economies during the 1997/98 crises in section 5. The final section draws up the conclusion.

5.2 Survey of Recent Contagion Literature

5.2.1 Definition of Contagion

We follow the mainstream literature to define contagion as a significant increase in cross-market linkages after a shock to one country (or group of countries). It is therefore regarded as contagion only if the magnitude of co-movement amongst different markets shows a drastic increase during periods of turmoil, compared with periods of stability. On the other hand, if the cross-market correlation does not increase significantly after the shock, then any continued level of market correlation during crises only suggests interdependence amongst markets, which exists in all states of the world.

One advantage of adopting this definition is that it can distinguish between alternative approaches to explaining how crises are transmitted across markets. The current literature about contagion is so extensive that it could be useful if we can differentiate some theories that assume that investors behave differently after a crisis, from others that argue the shock is propagated via a channel that exists pre-crisis, such as trade linkages. Furthermore, empirical works based on this definition of contagion can provide evidence how the transmission mechanisms behave during various currency crises, as it is argued that only those countries that are affected by crisis through a channel magnified by the crisis should qualify for multilateral bailouts.

5.2.2 Survey of Recent Evidence on Contagion

The literature that attempts to test for the effects of contagion is extensive. Based on the same definition as ours, four different methodologies are employed to measure how shocks are transmitted across markets: the cross-market correlation coefficient,
the ARCH (GARCH) model, the cointegration model, and direct estimation of a specific transmission mechanism.

Amongst them, the cross-market correlation coefficient is the most widely used method to test for contagion. The correlation coefficients between two economies are measured in both crisis and relatively stable periods, in order to see whether the coefficient strengthens after a shock to one country. Calvo and Reinhart (1996) adopt this approach to test for contagion in stock prices and Brady bonds after the 1994 Mexican peso crisis. They find that the cross-market correlation coefficients increase significantly for many emerging markets. Baig and Goldfajn (1998) also find evidence of contagion in asset returns during the 1997/98 East Asia Crises, based on significant increases in cross-market correlation.

However, Forbes and Rigobon (2002) argue that these tests are biased and inaccurate due to heteroscedasticity, as the cross-market correlation coefficients are conditional on market volatility. Therefore, during crises when markets are more volatile, estimated correlation coefficients tend to increase and be biased upward. When tests do not adjust for this bias in the correlation coefficients, they traditionally find evidence of contagion. In this chapter, we adopt a recommended formula to adjust the cross-market correlation coefficients for heteroscedasticity, and employ a non-linear Markov-Switching VAR model to deliver a better description of crisis transmission mechanism.

The second approach is to adopt an ARCH or GARCH framework to estimate the variance-covariance transmission mechanism between countries. Edwards (1998) examines linkages between bond markets after the 1994 Mexican peso crisis. He finds that there are significant spillovers from Mexico to Argentina, but does not explicitly test if this transmission mechanism changes significantly after the shock.

The third method tests for changes in the cointegrating vector between markets over a long period of time. One disadvantage of this approach is that the cross-market relationship over such a long period can change for a number of reasons other than contagion, such as greater trade integration.
The final approach attempts to measure how the vulnerability of an exchange rate regime in one country can be affected by shocks to another country, along with other economic variables. The literature following this approach is extensive and normally incorporates other empirical methodologies as well, e.g., Logit/Probit model, as discussed in Chapter 2 Literature Review. Many find that the probability of a crisis in one country is highly associated with speculative attacks on the currency of a neighboring country.

Also note that as in Chapter 2 Literature Review, the regime switching models are increasingly applied to currency crises over the past few years (see Piard (1999) and Jeanne and Masson (2000)). But most of them analyze the effects of a range of variables on the probabilities that a country switches from one regime to another. Our research, however for the first time, employs the Markov-switching VAR model to examine the crisis transmission mechanism in 1997/98 East Asia Crises. However, it is necessary to first go through the argument about the bias in the traditional test for contagion that is based on the conditional cross-market correlation coefficient.

5.3 Bias in the Conditional Cross-market Correlation Coefficient

The discussion how changes in market volatility during crises can bias the test for contagion based on the cross-market correlation coefficient, is initially motivated by Ronn (1998). He addresses the issue in his estimation of intra-market correlation in stocks and bonds. More recently, Boyer et al. (1999) and Loretan and English (2000) start to investigate the problem in more detail. They use their own statistical framework to address the bias when testing for contagion.

A more complete documentation of this issue is found in Forbes and Rigobon (1999) and Forbes and Rigobon (2002). They not only provide some initiative examples of how heteroscedasticity can affect the cross-market correlation coefficient (while the underlining mechanism remains unchanged), but also give the formal
proof of the bias. As a byproduct, a formula is proposed to adjust for the bias when testing for contagion.

Assume $x$ and $y$ are two stochastic variables that represent exchange market pressure in two different markets. The relationship between them is specified in the following equation:

$$y_t = \alpha + \beta x_t + \epsilon_t$$  \hspace{1cm} (5.1)

If we assume the absence of endogeneity and omitted variables:

$$E[\epsilon_t] = 0; \quad E[\epsilon_t^2] = c < \infty$$ \hspace{1cm} (5.2)
$$E[x_t \epsilon_t] = 0$$ \hspace{1cm} (5.3)

then we can write the conditional cross-market correlation coefficient as

$$\rho^* = \rho \sqrt{\frac{1 + \delta}{1 + \delta \rho^2}}$$ \hspace{1cm} (5.4)

or

$$\rho = \frac{\rho^*}{\sqrt{1 + \delta [1 - (\rho^*)^2]}}$$ \hspace{1cm} (5.5)

where $\rho^*$ is the conditional correlation coefficient, $\rho$ is the unconditional correlation coefficient (the underlying transmission mechanism), and $\delta$ is the relative increase in the variance of $x$ defined by:

$$\delta = \frac{\sigma^2_{xx}}{\sigma^2_{xx}} - 1$$ \hspace{1cm} (5.6)

It is clear in Equation (5.4) that the conditional correlation coefficient could increase in $\delta$. In other words, even though the unconditional correlation coefficient (the underlying transmission mechanism) remains unchanged, the conditional correlation coefficient becomes larger in crises when $x$ is more volatile. Therefore, the
traditional test for contagion that is based on the conditional correlation coefficients is biased and inaccurate—as the market tends to be more volatile during the crisis, the estimated conditional correlation coefficient tends to increase although the cross-market linkage remains the same as during more stable periods.

Equation (5.5) also provides us with a method to adjust for this bias, as it is a relatively good approximation of the unconditional cross-market correlation coefficient. Therefore, the null hypothesis that we would like to test in our analysis is that adjustment for heteroscedasticity does not change the results of traditional tests for contagion. The alternative hypothesis, on the other hand, is that adjustment proposed by Forbes and Rigobon (2002) has a significant impact on the test results. However, three criteria are required to justify this method for adjustment: (1) a major shift in market volatility; (2) clear identification of the original country where the shift is generated; and (3) including only one neighboring country each time in the regression to test for contagion.

5.4 The Variables and Data Set

5.4.1 The Variables of Interest

As in previous chapters, we establish the Market Pressure Index (MPI) to capture not only the actual depreciations of the domestic currency, but also speculative attacks that are warded off by increasing interest rates and/or selling international reserves. The MPI is a weighted sum of the nominal exchange rate depreciation, the change in market interest rate, and the negative percentage change in foreign exchange reserves according to their relative precision.

Krugman (1998) suggests in his Moral Hazard model that moral hazard (international over-borrowing) in the banking sector plays a crucial role in the 1997/98 East Asia Crises, acting as a main force for bubbles in asset markets. Therefore, we

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\(^1\) As in Chapter 3, the minus sign ensures that a reduction in foreign reserves translates to an increase in the value of the MPI.
include in our regression: (1) the Ratio of Domestic Bank Assets to Nominal GDP.

In Change and Velasco (1998)'s illiquidity model, the liquidity crunch experienced by the East Asian economies once the currency regimes collapse, is believed to be another cause of the crises. Thus, (2) the Ratio of Bank Loan to Bank Total Deposit is also considered in our model.

The declining export competitiveness in East Asian countries are widely blamed as a deeper cause for the 1997/98 turmoil, and the relevant variables are commonly selected as a leading indicator in empirical studies. Here, we also include (3) the Growth Rate of Export (% of previous year) in our regression.

5.4.2 The Data Set

Most of our data is extracted from the on-line IMF’s International Financial Statistics (IFS) and the Global Development Finance (GDF) 2004 CD. We also acquire data from the working papers published at the World Bank’s website.

The data set consists of monthly observations from July 1991 to December 1998 for Thailand, Malaysia, Philippines, Indonesia and South Korea. We do not adopt a longer sample period, because possible structural changes can happen in markets over a long period which invalidates our test for contagion. The end point of December 1998 is chosen as Malaysia started to re-adopt a pegged exchange rate regime and imposed strict controls on capital flows at the beginning of 1999.

