ESSAYS ON THE EMPIRICAL MODELLING OF MONEY DEMAND IN
PERIODS OF FINANCIAL LIBERALISATION: THE CASE OF INDONESIA

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by

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To my wife, Rahayu, and my son, Harry
But I shall let the little I have learnt go forth into the day in order that someone better than I may guess the truth, and in his work may prove and rebuke my error. At this I shall rejoice that I was yet a means whereby such truth has come to light.

Albrecht Dürer
Financial liberalisation is widely blamed for the instability of empirical money demand models. A stable money demand function (MDF) is a key element in the formulation of monetary policy. Structural breaks brought about by financial liberalisation can impair the predictability of the impact of changes in money on income, the price level and the interest rate, and render monetary policy less reliable. This in turn can affect individuals' expectations about future policy and ultimately alter the effects of a given policy measure. Thus, obtaining reliable estimates of the parameters of the MDF—through an appropriate modelling strategy that takes into account the structural change induced by financial liberalisation—is a crucial task.

This thesis contains three essays on the empirical modelling of money demand in periods of financial liberalisation. The empirical analysis uses a quarterly time-series dataset on Indonesian money, output, price, interest and exchange rates from 1983:1 to 2001:4. The first essay uses a univariate method to identify endogenous structural breaks in money and the determinants of money demand. The second essay extends this approach to the multivariate case to detect endogenous regime shifts in money demand. The third essay explicitly controls for financial liberalisation in the MDF.

We find evidence of a break occurring in the second quarter of 1991 and coinciding with a major government intervention in the money market known as the Sumarlin shock. We also find evidence of a break occurring in the last quarter of 1997 and coinciding with the severe economic crisis and a government intervention in the money market for which the Sumarlin shock of 1991 is the precedent. Finally, we show how modelling financial liberalisation as a deterministic drift process constitutes an improvement over the standard specifications in terms of yielding more constant and plausible values for the parameters of the MDF.
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CHAPTER 1
INTRODUCTION

During the last three decades, substantial changes occurred in the design and conduct of monetary policy in Indonesia. Before the early 1980s, Indonesia's central bank—like most central banks in the Southeast Asian region—relied on direct control measures such as interest rate ceilings, minimum reserve requirements, and exchange rate controls to implement monetary policy. However, in June 1983, the government embarked upon structural reforms and financial liberalisation which saw the abolition of interest rate controls, liberalisation of the capital market, and removal of entry restrictions of foreign banks into the domestic banking market. The aim was to increase saving and the level and efficiency of investment to support economic growth and achieve higher living standards. In parallel with these reforms, Indonesia's central bank started to rely increasingly on market-based measures to implement monetary policy. Following the interest rate liberalisation, the bank introduced the use of open market operations and started issuing its own debt instruments in the absence of government debt securities. The move was accompanied by the adoption of monetary targeting as the main framework for the conduct and implementation of monetary policy.

Implicit in monetary targeting is the notion of a stable money demand function (Poole, 1970). The existence of a stable and predictable relationship between monetary aggregates, economic activity, prices and interest rate is key to the theory
and application of macroeconomic policies as it provides a reliable link between changes in monetary aggregates and changes in the variables included in the money demand function (Siddiki and Morrissey, 2006; Ericsson, 1998; Ghatak, 1995). While economic theory supports the notion that money demand is likely to be a stable function of income, prices, and interest rates, this stability is predicated on an unchanging institutional environment.

One might expect a priori that the increasing globalisation of the Indonesian financial market and institutional changes in the conduct of monetary policy would have a destabilising impact on the underlying money demand function. The experience of many industrial countries that went through substantial episodes of financial reforms and deregulation during the early and mid-1980s revealed significant shifts in the money demand relationship, which made it difficult to retain intermediate monetary targets. Financial liberalisation, by improving the quality of economic signals, altering the institutional environment, and expanding the array of financial opportunities, creates potential for money demand instability. Interest rates that better reflect the return and riskiness of financial assets can prompt portfolio shifts in money demand. Measures that improve the functioning of financial markets can prompt portfolio shifts as well as alter the sensitivity of money to changes in income and interest rates. More generally, measures that promote financial development may result in the availability of new, attractive assets (for instance foreign assets if external capital flows are liberalised) leading to gradual portfolio shifts away from monetary assets (Tseng and Corker, 1991).
Changes in the institutional environment brought about by financial liberalisation are also widely perceived to be responsible for observed cycles in the income velocity of money (Bordo and Jonung, 1987; 1990). This places some practical limits on the notion of stability. In particular, it may be impossible to identify a stable money demand relationship for all time, but it may be possible to estimate a relationship for which the key parameters are reasonably constant over the sample period in question. This would provide some grounds for concluding that monetary developments remain predictable in the future, at least over the duration of the policy horizon (Tseng and Corker, 1991).

From an econometric perspective, the question of the stability of money demand has been intensively investigated using a variety of econometric techniques. Currently, the most popular framework for examining the behaviour of money demand is the cointegration framework. In this framework, the stationarity of money demand instead of the stability of money demand is evaluated. Earlier tests of the stability of money demand typically centered on whether the coefficients estimates were stable, that is, not subject to structural breaks (see, for instance, Aghevli et al., 1979). These tests did not consider the underlying time series aspects of the variables in money demand or the time series properties of their joint relationship prior to estimation as is now common in the cointegration framework.

From a policy perspective, finding that money demand is nonstationary is problematic since the ability of policymakers to prevent money market disruptions depends on whether a money demand relationship that is predictable can be iden-
tified. If money demand is nonstationary then the monetary authorities may not be able to use such a relationship to target money growth with accuracy. In such case, the central bank's ability *ex ante* to prevent a money market disequilibria from affecting the economy could be curtailed. Finding that money demand is not stable is also problematic since proper policy conduct may *ex post* be misguided in light of a structural break and there may be a period of learning before policy is readjusted to its proper course (Breuer and Lippert, 1996).

The issues of stationarity and structural stability, however, should not be treated independently. It is possible that tests that find money demand is nonstationary may be flawed because the estimations procedures used have not considered a structural break. Failure to take account of structural change may bias the results in favour of nonstationarity.

The thesis contains three essays on the empirical modelling of money demand in periods of financial liberalisation: the case of Indonesia. Indonesia, having experienced financial liberalisation as well as a banking crisis of real gravity, and given its monetary framework, is a well chosen subject for money demand analysis. Its population, four times Britain's, and its strategic significance makes it a country worthy of study in its own right. However, the Indonesian case also exemplifies the experiences of structural reforms, policy changes and financial innovations of many developing countries in Asia and around the world. The methods and key findings of the thesis should therefore prove useful in improving our understanding of the effects of structural change on money demand not only for Indonesia but
also for a variety of developing economies.

The empirical analysis uses a quarterly time-series dataset on Indonesian money, output, price, interest and exchange rates from 1983:1 to 2001:4. Our theoretical and empirical starting points are provided in chapter 2. We review theories of money demand and their econometric formulations. We also discuss the implications of financial liberalisation on the stationarity and structural stability of the money demand function.

In chapter 3 we present an overview of financial liberalisation in Indonesia. We review significant financial reforms and discuss changes in monetary policy. We also analyse the development of the financial sector and include quantitative measures of its growth and structural change.

In chapter 4 we use a univariate method to search for endogenous structural breaks in monetary aggregates and the determinants of money demand in Indonesia. We use the procedure outlined in Vogelsang and Perron (1998). The null hypothesis is that of a unit root against the alternative of stationarity with a break of unknown timing in the intercept and slope of the trend function of the series.

In chapter 5, we apply a multivariate method to search for endogenous regime shifts in Indonesian money demand. We use the cointegration test of Gregory and Hansen (1996), which allows for a structural break of unknown timing in the money demand function.

In chapter 6, the content of which has been published as a journal article:

we explicitly model financial liberalisation as a deterministic drift process. We follow the ARDL approach of Pesaran, Shin and Smith (2001), which avoids pretesting of the order of integration of the series in the money demand function.

Finally, we conclude in chapter 7 by assessing the empirical relevance of our results and discussing their implications for future research.
CHAPTER 2
A REVIEW OF THEORIES AND EMPIRICAL MODELS OF MONEY DEMAND

The purpose of this chapter is to offer a review of economic theories and econometric formulations of the demand for money. From an empirical point of view, the theoretical literature implies that the aggregate demand for money (the money demanded by households, firms and the government) depends positively on a scale (transaction) variable; negatively on one or more opportunity costs variables (usually some interest rates and/or the inflation rate). Differences between competing theories relate to the use of the appropriate scale and opportunity cost variables. The chapter finishes with a discussion on financial liberalisation and its empirical implications in terms of structural change and model misspecification.

2.1 Economic theories of money demand

Money serves four major functions—medium of exchange, store of value, unit of account, and source of deferred payment. Individuals hold money balances for various reasons—transactions, precaution and speculation. Each money demand theory places particular emphasis on one or more of these functions and underlying motives. This following discussion briefly reviews the theoretical literature on money demand. The reader is referred to Cuthbertson (1985), Goldfeld and Sichel (1990), and Laidler (1993) for comprehensive surveys.
2.1.1 The quantity theory and Keynes's theory

One early approach to the demand for money is the quantity theory of money, put forward by Fisher (1911), and the so-called Cambridge economists, Marshall (see Whitaker, 1975) and Pigou (1917). Both Fisher and the Cambridge economists are concerned primarily with money as a means of exchange and therefore provide models of the transactions demand for money.