5.5 Contagion Analysis using MS-VAR model

As discussed in Chapter 2 Literature Review, the regime switching models have been increasingly applied to the currency crises over the past few years. It is shown that over a long period, many macroeconomic and financial variables, including the MPI we construct, sometimes undergo episodes when the behaviours of the series change dramatically. Such events include wars, currency crises, and significant changes in government policies. The regime switching models can allow us to assume that
observed data is drawn from the different distributions conditional on the state-contingent parameter sets.

The main advantage of the regime switching framework over the Logit/Probit model is that it allows for a continuous dependent variable and avoids the arbitrary choice of the threshold value for a crisis. Moreover, the non-linear nature of the regime switching models is also appealing as the economy switches endogenously amongst different regimes.

5.5.1 The MS-VAR model

We adopt Hamilton's Markov-switching \(^2\) Vector Autoregressive model (MS-VAR) to examine the experiences of East Asian countries during the 1990's, while Thailand is selected as the origin country where the crisis spreads to other countries in the region. It is presumed in Hamilton's model that we cannot observe the shifts in regimes directly, but instead must draw probabilistic inferences about whether and when changes may occur, based on the observed behaviours of the MPIs.

Here, our Markov-switching model assumes regime shifts in the intercept term and also allows for heteroscedasticity in the disturbance term. It is also assumed that there are two regimes in our model—the crisis regime and the relatively stable regime. A VAR framework is employed in order to control for endogeneity. As discussed above, the MPI rather than the simple nominal depreciation rate is used so that we can encompass all the periods of both large depreciations and successful defences against severe attacks on exchange rates.

We also include in the model three economic variables based on the crisis theories and empirical works: (1) the negative growth rate of exports (% as the same month in previous year,); (2) the Domestic Bank Assets (% of Nominal GDP); and (3) the ratio of Bank Loan to Bank Total Deposit. The negative of the export growth rate ensures that its coefficient has a predicted positive sign, as a slowdown in exports

\(^2\)See Appendix A for more details. Also see Hamilton (1994), chapter 22, for a statistical review of Markov-switching models and Hamilton (1989) for an application of the MS model to US growth trend.
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increases the vulnerability of the economy to speculative attacks.

The MS-VAR model can be formally expressed as

$$x_t = v(s_t) + \phi(L)x_t + \psi(L)I_t + \eta_t$$

$$x_t = \{x_t^c, x_t^j\}'$$

$$I_t = \{exp_t^c, exp_t^j, lend_t^c, lend_t^j, lndp_t^c, lndp_t^j\}'$$

where $x_t^c$ is the MPI of the crisis source country; $x_t^j$ is the MPI of another East Asian country $j$; so $x_t$ is a transposed vector of MPI of both countries. $\phi(L)$ is a vector of associated polynomials in the lag operator. $exp_t^c$ and $exp_t^j$ are export growth rate for the crisis source country and country $j$ respectively; $lend_t^c$ and $lend_t^j$ are domestic bank assets (% of nominal GDP) for two countries; and $lndp_t^c$ and $lndp_t^j$ are the ratio of bank loan to total deposit for the crisis source country and country $j$ respectively; therefore, $I_t$ is a transposed vector of all exogenous variables for two countries, while $\psi(L)$ is a vector of associated polynomials in the lag operator. Finally, the conditional constant term (intercept) $v(s_t)$ is assumed to switch between two states:

$$v(s_t) = \begin{cases} v_1, & \text{if } s_t = 1 \text{ ('tranquil') } \\ v_2, & \text{if } s_t = 2 \text{ ('crisis') } \end{cases}$$

and $\eta_t \sim \text{NID}(0, \Sigma(s_t))$, where $\Sigma(s_t)$ is also presumed to be different in different regimes. Equation (5.7)-(5.9) form the measurement equation for Hamilton's state space model.

The state $s_t$ is assumed to follow a first-order Markov process so that we can write the transition equation as

$$\Pr(s_t = i|\xi_{t-1}) = \Pr(s_t = i|s_{t-1} = j) = p_{ij}$$

where $\xi_{t-1}$ is a vector representing all the information available at time $t - 1$, includ-
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The lagged values of $x_t$ and $I_t$. The essence of this method is the presumption that the future will in some sense be like the past. The measurement equation and the transition equation make up Hamilton's Markov-switching space model. The value $p_{ij}$ is known as the transition probability of switching from state $j$ to state $i$, and is assumed to be independent of time:

$$
\begin{align*}
    P_{12} &= \Pr(\text{tranquil in } t | \text{crisis in } t-1) \\
    P_{21} &= \Pr(\text{crisis in } t | \text{tranquil in } t-1) 
\end{align*}
$$

for $t \in \{1, +\infty\}$.

As shown in Appendix A, the Hamilton filter provides a nonlinear algorithm for generating inferences about a discrete-valued unobserved state vector. As a by-product of that algorithm, the log likelihood function for the observed data evaluated at the estimated parameters can be calculated. Then, the EM method is adopted to obtain maximum likelihood estimators for all the parameters in the MS-VAR model.

In general, the EM method (developed by Dempster et al. (1977)) maximizes the incomplete-data log likelihood via the iterative maximization of the expected complete-data log likelihood, conditional upon the observable data. Given the observed data and some initial estimators of the parameters in the model, the EM algorithm begins by calculating the smoothed state probabilities (i.e. the unconditional probability of a particular state). With the estimated smoothed state and transition probabilities, the expected complete-data log likelihood function is constructed. This is the 'E', expectation part of the algorithm. The expected complete-data log likelihood function is then maximized to obtain updated parameter estimators. This is the 'M', maximization part of the algorithm. Using this updated estimators, the smoothed probabilities are calculated again and substituted into the expected likelihood function, which is maximized again. This iterative procedure is repeated until convergence (in the parameter estimators or the likelihood function) is achieved. Then we can obtain the maximum likelihood estimators of the MS model parameters.
It is also possible to derive smoothed estimators of the state probabilities based on all the sample information. These smoothed probabilities represent the best estimation of the probabilities for a state in period \( t (s_t) \). Since the Markov-switching model includes the possibility that the threshold depends on the last regime, the smoothed probabilities are different from the filter probabilities. Kim (1994) develops a backward recursion algorithm to calculate this smoothed inference.

We will test for contagion from Thailand to other East Asian countries—Malaysia, Philippines, Indonesia and South Korea—during the 1997/98 crises separately. For each pair of countries, we estimate the MS-VAR model specified in equations (5.7) through (5.12). Using the variance-covariance estimators from the regression, we calculate the cross-market correlation coefficients during both the more stable periods and periods of turmoil. We then see whether there is a significant increase in the correlation coefficients (not adjusted for heteroscedasticity yet) during crises, and hence evidence for contagion.

However, these test results may not be accurate, as the correlation coefficients are biased due to heteroscedasticity. In other words, the increases in the estimated correlation coefficient might not reflect increases in real cross-market linkages but rather increases in market volatility. Therefore, we modify our tests for contagion to be based on the unconditional cross-market correlation coefficients which are calculated using equation (5.5). Finally, we see the impact of the adjustment on the test results for contagion in the 1997/98 East Asia crises.

### 5.5.2 Results

It is shown in the estimated results (Table 5.1 and Table 5.2) that the relatively stable states are characterized by low variances, while most periods of turmoil are depicted by much higher volatilities, except for Malaysia. In other words, the periods of stability are the times of little fluctuation in foreign exchange markets, but crises are the periods with high market volatility.
Table 5.1: MS-VAR Model Estimated Results Table 1