Fisher (1911) analyses the institutional details of the payments mechanism and therefore concentrates on the velocity of circulation of money. The basis of Fisher's theory is an identity linking the value of sales with the amount of money which changes hands. If \( Y \) equals the volume of transactions and \( P \) equals the average price level then \( PY = \) the value of transactions undertaken. The number of times the money changes hands, that is the velocity of circulation of money \( V \) multiplied by the fixed stock of money \( M \), must be equal to the value of transaction \( (MV = PY) \). One crucial assumption is that \( V \), being determined by technological and/or institutional factors, is relatively constant. More specifically, it is assumed that the payments mechanism is such that the velocity of circulation is constant in the short run and varies slowly and in a predictable way in the long-run, as payments mechanisms in the economy change.

The Cambridge economists are concerned with the determinants of the individual's desired demand for money. Although the transactions motive is still the main determinant of desired money holdings, they argue that one alternative to holding money is to hold an interest bearing asset called a bond. The higher the
rate of returns on bonds, relative to the marginal utility from money, the more
individuals are encouraged to switch some of their money holdings into bonds.
Assuming a proportionate relationship, one can write $M^d = KPY$. If one assumes
money market equilibrium then $M^s = M^d = KPY$. Hence $V = 1/K$ and since
$K$ depends upon the interest rate and wealth, the Cambridge economists consider
that the velocity of circulation is neither constant in the short run nor in the long
run.

Keynes (1936) broadly accepts the view of the Cambridge economists concerning
the transactions demand for money. He also introduces two other motives
for holding money — precautionary and speculative. The main predictions from
Keynes's theory are, first, that individuals do not hold a diversified portfolio of
assets: they would hold either all bonds or all money. Second, there is a downward
sloping aggregate demand for money function with respect to the interest rate.
Finally, the theory predicts the possibility of a liquidity trap, meaning that under
certain circumstances, the interest elasticity of the demand for money can become
infinite. Keynes's model of the demand for money has the important implication
that velocity is not constant but instead is positively related to interest rates, which
fluctuate substantially. His theory also rejects the constancy of velocity because
changes in people's expectations about the normal level of interest rates would
cause shifts in the demand for money that would cause velocity to shift as well.
2.1.2 Transactions theories

Transactions theories, set forth in the inventory theoretic models of Baumol (1952) and Tobin (1956) and in the later uncertainty version of Miller and Orr (1966, 1968) emphasise the role of money as a medium of exchange.\(^1\) Money is viewed essentially as an inventory held for transactions purposes. Transactions costs of switching between money and other liquid financial assets justify holding such inventories even though other assets offered higher yields.

Inventory-theoretic models provide precise forms of the demand for money function. In its simplest form, the Baumol (1952) and Tobin (1956) model assumes that individuals receive a known lump sum payment of \(Y\) (paid in bonds) at the beginning of a period and spend this amount uniformly over the period. The optimal average cash balance is given by

\[
M = \left( \frac{bY}{2r} \right)^{1/2},
\]  

(2.1)

where \(r\) is the interest rate on bonds and \(b\) is the brokerage charge or fixed transactions cost for converting bonds into cash. The model therefore yields a constant interest elasticity of \(1/2\). It is worth mentioning that the precise form of (2.1) and constancy of the interest and income elasticities is dependent on the assumed payment mechanism.

Miller and Orr (1966, 1968) provide another application of inventory theory to the transactions demand for money, which can also be interpreted as a model

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\(^1\) Other approaches based on the transaction motive include "cash-in-advance" models and the money in the utility function approach. The interested reader is referred to Lucas (1980) and Feenstra (1992).
of the precautionary motive for holding money since there is a minimum allowable
money holding below which a penalty must be paid. The key difference with the
Baumol-Tobin model is the assumption that cash flows are stochastic.

In its simplest form, the model assumes that cash flows follow a random walk
without drift in which, in a given time interval (say 1/t of a day, where t is the
frequency of transactions), there is a probability of a positive or negative cash
flow of m dollars (m is the amount the cash balance is expected to alter with a
probability of $\frac{1}{2}$). Given a lower bound below which money balances cannot drop
(normalised to zero), the optimal policy consists of an upper bound, h, and a return
level, z. Whenever money balances reach the lower bound, z dollars of bonds are
converted into cash; whenever the upper bound is reached, h - z dollars of cash are
converted into bounds. The optimal average cash balance (assuming a binomial
distribution for the net cash drain with a zero mean) is given by

$$M = \frac{4}{3} \left[ \frac{3b}{4r} \sigma^2 \right]^{1/3},$$

where $\sigma^2$ is the daily variance of changes in cash balance ($\sigma^2 = m^2t$).

Thus, like the Baumol-Tobin model, the Miller-Orr model also yields a constant
interest rate elasticity, with a value of 1/3 instead of 1/2. Since the scale variable
is $\sigma^2$, the income (or transactions) elasticity is ambiguous. It is either 2/3 if one
considers the size of each cash flow, or 1/3 if one considers the rate of transactions.

Two features of these models can be used to explain the impact of financial
innovation on the demand for money. The first, common to both the certainty
and uncertainty approach, is the role of transactions costs in determining the de-
mand for money. Financial innovation, by lowering the converting other assets into money, allows money holders to keep smaller money balances, meeting transactions needs by more frequent transfers of funds from higher-yielding alternatives. The second is the emphasis in the Miller-Orr version of the uncertainty of cash receipts and disbursements as an important determinant of money demand. Greater variance of cash flow leads to larger precautionary holdings of money holdings not to be caught short. Innovations in cash management techniques are viewed as reducing the variance of cash flow, thereby allowing firms to reduce their precautionary balances (Judd and Scadding, 1982).2

2.1.3 Portfolio theories

Portfolio theories of money demand interpret the demand for money more broadly as part of a problem of allocating wealth among a portfolio of assets which include money. Money and other assets are viewed as alternative ways of holding wealth, each yielding some mix of income and non-pecuniary service flows. In the case of money, these services include the ease of making transactions as well as other services with non-pecuniary yields such as liquidity and safety.

The individual wealth-holder in Tobin (1958) allocates her portfolio between money and bonds. There is a trade-off between the net income receivable on bonds and the risk associated with the total portfolio of bonds and money (perfectly liquid with zero return). For any given level of wealth it is possible to calculate the impact of a change in interest rate on both the interest income and the capital gain or

2 For extensions of the Baumol-Tobin and Miller-Orr approaches the reader is referred to Barro and Fisher (1976) and Cuthbertson and Barlow (1991).
loss associated with the holding of different quantities of bonds and money. Under the assumption of expected utility maximisation, the optimal portfolio mix can be shown to depend on the individual's degree of risk aversion and wealth, and the mean-variance characteristics of the risky asset's return distribution.

Tobin's analysis implies a negative interest elasticity of money demand providing another "rationalisation" of Keynes's liquidity preference hypothesis (Judd and Scadding, 1982). The approach has, however, two shortcomings. First, money does not actually have a yield that is riskless in real terms. Second, in many economies there exists a number of riskless assets paying a positive (or higher) rate of return than money. Under such conditions the model implies that individuals would only hold such assets and, therefore, that money is a "dominated asset" (Barro and Fisher, 1976).

As in Tobin's portfolio approach, Friedman (1956) offers a "restatement" of the quantity theory which regards the primary role of money as a form of wealth. According to Friedman, the demand for money can be analysed in the same way as the demand for any asset. Money yields utility in the form of a flow of services to the holder. Thus, the demand for money function contains a budget constraint (either permanent income or wealth); the price of the commodity itself (money) and its substitutes and complements (the rates of return on money and other assets categorised into three types: bonds, equity, and goods); other variables determining the utility attached to the services rendered by money relative to those provided by other assets; tastes and preferences. Friedman expresses his formulation of the
demand for money as follows:

\[
\frac{M^d}{P} = f(Y_p, w, r_b - r_m, r_e - r_m, \pi_e - r_m, u),
\]  

where \(M^d/P\) is the demand for real money balances; \(Y_p\) is a measure of wealth known as permanent income (the present discounted value of all expected future income); \(w\) is non-human wealth/wealth; \(r_m\) is the rate of return on money, influenced by two factors: the services provided by banks on deposits (e.g. provision of receipts in the form of cancelled checks, automatic paying of bills) and the interest payment on money balances; \(r_b\) and \(r_e\) are the expected rates of return on bonds (abstracting from the possibility of capital gains and losses) and equity (abstracting from the effects of on equity prices of changes in interest rates and the rate of inflation), respectively; \(\pi_e\) is the expected inflation rate, included as the rate of return on real assets; \(u\) is a portmanteau symbol standing for other variables affecting the utility attached to the services of money and also includes tastes and preferences.

Tobin's theory supports Keynes's proposition that the demand for money is sensitive to interest rates, suggesting that velocity is not constant and that nominal income might be affected by factors other than the quantity of money. In contrast, Friedman's analysis implies that changes in interest rates should have little effect on the demand for money. Friedman argues that any rise in the expected returns on other assets as a result of the rise in interest rates would be matched by a rise in the expected return on money. He also suggests that random fluctu-
ations in the demand for money are small and that the demand for money can be predicted accurately by the money demand function. These views when combined mean that velocity is predictable—even though it is no longer assumed to be constant—implying that a change in the quantity of money produces a predictable change in aggregate spending as in the quantity theory of money.

2.1.4 Transactions with a shopping time constraint

Microeconomic transactions models attempt to explain the holding of money for transactions purposes within general equilibrium models. The most prominent models of this kind are proposed by McCallum (1989) and McCallum and Goodfriend (1992). These models analyse the demand for money in terms of the shopping time saved in carrying out transactions through the use of money (as opposed to barter). Shopping time saved has value since it can be used to earn income or to obtain utility from other uses.