<table>
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<th>Thailand-Malaysia</th>
<th>Thailand-Philippine</th>
</tr>
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<tbody>
<tr>
<td></td>
<td>Tha</td>
<td>Mal</td>
</tr>
<tr>
<td>Intercept (Stable)</td>
<td>6.99 (3.35)</td>
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<tr>
<td></td>
<td>-1.27</td>
<td>0.06</td>
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<tr>
<td>Intercept (Crisis)</td>
<td>8.63 (3.86)</td>
<td>-5.53 (-3.05)</td>
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<td></td>
<td>1.87</td>
<td></td>
</tr>
<tr>
<td>AR_Tha(-1)</td>
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<td>0.01 (-0.09)</td>
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<tr>
<td></td>
<td>-0.03</td>
<td>0.05</td>
</tr>
<tr>
<td>AR_Mal/Phi (-1)</td>
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<td>0.002 (-0.006)</td>
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<tr>
<td>EXP_Tha</td>
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<tr>
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<td>1.15</td>
<td></td>
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<tr>
<td>EXP_Tha(-1)</td>
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<td>-0.63 (-0.95)</td>
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<tr>
<td></td>
<td>(-1.29)</td>
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<td>EXP_Mal/Phi</td>
<td>1.37 (1.82)</td>
<td>-1.37 (-2.03)</td>
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<td></td>
<td>(1.37)</td>
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<td>EXP_Mal/Phi(-1)</td>
<td>2.49 (2.20)</td>
<td>0.69 (-0.02)</td>
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<td>(4.07)</td>
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</tr>
<tr>
<td>BKASS_Tha</td>
<td>4.56 (1.10)</td>
<td>4.51 (3.22)</td>
</tr>
<tr>
<td></td>
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<tr>
<td>BKASS_Tha(-1)</td>
<td>-3.63 (-2.22)</td>
<td>-3.63 (-2.22)</td>
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<tr>
<td>BKASS_Mal/Phi(-1)</td>
<td>-7.82 (-2.06)</td>
<td>0.23 (-0.26)</td>
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<tr>
<td>BKLIQ_Tha</td>
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<td>27.65 (12.16)</td>
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<td>(-1.16)</td>
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<table>
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<th>MPI</th>
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<tr>
<td></td>
<td>Std. Dev. (Stable)</td>
<td>Std. Dev. (Crisis)</td>
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<td>MPI</td>
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<td>(Crisis)</td>
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<td>-2.13</td>
<td>5.58</td>
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<td></td>
<td>(-0.29)</td>
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<td>(-1.39)</td>
<td>(-3.69)</td>
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<td>BKLQ_Mal/Phi(-1)</td>
<td>-0.97</td>
<td>-3.80</td>
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<tr>
<td></td>
<td>(-0.48)</td>
<td>(-3.03)</td>
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<tr>
<td></td>
<td>(2.96)</td>
<td>(1.52)</td>
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</tbody>
</table>

Regime Duration (months)

| Stable Regime | 62.31 | 17.38 |
| Crisis Regime | 9.52  | 10.59 |

Transition Prob. (%)

| (Stable → Stable) | 0.98 | 0.94 |
| (Crisis → Crisis) | 0.89 | 0.91 |
| (Stable → Crisis) | 0.02 | 0.06 |
| (Crisis → Stable) | 0.11 | 0.09 |

Classified

<table>
<thead>
<tr>
<th>Crisis periods</th>
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SC criterion 5.11 6.67

LR Linearity test 76.01 63.61
Notes: Estimated Model: Non-linear Markov-switching Vector Autoregressive model (MS-VAR) with the regimes-dependent intercept and heteroscedasticity in the disturbance term. Number of regimes: 2. Order of VAR: chosen according to the (lowest) information criteria. Number of exogenous variables: 3. All results are obtained using H-M. Krolzig's MSVAR package for Ox. Number in parenthesis below regression coefficient is t-value. Number in parenthesis besides variable indicates number of lag(s) of AR term. Number in square bracket below test statistic is p-value.

EXP—Growth Rate of Export (% of previous year).

BKASS—Ratio of Domestic Bank Assets to Nominal GDP.

BKLIQ—Ratio of Bank Loan to Bank Total Deposit.

Table 5.2: MS-VAR Model Estimated Results Table 2

<table>
<thead>
<tr>
<th>MPI</th>
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<th>Thailand-South Korea</th>
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<tr>
<td></td>
<td>Tha</td>
<td>Ind</td>
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<tr>
<td>Intercept (Stable)</td>
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<td>(0.50)</td>
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<td>Intercept (Crisis)</td>
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<td>(0.97)</td>
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<td>AR_Tha(-3)</td>
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<td>(-0.76)</td>
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Table 5.2: MS-VAR Model Estimated Results Table 2

<table>
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<tr>
<th>MPI</th>
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<th>Thailand-South Korea</th>
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<td>EXP. Tha</td>
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<td>(1.22)</td>
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<td>(1.12)</td>
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<td>BKASS. Ind/Kor(-1)</td>
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<td>7.96</td>
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<td>(1.63)</td>
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<td>BKLIQ. Tha(-1)</td>
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<td>(-1.34)</td>
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<td>4.84</td>
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<td>(1.62)</td>
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<td>BKLIQ. Ind/Kor(-1)</td>
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<td>Std. Dev. (Stable)</td>
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<td>0.47</td>
<td>0.36</td>
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Table 5.2: MS-VAR Model Estimated Results Table 2

<table>
<thead>
<tr>
<th>MPI</th>
<th>Thailand-Indonesia</th>
<th>Thailand-South Korea</th>
</tr>
</thead>
<tbody>
<tr>
<td>Std. Dev. (Crisis)</td>
<td>2.09</td>
<td>3.17</td>
</tr>
<tr>
<td>Regime Duration (months)</td>
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<tr>
<td>Stable Regime</td>
<td>23.50</td>
<td>13.71</td>
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<tr>
<td>Crisis Regime</td>
<td>4.44</td>
<td>2.15</td>
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<tr>
<td>Transition Prob. (%)</td>
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<td></td>
</tr>
<tr>
<td>(Stable → Stable)</td>
<td>0.96</td>
<td>0.93</td>
</tr>
<tr>
<td>(Crisis → Crisis)</td>
<td>0.77</td>
<td>0.54</td>
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<tr>
<td>(Stable → Crisis)</td>
<td>0.04</td>
<td>0.07</td>
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<td>(Crisis → Stable)</td>
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<td>0.46</td>
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<tr>
<td>Crisis periods</td>
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<tr>
<td>SC criterion</td>
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<td>5.97</td>
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<tr>
<td>LR Linearity test</td>
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<tr>
<td>72.94</td>
<td>104.75</td>
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<tr>
<td>[0.000]</td>
<td>[0.000]</td>
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</table>

Notes: Estimated Model: Non-linear Markov-switching Vector Autoregressive model (MS-VAR) with the regime-dependent intercept and heteroscedasticity in the disturbance term. Number of regimes: 2. Order of VAR: chosen according to the (lowest) information criteria. Number of exogenous variables: 3. All results are obtained using H-M. Krolzig's MSVAR package for Ox. Number in parenthesis below regression coefficient is t-value. Number in parenthesis besides variable indicates number of lag(s) of AR term. Number in square bracket below test statistic is p-values.

EXP—Growth Rate of Export (% of previous year).
BKASS—Ratio of Domestic Bank Assets to Nominal GDP.
BKLIQ—Ratio of Bank Loan to Bank Total Deposit.
Empirical Tests of Contagion in 1997/98 East Asia Crises

The adoption of formula (5.5) to adjust for heteroscedasticity is justified in that the three criteria mentioned above are satisfied. The variances of the MPI in all East Asian economies increase significantly in the 1997/98 East Asia Crises, compared with the tranquil periods. The country where the crisis originally breaks out is Thailand—the devaluation of Thai Bhat in September 1997 is widely regarded as the start point of the financial/currency turmoil for the region. Finally, we only test for contagion from Thailand to one of its neighboring countries each time in our MS-VAR model.

It is seen that duration of a relatively stable state is much longer than a period of crisis. The chances that the East Asian countries remain in the same state—no matter it is a regime of turmoil or stability—are also much higher than when countries switch between different states. In other words, the economies in East Asia that enjoy a relatively stable state would stay in ‘calm’ for the next period, while the countries that find themselves struggling in crises could be dragged even deeper into crises in the following month.

Our results show that the hypotheses of linear specification are rejected at even the 1% significant level in all cases. This confirms our assumption that time series variables such as the MPI undergo regime changes when facing dramatic events like currency crises, and hence justifies our usage of the Markov-switching model to depict crisis transmission mechanisms. The three exogenous variables that we include in our estimation to help explain movements in exchange markets based on crisis theories and previous works have frequent statistically significant coefficients, which reduces the risk of having omitted variables in our models.

As discussed above, the unconditional smoothed probability is the inferred probability of a particular state in a period which is estimated based on all the sample information. A month is then classified as a month of turmoil if the smoothed probability of being in crisis is greater than 50% and vice versa. It is noted that 50% is not necessarily the only choice of threshold, as the inferred smoothed probabilities are all close to either zero or one.
Empirical Tests of Contagion in 1997/98 East Asia Crises

Table 5.3: Conditional Cross-market Correlation Coefficients

<table>
<thead>
<tr>
<th></th>
<th>Thailand-Malaysia</th>
<th>Thailand-Philippines</th>
<th>Thailand-Indonesia</th>
<th>Thailand-South Korea</th>
</tr>
</thead>
<tbody>
<tr>
<td>Stable periods b</td>
<td>0.33</td>
<td>0.04</td>
<td>0.09</td>
<td>0.05</td>
</tr>
<tr>
<td>Turmoil periods</td>
<td>0.50</td>
<td>0.12</td>
<td>0.30</td>
<td>0.04</td>
</tr>
<tr>
<td>Contagion c</td>
<td>C</td>
<td>C</td>
<td>C</td>
<td>N</td>
</tr>
</tbody>
</table>

a This table reports conditional (unadjusted) cross-market correlation coefficients for Thailand and each of other countries in the sample.

b Both stable and crisis periods are defined by the Markov-switching Vector Autoregressive model (MS-VAR).

c 'C' indicates that contagion occurs, while 'N' shows no crisis contagion effect.