The model in McCallum and Goodfriend (1992) has a representative household maximising present and future utility for the consumption of goods and leisure subject to the usual intertemporal budget constraint. Consumption goods can be obtained in exchange for income only by shopping for them; the household faces a shopping time constraint given by:

\[ s_t = \psi(c_t, m_t) \]  

(2.4)

The shopping time constraint is the amount of time required to carry out purchases and is increasing in total transactions, \( c \), but decreasing in the quantity of real money balances carried on shopping trips, \( m \). It follows that a decision to hold
more money now, *ceteris paribus*, reduces shopping time, leaving more time for current leisure and/or increased labour supply and future real income. The model yields the following money demand relationship:

$$\frac{M_t}{P_t} = L(c_t, i_t),$$

(2.5)

where $i_t$ is the nominal interest rate, which reflects the cost of holding money rather than bonds. For reasonable specifications of the utility and shopping time functions, $L(\cdot, \cdot)$ will be increasing in its first argument and decreasing in the second.

Arrau et al. (1995) use a framework similar to McCallum and Goodfriend in which financial innovation enters the shopping time (transactions technology) function in the form of a technological parameter. For every unit of the consumption good bought by the household, "$H$" units of the consumption good must be spent, which is represented by:

$$H(m_t, c_t, \theta_t) = \frac{1}{c_t^{1-\phi}} h \left( \frac{m_t}{c_t^{\phi}}, \theta_t \right),$$

(2.6)

where $m$ is real balances and financial innovation is represented by changes of the technological parameter $\theta$ through time.

The cost function is increasing in $\theta_t$; a reduction in $\theta$ reduces the cost of transactions and is associated with (positive) financial innovation. Financial innovation in this formulation is analogous to productivity growth in a production function. The parameter $\phi$ represents the degree of scale economies in transactions. For all $\phi < 1$ money holdings required per unit of consumption (for a given $H$) are decreasing in the level of $c$. A closed form for the demand for money can be achieved
assuming $h$ has the following form (omitting time subscripts):

$$h \left( \frac{m}{c^\alpha}, \theta \right) = K\theta + \frac{1}{\alpha} \left\{ \frac{m}{c^\alpha \theta} \log \left( \frac{m}{c^\alpha \theta} \right) - \frac{m}{c^\alpha} \right\}, \quad (2.7)$$

where $K$ denotes a constant. This formulation yields the following money demand function:

$$\log(m_t) = \log(\theta_t) + \phi \log(c_t) - \alpha \frac{i_t}{1 + i_t}, \quad (2.8)$$

where $-\alpha$ is the semi-elasticity of the interest rate and $\phi$ is the elasticity of consumption. Equation (2.8) indicates that $\frac{i_t}{1 + i_t}$ is the relevant measure of opportunity cost. However, the most commonly used $i$ can be justified by changing the timing of the household problem. Furthermore, since this is a model of household money demand the relevant scale variable is consumption; inclusion of firms and government in the analysis would lead to the use of other scale variables.

### 2.2 Econometric formulations

There is a large body of empirical work on the demand for money. Initially, this work was confined primarily to industrial countries, especially the United States and the United Kingdom. More recently, there has been a renewed interest among industrial and developing countries alike (see for example Siklos, 1993; Arrau et al., 1995; Ericsson and Sharma, 1998; Dekle and Pradhan, 1999; Brissimis et al., 2003, Binner et al., 2004). One of the significant contributors of the recent boom in empirical research on money demand is the major advancements made in time series econometrics from the late 1980s onwards, which have motivated researchers to revisit the empirical models built previously (Sriram, 1999).
This discussion provides an historical background to the debate on the stability of the demand for money starting from the mid-1970s, as well as a brief review of the various models and procedures which have been used to estimate money demand.

2.2.1 Historical development

Judd and Scadding (1982) noted that, prior to 1973, the evidence on the demand for money was interpreted as showing that a stable money demand did exist. This evidence has been extensively surveyed elsewhere (see for example Laidler, 1993). The substantial body of empirical research which had accumulated over the postwar period sought to discriminate among the competing hypotheses suggested by the different theories of the demand for money surveyed above. However, this literature showed that it was not possible to distinguish empirically, with any degree of precision, between competing hypotheses about the demand for money (Judd and Scadding, 1982).

Most of the research up until the 1980s was carried out using partial adjustment models (PAMs) in which it is assumed that because of adjustment costs, lagged money needs to be included in the money demand function for the desired level of money holdings to match the actual level. Using the partial adjustment framework, Goldfeld (1973) found evidence of a stable demand function for US narrow money (M1) over the period 1952:2-1973:4, which was positively related to real GNP, negatively related to a representative short-term market rate (Treasury bill or commercial paper rate) and the rate on savings deposits, and positively
related to lagged money. This specification became the conventional demand for money function used by economists but was shown unable to explain a new set of events which emerged in the United Kingdom and the United States in the mid-1970s. The latter included changes in regulations concerning interest rate ceilings on the deposits of commercial banks, innovations in short-term financial markets associated with improvements in corporate cash-management techniques, increases in the rate of inflation and interest rate compared with previous postwar experience, and a greater emphasis on monetary targeting by the central bank.