Table 5.4: Unconditional Cross-market Correlation Coefficients

<table>
<thead>
<tr>
<th></th>
<th>Thailand-Malaysia</th>
<th>Thailand-Philippines</th>
<th>Thailand-Indonesia</th>
<th>Thailand-South Korea</th>
</tr>
</thead>
<tbody>
<tr>
<td>Stable periods b</td>
<td>0.35</td>
<td>0.04</td>
<td>0.09</td>
<td>0.05</td>
</tr>
<tr>
<td>Turmoil periods</td>
<td>0.31</td>
<td>0.03</td>
<td>0.08</td>
<td>0.01</td>
</tr>
<tr>
<td>Contagion c</td>
<td>N</td>
<td>N</td>
<td>N</td>
<td>N</td>
</tr>
</tbody>
</table>

a This table reports unconditional (after adjusted for heteroscedasticity) cross-market correlation coefficients for Thailand and each of other countries in the sample.

b Both stable and crisis periods are defined by the Markov-switching Vector Autoregressive model (MS-VAR).

c 'C' indicates that contagion occurs, while 'N' shows no crisis contagion effect.


The estimated conditional and unconditional cross-market correlation coefficients for each pair of countries during both the stable and crisis periods are shown in Table 2 and 3 respectively.

We here report the absolute value of the estimated correlation coefficients.
Thailand and Malaysia

Several patterns are immediately apparent in the test results for a contagion effect between Thailand and Malaysia. Firstly, the cross-market correlation coefficient between the two countries during the relatively stable periods is unsurprisingly high, given that they are both key members of ASEAN and share a similar economic structure and development. In other words, the Thai foreign exchange market is highly correlated with the Malaysian market during the periods of stability (0.35).

Secondly, the correlation coefficient increases dramatically to 0.80 during the periods of turmoil—more than double what it is in a relatively stable state. This shows an even higher level of comovement between the two markets during crises and thus clear evidence of contagion. Therefore, based on the conditional cross-market correlation coefficient (not adjusted for heteroscedasticity), our test finds that contagion spreads from the crash of Thai exchange market to Malaysia.

However, it is shown that the adjustment for the bias due to heteroscedasticity has a significant impact on the estimated correlation coefficient and the resulting test for contagion. The unconditional cross-market correlation coefficient is now 0.31 during the periods of turmoil, and smaller than the correlation coefficient during the periods of stability (0.35). Thus, there is no significant increase in the unconditional correlation coefficient after the shock to Thailand. In other words, the evidence of contagion from the primary test disappears after we adjust for heteroscedasticity. This is exactly what Forbes and Rigobon (2002) find in their research in stock market returns after the 1997/98 East Asia Crises.

Furthermore, the results also highlight the way by which we define contagion. The continued high level of comovement between the Thai foreign market and the Malaysian market in periods of turmoil is not considered as evidence of contagion. Instead, the two economies are regarded as highly interdependent upon each other in all states of the world.
Thailand and Philippines

The test results for contagion between Thailand and Philippines shows similar patterns as in the previous case. The low cross-market correlation coefficient (0.04) between the two countries during relatively stable periods could be attributed to their different economic structure and development—the Thai economy is considered more developed than that of the Philippines.

It is also shown that the correlation coefficient increases significantly to 0.12 for the two economies during the periods of turmoil. This is the evidence of contagion from Thai crash to Philippines, if we base our test on the conditional cross-market correlation coefficient (not adjusted for the bias).

But the adjustment for heteroscedasticity has a significant impact on the test for contagion. Once again, we find that the higher level of market comovement between the two countries during crises could result from higher market volatilities (heteroscedasticity) instead of an increase in real cross-market linkages. After we adjust for bias and base our test for contagion on the unconditional correlation coefficient, there is no significant changes (increases) in the magnitude of the propagation mechanism between the Thai exchange market and the Philippines market—in other words, no contagion. The unconditional cross-market correlation coefficient is now 0.03, smaller than the correlation coefficient in the periods of stability (0.04).

Thailand and Indonesia

The same story repeats itself in the case of Thailand and Indonesia. During relatively stable periods, the Thai exchange market is correlated with the market in Indonesia at a low level (0.09), also because of different economic structure and development in two countries. But the cross-market correlation coefficient increases dramatically to 0.29 after a shock to the Thai market, more than three times as high as in periods of stability. Therefore, the test based on the conditional correlation coefficient finds the evidence that contagion occurs from the collapse of Thai exchange rate regime to Indonesia.
However, it is shown that the test results are entirely different after we adopt the proposed formula to adjust for heteroscedasticity. The unconditional cross-market correlation coefficient during periods of turmoil is now 0.06, smaller than 0.09 during times of stability. In other words, there is no more evidence of contagion between Thailand and Indonesia after we adjust our test for the bias due to heteroscedasticity.

Thailand and South Korea

South Korea is the third largest economy in East Asia and thus has a much more advanced economy than Thailand. This fact is reflected by the low cross-market correlation coefficient (0.05) during the stable periods.

In contrast to the previous cases, we do not find any evidence of contagion between Thailand and South Korea. The conditional (unadjusted) correlation coefficient decreases rather than increases to 0.04 during periods of crises. Furthermre, after adjusting for different market volatilities in crises, the correlation coefficient becomes even smaller (0.01) during the periods of turmoil. Therefore, there is no contagion occurring from the crash in Thai exchange market to South Korean market.

5.6 Conclusion

This chapter has shown the significant impact that the adjustment for the bias due to heteroscedasticity may have on the test for contagion.

The rapid spread of the crisis from Thailand to other countries in East Asia during 1997/98 crises raises an interesting question: do these highly correlated co-movements provide any evidence of contagion? Here, we adopt the definition of contagion that is widely used in the mainstream literature: a significant increase in cross-market linkages after a shock to one country (or group of countries). Therefore, contagion occurs if the cross-market correlation coefficient increases significantly during the periods of turmoil, compared with in periods of stability. However, as
Forbes and Rigobon (2002) argue, the traditional method of testing for contagion does not adjust for the bias due to increased market volatility during crisis periods, and thus gives inaccurate results.

Using the proposed formula to adjust for heteroscedasticity, we base our test for contagion on the unconditional cross-market correlation coefficient. For the first time, we employ the non-linear Markov-switching Vector Autoregressive (MS-VAR) model to depict the crisis transmission mechanism. Compared with conventional techniques, the MS-VAR model can allow us to avoid an arbitrary choice of threshold level, and its non-linear feature is also appealing. It is shown in the estimated results that the hypotheses of linear specification are rejected even at the 1% significant level in all cases, which justifies our usage of the MS-VAR model.

Not surprisingly, the tests based on the conditional cross-market correlation coefficient find evidence for contagion effects. In three of four cases, the correlation coefficients increase significantly during the turmoil periods, compared with periods of stability. However, as in Forbes and Rigobon (2002), all evidence for contagion soon disappears after we adjust our tests for the bias due to heteroscedasticity. The unconditional cross-market correlation coefficients are much smaller than their conditional counterparts during periods of turmoil, and even smaller than the correlation coefficients in periods of stability. Instead, the high level of market co-movements that exist in all states of the world shows real economic linkages—interdependence—between Thailand, Malaysia, Philippines and Indonesia.
Empirical Tests of Contagion in 1997/98 East Asia Crises

Figure 5.1: The Filtered and Smoothed Regime Probabilities for Thailand and Malaysia during 1991.10–1998.9

Figure 5.2: The Filtered and Smoothed Regime Probabilities for Thailand and Philippines during 1991.10–1998.9
Empirical Tests of Contagion in 1997/98 East Asia Crises

MSVAR: Regime probabilities
MSIH(2)-VARX(3), 1992 (1) - 1998 (9)

Figure 5.3: The Filtered and Smoothed Regime Probabilities for Thailand and Indonesia during 1991.10–1998.9

MSVAR: Regime probabilities
MSIH(2)-VARX(3), 1992 (1) - 1998 (9)

Figure 5.4: The Filtered and Smoothed Regime Probabilities for Thailand and South Korea during 1991.10–1998.9
The last 30 years saw frequent and severe speculative attacks on the exchange rate regimes in both developed and developing countries, such as the collapse of the European Exchange Rate Mechanism (ERM) in 1992/93, the 1994 Mexican Peso devaluation, the East Asia Financial/Currency crises in 1997/98, and the more recent Argentina turmoil (2001). The scale and impact of these episodes have prompted a huge amount of research interest as scholars are trying to provide different explanations for these ever-changing phenomena. The mainstream crisis literature includes the so-called three generation crisis models that we have discussed in detail in Chapter 2 Literature Review.

The initial approach of 'bad' fundamentals (e.g., Krugman (1979), Flood and Garber (1984)), indicating that crises can be created by expansionary fiscal and/or monetary policies, seems promising when tested to fit the crisis episodes in Latin American countries in the late 1970's and 1980's. This is followed by game-theoretical models with multiple equilibria (e.g., Obstfeld (1994), Obstfeld (1996)) that were initiated by the 1992/93 ERM crises. The second generation crisis models show that as a combination of the conflicting goals of government's policies with private
sector's perfect foresight can create multiple equilibria—the economy can be driven from an equilibrium where the fixed exchange rate regime is feasible to another one where it is not, even when the fundamentals are sound. The 1997/98 East Asia Crises prompted the third generation models, including models attempting to explain crises in terms of moral hazard (Krugman (1998)) and illiquidity (Change and Velasco (1998)), and models exploring the crisis transmission mechanism (Contagion, e.g., Baig and Goldfajn (1998), Glick and Rose (1999)). It is therefore natural to assume that the different generation crisis models are at best explaining the causes (and other aspects) of different currency crises by which they were originally inspired—crises occurring in Latin American countries seem mostly resulting from the governments' irresponsibly expansionary policies and the unsustainable credit expansion; conflicting goals of government's policies when combined with private sector's perfect foresight in Europe cause the ERM crashes; and an unhealthy banking system with problems of both moral hazard and liquidity crunch is at the center of East Asian turmoil.

In Chapter 3, we have searched for the stylized facts associated with currency crises in the three regions: Latin America, Europe and East Asia in order to see how good the three generation crisis models are in addressing them respectively. We use the country-specific Market Pressure Index (MPI) rather than the simple nominal depreciation rate to capture not only the actual devaluations but also 'failed' attacks, which are fended off by raising interest rates and/or selling reserves. The pioneering approach we adopt here to define currency crises is a combination of two non-linear state models—the SETAR (Self Exciting Threshold Autoregression) model and the Markov-switching model. Both models allow us to avoid an arbitrary choice of threshold level (i.e., 10% or a one standard deviation move), by allowing the data itself to reveal what move represents abnormal behaviour and thus signal a crisis.

The estimated results from our Logit model that link the binary crisis variable to an extensive range of economic variables selected on the basis of the three generation crisis models, are limitedly promising. On the one hand, we do find that
to some extent, the three generation models address the different perspectives of currency crises occurring in the three regions respectively. It is shown that a rapidly expanding stock of credit and large deficits in current account are important indicators for Latin American crises, consistent with the first generation models that emphasize ‘bad fundamentals’. We also find that the turmoil episodes in Europe are highly associated with its real economic cycle, i.e., the pressure in foreign exchange markets piles up when the economy slides into recession. This is the time when the discretionary governments are tempted to relax their policies to boost output according to the second generation models. As for East Asia, the (international) over-borrowing and lack of export competitiveness are shown to play an important role, which corresponds the predictions made by the third generation models. Our approach is also an example where the theories are neatly linked with empirical evidence.

However, the picture is far from perfect. It is seen that some variables appear not to have the predicted signs in the regressions, and others that are suggested by the theories are not statistically significant at the conventional level (10%). Due to the rigidity of our econometric model, it is unfair to simply conclude that the three generation crisis models are ineffective or inefficient at explaining the various currency crises across the world. But at the least, our analysis has shown that there is some room for improvement, which is the task laid in the chapters which follow. It is known that currency crises are so complicated and ever-changing phenomena that it is not possible for the three generation crisis models to address all issues surrounding the crises. In fact, we look at the issue of credibility of the exchange rate regime in currency crises in the following chapter.

Chapter 4 sheds some light on the two ambivalent impacts that public debt, especially debt maturity, has on currency crises: the ‘debt-burden’ effect and the ‘signalling’ effect. When discussing monetary policy, traditional economic theories focus on the monetary authority’s reputation, and insist that resisting a crisis can enhance the authority’s anti-inflation reputation, and hence the probability that the
Concluding Remarks

exchange rate will be maintained in the future. However, as Benigno and Missale (2001) argue, tight monetary policies aiming to defend the parity can, at the same time, increase the actual debt burden and thus impair the ability to withstand future crises. Thus, whether the exchange rate regime can gain credibility from a current devaluation may depend on the importance of the debt-burden effect relative to the signalling effect.

We adopt a three-period stochastic version of the Barro and Gordon (1983) model, and derive solutions by solving the optimal problem facing the monetary authority. We extend Benigno and Missale (2001)'s analysis to introduce non-deflatable debt, and also include a new scenario where the monetary authority can commit itself to a two-period rule—which is set up in period 0 and known to the public. The latter provides us with the socially optimal solution regarded as a benchmark. Another two scenarios have also been examined—where the authority cannot commit and its preference is publicly known, and where it also adopts discretionary policies but its type is not known to private sector. It is shown that a current devaluation may improve the prospects that the future exchange rate regime could be maintained by reducing the real value of public debt, as long as the debt-burden effect prevails over the signalling effect. However, if there is great uncertainty about the authority's preferences, a current depreciation may instead give a signal of a weak authority, and thus lead to higher expectations of a further devaluation. All these findings confirm the results of Benigno and Missale (2001), but in a more comprehensive manner. The key conclusions are also demonstrated by the numerical simulations using Matlab.

Our theoretical model also produces two testable predictions about the effect of debt maturity on interest rates. In countries where the debt-burden effect dominates, a larger proportion of short-term debt strengthens the incentive to roll it over at a lower-than-expected interest rate, and thus leads to a higher probability of a current devaluation (higher short-term interest rate). As the forward rate is independent of debt maturity, we have a flatter yield curve for debt-burden countries. But in
countries where there is substantial uncertainty about the authority’s preferences, the incentive to send a strong signal of the authority’s toughness implies that a higher ratio of short-term debt lowers the current interest rate, and leads to a steeper yield curve.

But there exist some obstacles for us to test the predictions in empirical work, in particular how to differentiate countries where the debt-burden effect prevails from countries where signalling dominates. As in Chapter 3, we construct the continuous Market Pressure Index (MPI) that incorporates changes in the exchange rate, reserves and interest rates, to gauge the vulnerability of exchange rate regime. We then employ both the SETAR (Self Exciting Threshold Autoregression) model and the Markov-switching model to identify the periods when foreign exchange markets are under severe speculative attacks. The actual devaluation episodes are defined by the regime changes in the Reinhart-Rogoff index. Then, the periods of successful defences of the parity are periods when the high pressure on domestic currency does not lead to an actual devaluation. Finally, we can draw the line between the debt-burden countries and the signalling countries, according to different patterns shown by their interest rates facing currency crises and/or successful defences. In this way, we distinguish between the two types of countries for our 11 emerging economies between 1993Q3—2001Q4.

Based on the two different country groups, we can conduct our tests for the two predictions by using GMM (General Method of Moments) for the dynamic panel data model. The estimated results show supportive evidence for our predictions. Overall, our main contributions here have been laying down a brand-new method to distinguish between the actual devaluations and the successful defences, and for the first time, providing empirical evidence for the ambivalent impacts of a current devaluation on the credibility of future exchange rate regime. In an earlier paper, Drazen and Masson (1994) suggest that the persistent effect of unemployment could provide another channel whereby current actions affect future policy. The direction of future research is, therefore, looking for other factors that can act as
an intermediate between the current and future, which are discussed in more detail later in this chapter.

Another area where our research project has made some progress is with regard to the crisis transmission mechanism. We follow the mainstream literature to adopt a narrow definition of contagion: a significant increase in cross-market linkages after a shock to one country (or group of countries). It is argued that the traditional approach to test for contagion is inaccurate because it is based on the conditional cross-market correlation coefficient that has not been adjusted for heteroscedasticity. During periods of crisis when markets are more volatile, the correlation coefficient intends to increase and be biased upward and thus tests based on it traditionally find evidence of contagion. In this chapter, we adopt a formula proposed by Forbes and Rigobon (2002) to adjust for the bias. In other words, we have based our tests for contagion on the unconditional cross-market correlation coefficient.

The main contribution we have made to the existing literature is, for the first time, to employ a non-linear Markov-switching VAR (MS-VAR) model to depict the crisis transmission mechanism. The advantage of the MS-VAR model over the linear models is that it allows us to assume observations are drawn from the different distributions conditional on the state-contingent parameter sets, which is more appropriate for time series variables (i.e. the MPI) especially when there happen special events like currency crises during the period. Instead of imposing an arbitrary threshold, i.e. 10% or a one standard deviation move, the MS-VAR model allows the data itself to reveal what move represents abnormal behaviour and therefore signal a crisis. The estimated results have justified our usage of the MS-VAR model, as the linear hypotheses are rejected even at 1% significant level in all cases.