In the UK, Haache (1974) showed that his money demand function was unable to predict accurately outside of sample for the period after 1971, and recorded significant negative forecasting errors for broad and narrow money. In the US, Goldfeld (1976) came to the same conclusion showing that forecasts seriously overpredicted real money balances from 1974. Goldfeld labeled this phenomenon "the case of the missing money". Other industrial countries experienced a similar problem.

The systematic overprediction of real money balances (or underprediction of velocity) by standard money demand functions stimulated an intense search for a stable money demand function which dominated the research agenda of the next twenty years or so. This search initially took two directions. The first direction focused on whether an incorrect definition of money could explain why the demand for money function had become unstable. For example, Garcia and Pak (1979),

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3 The forecasts were out-of-sample dynamic simulations, which used actual interest rates and income but last period's predicted money balances as the lagged dependent variable.
Wenninger and Sivesind (1979a, 1979b), and Tinsley et al. (1978, 1981) looked for instruments that had been incorrectly left out of the definition of money used in the US money demand function, most notably overnight repurchase agreements (RPs).\textsuperscript{4} These authors argued that for many large corporations, transactions costs between money and RPs were so low that RPs were effectively money.

The second direction of search for a stable money demand function was to look for new variables to include in the money demand function. Hamburger (1977) specified the demand for real M1 in the US as a function of real income, lagged M1, and three rates of return-the commercial bank savings deposit rate, the US government bond rate, and the dividend-price ratio (as a proxy for the rate of return on equities and thus on physical capital).\textsuperscript{5} Although this specification was found to be stable, its stability also depended on the assumption that the income elasticity of the demand for money was unity, which led to strong criticisms (see for example Hafer and Hein, 1979).

Heller and Khan (1979) expanded the opportunity costs of money in their money demand function to include the entire term structure of interest rates and found that this produces a stable money demand function. However, Porter and Mauskopf (1978), who re-estimated the Heller and Khan equation and dynamically simulated it over 1973:3-1977:4, found that the cumulative prediction error by 1977:4 was on the same order of magnitude produced by the Goldfeld equation.

\textsuperscript{4} These are one-day loans with little default risk because they are structured to provide Treasury bills as collateral.

\textsuperscript{5} Including the dividend-price ratio in the list of interest rate arguments allowed for the possibility of substitution between money on the one hand and real capital and commodities on the other.
The problems associated with conventional money demand functions worsened in the 1980s in the face of a surprising slowdown in M1 velocity in the US, which conventional money demand functions could not predict. The relative stability of M2 velocity at the time suggested that money demand functions in which the money supply was defined as M2 might perform substantially better than those in which the money supply was defined as M1. For instance, Small and Porter (1989) found that their M2 money demand function performed well in the 1980s, with M2 velocity moving quite closely with the opportunity cost of holding M2, which was defined as the market interest rate minus an average of the interest paid on deposits and financial instruments that make up M2. However, in the early 1990s, M2 growth underwent a dramatic slowdown, which traditional money demand functions again could not explain.

A new strand of research identified a number of theoretical and econometric problems associated with the partial adjustment framework. PAMs were shown to suffer from specification problems and highly restrictive dynamics (see for example Cooley and LeRoy, 1981; Goodfriend, 1985; Hendry, 1979). To counter these problems, two major solutions were proposed—modifying the theoretical base and improving the dynamic structure. The former led to BSMs (see for example Laidler, 1984; Cuthbertson and Taylor, 1987; Milbourne, 1988), and the latter to ECMs.6 ECMs have been used in a single equation context (see for example Mehra, 1993; Hendry and Ericsson, 1991a, 1991b; Baba et al., 1992; Hess et al., 1994; Janssen, 1996; Thomas, 1996) or in a vector error correction model (VECM) system context.

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6 Hendry et al. (1984) showed that PAMs and BSMs form special cases of ECMs.
While BSMs ran into empirical difficulties, ECMs, because of their success, rapidly became the primary tool to analyse the demand for money. Rose (1985) obtained an ECM for US narrow money demand with constant parameters over the 1970s, thus showing how Goldfeld’s episode of missing money of the mid-1970s resulted from dynamic misspecification. However, Rose’s model broke down in the 1980s. Using a similar empirical model but accounting for financial innovations, Baba et al. (1992) found a money demand function with a constant parameterisation for 1960-1988. MacDonald and Taylor (1992) found a constant US money demand function for annual data over 1874-1975. A similar chronology of ECMs also exists for money demand in the UK. Relevant papers include Hendry and Ericsson (1991a, 1991b), Ericsson et al. (1994), and Ericsson et al. (1998).