Our research focuses on the experience of a rapid spread of crises from Thailand to other countries in East Asia in 1997/98. Our initial tests find evidence of contagion between Thailand, Malaysia, Philippines and Indonesia, since the conditional (unadjusted) cross-market correlation coefficients increase significantly during the periods of turmoil in comparison to more stable periods. However, the adjustment
for heteroscedasticity has a significant impact on the correlation coefficients and the resulting tests for contagion. As in Forbes and Rigobon (2002), we have found that all evidence of contagion disappears after we base our tests on the unconditional correlation coefficients for the 1997/98 East Asia crises. The high level of market co-movements between Thailand, Malaysia, Philippines and Indonesia, on the other hand, indicate the close economic linkages-interdependence-amongst them, which exist in all states of the world.

At this point, we think it would be useful to remind ourselves of the purpose of our research that is set up at the beginning of the project. After we discover that the so-called three generation crisis models are not so good at explaining currency crises across the world as they should be, two areas were singled out for this project to make some progress. One is the issue of credibility. It is shown that through the channel of public debt, a current devaluation may have two ambivalent effects on the credibility of a future exchange rate regime. The original theoretical model has been extended, and empirical proof is provided for its two predictions for the first time. The other area is contagion where we adjust our tests for bias due to heteroscedasticity, and also for the first time, employ the non-linear Markov-switching VAR model to analyze the crisis transmission mechanism. Therefore, it is our belief that the goal of the research project has been achieved.

The stock of knowledge that we hope we can contribute to (summarized in Chapter 2 Literature Review and re-discussed later in relevant chapters) has provided us with the necessary background for this study, and has also offered an invaluable reference during our research. The collection of data that has been used in the project has not been a painless process, but the clarity of some of the results has been extremely rewarding. In addition, we have ensured that the methodology we follow in each chapter is appropriate for the problem at hand.

In writing up this thesis, a big value has been placed on transparency. Every effort has been made to provide the reader with sufficient information on the construction of variables and, in general, on every aspect of the work, including any
limitations (which follow in this chapter). Additionally, we have tried to base our claims on facts. Where necessary, we incorporated graphs and tables to support an argument. Appendices as well as footnotes are used to provide further details supplementary to the material in the main text.

Of course, this project is not immune to limitations. Generally, the unavailability of data in some cases forms a natural restriction to our econometric exercises in Chapters 3, 4 and 5. This restriction applies to the majority of empirical studies, the quality of which is judged against the relevancy and adequacy of the data for the question at hand. As mentioned earlier, we have made every effort to obtain reliable and extensive data-sets, which, we believe, contain sufficient information for our research purposes. Furthermore, we aim to constantly update our data-base to include more countries and longer time series, in order to check for possible new trends and test the robustness of our results.

Some specific criticisms also apply to some aspects of our work. In what follows, and before outlining our future research agenda, we briefly present two limitations of this project. In Chapter 3, the main criticism is the rigid specification of our macro econometric model, and the choice of variables is limited by the relevant crisis theories. The reason we have chosen to take this approach is that it is a neat way to link the crisis theories to the empirical works. The implication of this strategy is that there is a cost in terms of accuracy and completeness, since capturing economic reality exactly and fully in an econometric model (i.e., a prediction model) is an extremely demanding task including a much more sophisticated form and more explanatory variables. We have not undertaken this task here, as we aim to see the capabilities of the three generation crisis models in interpreting the causes of different crises across the world—by which they were originally inspired. Therefore, our model is rich enough for our purpose and our approach is suitable and robust for extraction of information relevant to the research.

Another limitation lies in Chapter 5, concerning the assumption of no endogeneity or omitted variables in our model. In other words, the formula that we adopt to
adjust for heteroscedasticity would not be valid if there are exogenous shocks and/or feedback from infected countries to origin country. However, it has been shown in Forbes and Rigobon (2002) that without this assumption, the adjusted correlation coefficient is still a relatively good approximation of the unconditional correlation coefficient if changes in the variances are large and it is possible to identify the country where the shock originates. The intuition is what the simultaneous equation literature calls near identification. It is shown that the variances of the Market Pressure Index (MPI) in nearly all East Asian economies increase by more than 5 times during the 1997/98 Financial/Currency Crises, and the source of the shock is clear—the devaluation of Thai Bhat in early September 1997 is widely regarded as the beginning of the turmoil. Moveover, we only test for contagion from Thailand to other countries in the region in our VAR framework.

We finally turn to the issue of future research. This will develop in two directions. Firstly, we would like to further develop our theoretical model on how current actions affect future policy in Chapter 4. As mentioned before, Drazen and Masson (1994) suggest that the persistent effect of unemployment could provide a channel, while Benigno and Missale (2001) considers the public debt as another factor. Regarding the 1997/98 East Asia Financial/Currency Crises, one would recommend that the domestic credit provided by banking system can also be a bridge to link current actions to the future. An outline of our initial thinking follows.

A similar theoretical model could be set up to show that successful defence of the exchange rate would worsen the fundamentals, by sending a wrong signal to already booming markets to encourage more aggressive lending and thus lead to an increase in the Non-Performing Debts (NPD) in banking system. We would still consider a three-period open economy version of the Barro and Gordon (1983) model, where an inflation surprise, due to an unexpected devaluation, increases output both through the standard price-output effect and through the reduction of repayment of foreign debts. In the model, the domestic bank finances its both one-period (short-term) loans and two-period (long-term) loans to domestic clients by borrowing only short-
term loans from abroad at the beginning of each period.

Following the same logic as in Chapter 4, the monetary authority finds itself facing a dilemma in the first period. On the one hand, resisting a crisis could enhance the authority's reputation so that the foreign deposits would not be pulled out. In that case, the more profitable long-term investment would not be forced to liquidate to compensate the short-term foreign borrowings (Change and Velasco (1998)). Therefore, the market would expect that the parity will be maintained in the second period as everything goes as planned. However, on the other hand, defending the parity would send a wrong signal to the over-booming asset markets, encouraging more risky lendings by moral hazard, which is transformed into an excessive capital accumulation. The excessive investment would inevitably lead to a sharp decrease in returns to capital and thus piling up of non-performing debts in the banking system. Eventually, the bubbles in asset markets would burst and the domestic currency would be forced to devalue.

The recent experience of the emerging economies in East Asia could provide us with some evidence for the two ambivalent effects—a tough policy in a current period may improve the credibility of the exchange rate regime by avoiding the liquidation of long-term investments, or may send a wrong signal to encourage an excessive accumulation of capital in asset markets. But the methodology will have to be left for future discussion.

The second direction we are willing to take is to go deeper to look at our analysis of currency crises in the context of real business cycles. The business cycle is the fluctuation between periods of relatively rapid economic growth alternating with periods of relative stagnation or decline. The periods of rapid growth are associated with increases in productivity, and in consumer confidence. These growth periods usually end with the failure of speculative investments built on a bubble of confidence that bursts or deflates at the end. Some economists argue that the currency crises are rather a symptom of this underlying real (in both senses of the word) malady. We believe that further study of the relationship between the real business cycle and
currency crises in a highly integrated global environment will prove to be a fruitful area of research.
Many macroeconomic or financial variables undergo episodes for a long period in which the behavior of the series seems to change quite dramatically. Such apparent changes in the time series process can result from events such as wars, financial panics, or significant changes in government policies.

But in econometrics, we usually assume that observed data has been drawn from the same distribution conditional on some constant parameter set. As indicated above, it is very unlikely that economic time series can be characterized in such a way.

The standard econometric way to solve that problem is trying to detect the existence of these changes in regime using different types of parameter constancy tests, and then impose dummy variables to account for these changes. But this procedure might be very rigid and may lead to the use of models with too many dummy variables.

As a result, state space models have been launched to tackle this problem. The general state space model consists of two equations: the measurement equation and the transition equation. The Kalman filter of Kalman (1960) and Kalman and
Markov-Switching Model Briefing

Bucy (1961) is an algorithm for generating minimum mean square error forecasts in a state space model. But whereas the Kalman filter is a linear algorithm for generating estimates of a continues unobserved state vector, Hamilton (1989) presented another approach to provide nonlinear inference about a discrete-valued unobserved state vector.

It is presumed in Hamilton model that econometricians can not observe the shifts in regime directly, but instead must draw probabilistic inference about whether and when changes may have occurred based on the observed behavior of the series.