2.2.2 Partial adjustment models

The partial adjustment framework assumes that individuals constantly adjust their current money holdings ($M_t$) to the desired long-run equilibrium ($M^*_t$), which is given and depends upon a vector of variables (e.g. interest rates, income). Agents face the costs of being out of equilibrium, which can be explained in terms of interest income foregone or inability to purchase goods. Assuming the costs of being above and below equilibrium are equal, these costs are represented by the quadratic

\[ \text{costs} = K (v - a)^2 \]

where $v$ is the deviation from equilibrium and $a$ is the adjustment parameter. The adjustment parameter $a$ represents the speed of adjustment back to equilibrium. The costs function is quadratic, implying that the adjustment process is smooth and that deviations from equilibrium are penalized.

\[ \text{costs} = K (v - a)^2 \]

The adjustment of actual to desired money holdings can be in real terms as in the real partial adjustment models (RPAMs), with the lagged money variable in the form of $M_{t-1}/P_{t-1}$ (where $M$ are nominal money balances and $P$ is the price level), or in nominal terms as in the nominal partial adjustment models (NPAMs), with lagged money in the form of $M_{t-1}/P_t$. The reader is referred to Cuthbertson (1985, chapter 6) and Boughton (1992) for a discussion of the deficiencies (e.g. overshooting and implausible lags of adjustment) associated with both types of models.
term $a(M_t - M_t^*)^2$, where $a$ is the cost per unit of disequilibrium. Individuals also face the costs of adjusting their portfolio of assets, which is also represented by a quadratic term $b(M_t - M_{t-1})^2$. Given $M_t^*$ and $M_{t-1}$ an individual chooses $M_t$ to minimise (one period) total costs which is the sum of the two quadratic terms. The first order solution to this minimisation problem gives

$$M_t - M_{t-1} = \gamma(M_t^* - M_{t-1}), \quad (2.9)$$

where $\gamma = \frac{a}{a+b}$.

First, assuming that an increase in the demand for money is always met by an increase in supply (as for example under a constant interest rate target by the central bank) then the short-run desired demand derived in equation (2.9) is equal to actual money balances and the money demand function can be estimated. Second, assuming that the long-run desired demand for money $M_t^*$ has a linear form and depends on current income $Y_t$, and the expected return on alternative assets $R_t^e$, the long-run money demand function can be written as

$$M_t^* = m_yY_t - m_rR_t^e + u_t \quad (2.10)$$

where $u_t$ is an additive stochastic error fulfilling the usual stationarity conditions.

Substituting equation (2.10) in equation (2.9) and rearranging yields

$$M_t = b_1Y_t + b_2R_t^e + b_3M_{t-1} + u_t, \quad (2.11)$$

where $b_1 = \gamma m_y$, $b_2 = -\gamma m_r$, and $b_3 = 1 - \gamma$ are the estimated coefficients and $u_t = \gamma u_t$. 
Equations such as (2.11) have generally been estimated with ordinary least squares (OLS) using the Cochrane-Orcutt technique to adjust for serial correlation. In general, all estimates showed very low short-run elasticities for income (around 0.1) and interest rates (around -0.05), and a coefficient close to unity for the lagged dependent variable. In contrast, the long-run interest elasticity was often found to be much higher (around -0.3).

The partial adjustment framework has raised a number of criticisms, both on theoretical (see for example Goodfriend, 1985) and empirical grounds. Econometric problems include serial correlation, heteroskedasticity, simultaneity bias, model misspecification, and spurious regressions due to the nonstationarity of the data (Goodfriend, 1985; Yoshida, 1990).

2.2.3 Buffer stock models

There are a number of fairly distinct applied approaches to buffer stock money (see for surveys Laidler, 1984; Milbourne, 1987; Mizen, 1994). The following discussion gives a brief account of some general considerations underlying the idea of money as a buffer stock drawing from Mizen (1997).

Buffer stock models have two common basic assumptions, which are an exogenous money stock, that is, the money stock is primarily influenced by supply factors, and a disequilibrium real balance effect. This is in contrast with PAMs, which assume that the money stock is demand determined in the short run and that the money market is in equilibrium with endogenous income, interest rates and price level variables adjusting to clear the market. BSMs explain the demand
for money in the context of individual optimisation and the microeconomics of adjustment in the market for money. Expectations are introduced into the analysis such that unexpected events in the present or expected events in the future induce deviations from long-run equilibrium. Since adjusting money balances is costly, it can be optimal for an individual, when faced with a deviation of money balances from long-run equilibrium, to allow that departure to persist in the short run. Hence, BSMs reject the notion that all individuals hold their long-run level of money balances at all times and that the money market in aggregate is continuously in equilibrium. Instead, BSMs adopt the view that temporary departures can be rational and optimal both at the individual and the market level.