A variable $y_t$ is assumed to be a linear function of a vector of variables $x_t$ with coefficients that depend on the state or regime in period $t$. There are a discrete number of states, $n$. Formally

$$y_t = x_t \beta_{st} + \epsilon_t$$  \hspace{1cm} (A.1)

where $s_t$ is the state in period $t$ which can take one of $n$ possible values, $1, \ldots, n$. Defining $\alpha_i$ as the $n \times 1$ vector with $i$th element equal to one when $s_t = i$ and all other elements equal to zero, we can rewrite the measurement equation (A.1) as

$$y_t = x_t' B \alpha_t + \epsilon_t$$  \hspace{1cm} (A.2)

where $B = [\beta_1 : \cdots : \beta_n]$ and $\text{var}(\epsilon_t) = \sigma^2$.

The state $s_t$ is assumed to follow a first order Markov process so that we can write the transition equation as

$$\Pr(s_t = i|\xi_{t-1}) = \Pr(s_t = i|s_{t-1} = j) = p_{ij}$$  \hspace{1cm} (A.3)

where $\xi_{t-1}$ is a vector representing all the information available at time $t - 1$ which includes lagged values of $y_t$ and $x_t$. The reason for the assumption of Markov chain is that over a deterministic specification for a process with permanent regime changes, one could generate meaningful forecast prior to the change that take into account the possibility of the regime change. The essence of this scientific method
is the presumption that the future will in some sense be like the past. The value \( p_{ij} \) is known as the transition probability of moving to state \( i \) from state \( j \) and is independent of time. If

\[
\Pi = \begin{bmatrix}
p_{11} & \cdots & p_{1n} \\
\vdots & \ddots & \vdots \\
p_{n1} & \cdots & p_{nn}
\end{bmatrix}
\]

then we can rewrite (A.3) as

\[
E(\alpha_t|\alpha_{t-1}) = \Pi \alpha_{t-1}
\]

\[
\alpha_t = \Pi \alpha_{t-1} + \eta_t
\]

(A.4)

where \( \eta_t \) is uncorrelated with \( \alpha_{t-1} \) or \( \xi_{t-1} \) and is not normally distributed.

The Hamilton filter is an iterative algorithm for calculating the distribution of the discrete state variable \( \alpha_t \).

Let \( \alpha_t \) be \( E(\alpha_t|\xi_t) \) with \( i \)th element given by

\[
Pr(s_t = i|\xi_t)
\]

and \( \alpha_{q,t-1} \) be \( E(\alpha_t|\xi_{t-1}) \) with \( i \)th element given by

\[
Pr(s_t = i|\xi_{t-1})
\]

Then the Hamilton filter comprises two recursive equations: obtaining an optimal inference about the current state given the past values of the variable that is to be forecast, the updating equation, \( \alpha_t = h(\alpha_{q,t-1}) \); the prediction equation, using the outcome of the filter to generate future forecasts of this variable, \( \alpha_{q,t-1} = g(\alpha_{t-1}) \).

The Hamilton filter prediction equation follows from (A.4) and is simply

\[
\alpha_{q,t-1} = \Pi \alpha_{t-1}
\]

(A.5)
Through the filter introduced by Hamilton (1989), the optimal inference of \( \alpha_t \) on the basis of the information set in \( t \) consisting of the observed values of \( Y_t \), could be calculated as the following (nonlinear) updating equation.

\[
\alpha_t = \frac{\mathbf{v}_t \boldsymbol{\alpha}_{t|t-1}}{\mathbf{v}_t' \mathbf{a}_{t|t-1}} 
\]

where \( \mathbf{v}_t \) is the \( n \times 1 \) vector with \( i \)th element given by \( f(y_t|\xi_t = i, \mathbf{x}_t, \mathbf{a}_{t-1}) \). The five-step filter weights for each regime the conditional density of the observation \( y_t \), given the population parameters of regime \( m \), with the predicted probability of being in regime \( m \) at time \( t \) given the information set \( Y_{t-1} \).

The optimal inference and forecast for each date \( t \) in the sample can be found by iterating on the above pair of equations.

Moreover, a starting value \( \alpha_0 \) is needed to calculate \( \alpha_t \) through the iterative algorithm. Several options are available for initializing the filter.

If the Markov process is stationary and ergodic, then \( E(\alpha_t) = E(\alpha_{t-1}) \) and from the transition equation (12)

\[
E(\alpha_t) = \Pi E(\alpha_t) 
\]

This can be used to define the vector of unconditional probabilities \( \alpha \) by solving

\[
\alpha = \Pi \alpha 
\]

Then one approach is to set \( \alpha_0 \) equal to \( \alpha \).

Another option is to set

\[
\alpha_0 = \rho 
\]

where \( \rho \) is a fixed \( (N \times 1) \) vector of nonnegative constants summing to unity, such as \( \rho = N^{-1} \cdot 1 \).

The population parameters that describe the time series governed by (A.2) and (A.4) consist of \( \mathbf{B}, \sigma^2 \) and the various transition probabilities \( \Pi \). Collect these
parameters in a vector $\theta$.

The log likelihood function $L(\theta)$ for the observed data $y_T$ evaluated at the value of $\theta$ that was used to perform the iterations can also be calculated as a by-product of this algorithm from

$$L(\theta) = \sum_{t=1}^{T} \log f(y_t|x_t, y_{t-1}; \theta)$$  \hspace{1cm} (A.7)

Then the value of $\theta$ that maximizes the log likelihood can be found numerically using the EM method, developed by Dempster et al. (1977). Given the observed data and an arbitrary initial guess for the value of $\theta$, $\theta^{(0)}$, the EM algorithm begins by calculating the smoothed state probabilities (i.e. the unconditional probability of a particular state, discussed later in this section). With the estimated smoothed state and transition probabilities, the expected complete-data log likelihood function is constructed, which is the "E", expectation part of the algorithm. The log likelihood function is then maximized to obtain an updated parameter estimate. This is the "M", maximization part of the algorithm. The iterative procedure continues until the change between $\theta^{(m+1)}$ and $\theta^{(m)}$ is smaller than some specified convergence criterion. Then the maximum likelihood estimate $\hat{\theta}$ is equal to $\theta^{(m)} = \theta^{(m+1)}$.

Once the unknown parameters of the model $B$, $\Pi$, and $\sigma^2$ have been estimated, it is possible to derive estimates of the state vector based on all the sample information. These smoothed estimators, given by $\alpha_{QT} = E(\alpha_Q|\xi_T)$, represent the best estimate of the probability that the model was in state $s_i$ in period $t$. since Markov-Switching Model includes the possibility that the threshold depends on the last regime, the smoothed probabilities are different from the transition probabilities we discussed above.

Developed by Kim (1994), the backward recursion algorithm used to calculate smoothed inferences, can be written as

$$\alpha_{Q|T} = \alpha_{Q|T} \odot \Pi(\alpha_{Q+1|T} \odot \alpha_{Q+1|T})$$  \hspace{1cm} (A.8)
where $\odot$ is the element-by-element multiplication operator and $\ominus$ denotes the element-by-element division. This algorithm is started with $a_{T_1}$, iterating on backward for $t = T - 1, T - 2, \ldots, 1$.

In practice, the Markov-Switching Vector Autoregressive (MS-VAR) models are mostly used for its flexibility. There are two major types of MS-VAR:

- *shift in the mean (MSM-VAR)*: once and for-all jump in the time series

  \[ y_t - \mu(s_t) = A_1(s_t)(y_{t-1} - \mu(s_{t-1})) + \ldots + A_p(s_t)(y_{t-p} - \mu(s_{t-p})) + u_t; \]

- *shift in the intercept (MSI-VAR)*: smooth adjustment of the times series

  \[ y_t = u(s_t) + A_1(s_t)y_{t-1} + \ldots + A_p(s_t)y_{t-p} + u_t. \]

However, some economists have pointed about that the numerical optimization of a very complicated non-linear function in Hamilton's filter could suffer from several specific types of failures. In particular, the parameter vector may change direction at ever increasing speed toward absurd values, while still increasing log likelihood at each step.
We must update the probability $q_t$ at $t=1$ based on the history $h = [D \text{ or } F, u_1]$. Using Bayes Law:

$$\Pr(A|B \cap u_1) = \frac{\Pr(B|A \cap u_1) \Pr(A \cap u_1)}{\Pr(B|u_1)} \tag{1.1}$$

we have that

$$q_1(F \cap u_1) = \frac{\Pr(T|F \cap u_1)}{\Pr(F|T \cap u_1) \Pr(T \cap u_1)} = \frac{(1 - \rho_1^T)q_0}{\Pr(F|T \cap u_1) \Pr(T \cap u_1) + \Pr(F|W \cap u_1) \Pr(W \cap u_1)} = \frac{(1 - \rho_1^T)q_0}{(1 - \rho_1^T)q_0 + (1 - \rho_1^W)(1 - q_0)}$$
and

\[ q_1(D \cap u_1) = \frac{\Pr(T \mid D \cap u_1)}{\Pr(D \mid T \cap u_1) \Pr(T \cap u_1)} = \frac{\rho_1^T q_0}{\rho_1^T q_0 + \rho_1^W (1 - q_0)} \]

(4.34) now becomes

\[ \rho_1^S(h) = \frac{1}{2} - \frac{1}{2v} \frac{\theta^d}{2m} + \frac{dm}{2} + \Pi(h) \]  

(2)

with \( \Pi(h) \) still given by (4.18). Hence combining (2) and (4.47) we arrive at

\[ \begin{align*}
E_1(x_2(F \cap u_1) - E_1(x_2(D \cap u_1)) &= \frac{d}{2v - dS} \left[ -(1 - q_1(F \cap u_1))\frac{\theta^w d}{2m} - q_1(F \cap u_1)\frac{\theta^T d}{2m} \right] \\
&+ \frac{d}{2v - dS} \left[ Bd + (1 - q_1(D \cap u_1))\frac{\theta^w d}{2m} + q_1(D \cap u_1)\frac{\theta^T d}{2m} \right] \\
&= \frac{d^2}{2v - dS} \left[ B + (q_1(D \cap u_1) - q_1(F \cap u_1))\frac{\theta^w}{2m} + (q_1(D \cap u_1) - q_1(F \cap u_1))\frac{\theta^T}{2m} \right] \\
&= \frac{d^2}{2v - dS} \left[ B + (q_1(D \cap u_1) - q_1(F \cap u_1))\theta^T \right] (\theta^T - \theta^w) \\
\end{align*} \]  

(3)

If \( q_0 = \frac{1}{2} \) (uniform probability), then

\[ q_1(D \cap u_1) - q_1(F \cap u_1) = \frac{\rho_1^T}{\rho_1^T + \rho_1^W} - \frac{1}{2 - \rho_1^T - \rho_1^W} = \frac{\rho_1^T(2 - \rho_1^T - \rho_1^W) - (1 - \rho_1^T)(\rho_1^T + \rho_1^W)}{(\rho_1^T + \rho_1^W)(2 - \rho_1^T - \rho_1^W)} = \frac{\rho_1^T - \rho_1^W}{(\rho_1^T + \rho_1^W)(2 - \rho_1^T - \rho_1^W)} \]
Hence

\[
E_1 \pi_2(F \cap u_1) - E_1 \pi_2(D \cap u_1) = \frac{d^2}{(2v - dS)} \left[ \mu B + \frac{\rho_T^1 - \rho_W^1}{(\rho_T^1 + \rho_W^1)(2 - \rho_T^1 - \rho_W^1)} \frac{(\theta^T - \theta^W)}{2m} \right] \quad (4)
\]

Note that (4) agrees with (7) of Drazen and Masson (1994). If we focus on an equilibrium in which a tough authority never devalues then \(q_1(D \cap u_1) = 0\) and (3) implies

\[
E_1 \pi_2(F \cap u_1) - E_1 \pi_2(D \cap u_1) = \frac{d^2}{2v - dS} \left[ \mu B - q_1(F \cap u_1) \frac{(\theta^T - \theta^W)}{2m} \right] \quad (5)
\]

Then using (4.35) and (4.36), we have

\[
\rho_T^1(F \cap u_1) - \rho_T^1(D \cap u_1) = \frac{d \mu B}{2v} + \left[ E_1 \pi_2(F \cap u_1) - E_1 \pi_2(D \cap u_1) \right] \frac{S}{2v}
\]

\[
= \frac{d \mu B}{2v} + \frac{d^2 S}{2v(2v - dS)} \left[ \mu B - q_1(F \cap u_1) \frac{(\theta^T - \theta^W)}{2m} \right]
\]

(6)
Dynamic Panel Model Estimation (GMM) Briefing

The methodology we use to estimate the dynamic panel model is the Generalised Method of Moments (GMM) (Arellano and Bond (1991)). The general model that can be estimated by the GMM is a single equation with individual effects of the form:

\[ y_{it} = \sum_{k=1}^{p} \alpha_k y_{i(t-k)} + \beta'(L)x_{it} + \lambda_t + \eta_i + \nu_{it}, \quad t = q + 1, \ldots, T_i; \quad i = 1, \ldots, N, \quad (1) \]

where \( \eta_i \) and \( \lambda_t \) are respectively individual and time specific effects, \( x_{it} \) is a vector of explanatory variables, \( \beta'(L) \) is a vector of associated polynomials in the lag operator and \( p \) is the maximum lag length in the model. The number of time periods available on the \( i \)th individual is \( T_i \) and the number of individuals is \( N \).

Identification of the model requires restrictions on the serial correlation properties of the error term \( \nu_{it} \) and/or on the properties of the explanatory variables \( x_{it} \). Only serially uncorrelated or moving average errors are explicitly allowed for the validity of the GMM estimators. The \( \nu_{it} \) are assumed to be independently distributed.
across individuals with zero mean, but arbitrary forms of heteroscedasticity across units and time are possible. This provides the justification of our empirical analysis where both the interest differential relative to USA and the yield curve have different volatilities across the different countries and time periods.

The levels \( x_{it} \) are initially supposed to be correlated with \( \eta_i \) but predetermined even not strictly exogenous, in the sense that \( \mathbb{E}(x_{it} \eta_i) \neq 0 \) for \( s < t \) and zero otherwise. Then only \( x_{i1}, x_{i2}, \ldots, x_{i(s-1)} \) are valid instruments in the differenced equation for period \( s \).

The \((T_i - q)\) equations for individual \( i \) can be written conveniently in the form:

\[
y_i = W_i \delta + \epsilon_i \eta_i + v_i, \tag{2}
\]

where \( \delta \) is a parameter vector including the \( \alpha_k \)'s, the \( \beta \)'s and the \( \lambda \)'s, and \( W_i \) is a data matrix containing the time series of the lagged variables, the \( x \)'s and the time dummies. Lastly, \( \epsilon_i \) is a \((T_i - q) \times 1\) vector of ones. The GMM estimator of \( \delta \) has the form:

\[
\hat{\delta} = \left( \sum_i W_i^* Z_i A_N (\sum_i Z_i^t W_i^*) \right)^{-1} \left( \sum_i W_i^* Z_i A_N (\sum_i Z_i^t y_i^*) \right) \tag{3}
\]

where

\[
A_N = \left( \frac{1}{N} \sum_i Z_i^t H_i Z_i \right)^{-1},
\]

and \( W_i^* \) and \( y_i^* \) denote the first difference transformation of \( W_i \) and \( y_i \). \( Z_i \) is a matrix of instrumental variables and \( H_i \) is a possibly individual specific weighting matrix.

Here we only focus on the one-step estimators which use some known matrix as the choice for \( H_i \). For a first-difference procedure, the one-step estimators use

\[\text{DPD can be used to compute a range of linear GMM estimators of different types of transformations.}\]
Dynamic Panel Model Estimation (GMM) Briefing

\[ H_i = H_i^D = \frac{1}{2} \begin{pmatrix} 2 & -1 & \ldots & 0 \\ -1 & 2 & \ldots & 0 \\ \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & \ldots & -1 \end{pmatrix}, \] (4)

while a two-step estimators use

\[ H_i = \hat{\psi}_i^r \hat{\psi}_i^r, \] (5)

where \( \hat{\psi}_i^r \) are one-step residuals. It is worth noting that, when the \( v_t \) are heteroscedastic, simulations suggest that the asymptotic standard errors for the two-step estimators can be a poor guide for hypothesis testing in typical sample sizes. That is the case in our empirical work where both the interest rate differential and yield curve are more volatile during currency crises. Therefore, we only produce the one-step GMM estimators, with the asymptotic heteroscedasticity-consistent standard errors, based on which the inference tests are more reliable \(^2\).

The assumption of no serial correlation (or just moving average) in the \( v_t \) is essential for the consistency of the one-step GMM first-difference estimators, which instrument the lagged dependent variable with further lags of the same variable. Sargan test of over-identifying restrictions is adopted here to test for the serial correlation in the error term. That is, if \( A_N \) is chosen optimally for any given \( Z_i \), the statistic

\[ S = \left( \sum_i \hat{\psi}_i^r Z_i \right) A_N \left( \sum_i Z_i \hat{\psi}_i^r \right) \]

is asymptotically distributed as a chi-square with as many degrees of freedom as over-identifying restrictions, under the null hypothesis of the validity of the instruments. The Sargan test based on the two-step GMM estimators is reported because

\(^2\)see §3.4, Arellano and Bond (1991), and Blundell and Bond (1998) for further discussions.
only that is heteroscedasticity-consistent.\footnote{Again, Arellano and Bond (1991) provides a comprehensive discussion of the several test procedures.}
Bibliography


Krugman, P. (1999). Balance sheets, the transfer problem, and financial crises. mimeo, MIT.