Carr and Darby (1981), starting from the view that money is subject to exogenous shocks, argue that money balances serve as a "shock absorber" or buffer stock while money holders choose their new portfolios. They implement this idea by starting with the real partial adjustment model and modifying it with "surprise money", which allows money supply shocks and temporary income to push real money balances off the individual demand for money function in the short run. The resulting functional form is an equation which allows for positive and negative variations to individual money balances in response to unexpected events in the short run. This equation is given by

\[
m_t = (1 - \lambda)m_{t-1} + \lambda m^*_t + p_t + b^T_t + f(m_t - m^*_t) + u_t,
\]

\[ (2.12) \]

\[ ^8 \text{Buffer stock models based on the inventory principle (see, for example, Akerlof and Milbourne, 1980; Milbourne, 1987) are surveyed by Mizen (1994) and not reviewed here. These models have much in common with the Miller-Orr approach discussed above.} \]
where \((m_t - m^*_t)\) are unexpected nominal money supply shocks defined as actual money supply, \(m_t\), minus anticipated money supply, \(m^*_t\); \(y^*_t\), is transitory income defined as current minus permanent income and included since any unexpected variations to income are temporarily held as money;\(^9\) \(m^*_t\) are long-run desired money holdings described as a function of permanent income and interest rates; \(\lambda\), \(b\), and \(f\) are coefficients derived from the parameters of the model, and \(u_t\) is the white noise error term. Carr and Darby recognise that \(u_t\) is correlated with \((m_t - m^*_t)\) and OLS estimation yields inconsistent estimates of the parameters. Hence, they use an instrumental variables procedure, namely two-stage least squares (2SLS).

Although Carr and Darby did not examine the issue of the missing money episode, one can see from equation (2.9) that in periods of relative "calm" the partial adjustment model would be expected to perform reasonably well, but when the expectational errors become more important in more "turbulent" times (when the coefficients \(f\) and \(b\) are significantly different from zero), they would be the cause of serious forecast errors as observed in the mid-1970s. However, both the estimation procedure of Carr-Darby and their conclusions have been extensively criticised (Cuthbertson and Taylor, 1986; MacKinnon and Milbourne, 1988).

Cuthbertson and Taylor (1987) propose a model in which departures from current period equilibrium result not just from unexpected events but from anticipated events ahead of the current time period, such as expectations of future

\(^9\) Carr and Darby (1981) argue that permanent income is the appropriate concept for the long-run demand for money and that transitory income will temporarily be kept as money until adjustment can occur.
monetary policy. The individual faces the cost of being out of equilibrium and the cost of adjusting balances in the future as well as the present. Money holdings are composed of planned components based on the minimisation of an intertemporal cost function and unplanned components resulting from unanticipated shocks. This approach assumes multi-period quadratic costs of adjustment. The resulting money demand function is given by

\[ m_t = \lambda m_{t-1} + (1 - \lambda)(1 - \lambda D) \sum_{i=1}^{\infty} (\lambda D)^i E_{i-1} m^*_t + m^*_t + e_t, \quad (2.13) \]

where \( \lambda \) is a parameter of the model, \( m^*_t \) are unexpected shocks to money holdings, \( D \) is a known discount factor, \( m^* \) are long-run money balances, and \( e_t \) is the white noise error term.

Milbourne (1987, 1988) summarises the various theoretical and empirical criticisms raised by the buffer-stock approach. These criticisms help explain why "subsequent work favours the ECM interpretation" (Cuthbertson, 1997).

2.2.4 Error correction models

The important feature of the error correction model is that it provides significant emphasis on the time series characteristics of the data and the underlying data generating process. Unlike the PAM and BSM, which severely restrict the lag structure by imposing implausible lags of adjustment or relying solely on economic theory, the ECM allows economic theory to define the long-run equilibrium while the short-term dynamics is determined from the data. The following discussion on ECMs draws from Alogoskoufis and Smith (1991).
The current popularity of ECMs is largely due to David Hendry, whose work was influenced by Phillips (1954, 1957) and Sargan (1964). One of the most influential of Hendry's contributions on error correction models is Davidson et al. (1978), in which the ECM form for the relationship between consumers expenditure and income is introduced.\(^\text{10}\) In Hendry's work (see for example Hendry et al., 1984), the best known representation of the ECM in a simple money demand model with three variables, \(m - p\) (real money) \(y\) (income) and \(R\) (interest rate) is a log-linear equation of the form

\[
\Delta(m - p)_t = \alpha + \beta_1 \Delta y_t + \beta_2 \Delta R_t - \gamma((m - p)_{t-1} - y_{t-1} - R_{t-1}) + \varepsilon_t \tag{2.14}
\]

where \(\varepsilon_t\) is a white noise error term.

The static long-run solution of this equation (when \((m - p)_t = (m - p)_{t-1} = m - p, y_t = y_{t-1} = y\) and \(R_t = R_{t-1} = R\) is

\[
(m - p) = \alpha/\gamma + y + R
\]

Since most economic time-series are highly trended with stationary growth rates, that is, they are integrated of order one, \(I(1)\), the two sides of equation (2.14) are of different orders of integration, with \(\Delta(m - p)_t\) stationary (integrated of order zero) and \((m - p)_{t-1}, y_{t-1}\) and \(R_{t-1}\) all \(I(1)\), unless the linear combination \((m - p)_{t-1} - y_{t-1} - R_{t-1}\) is also stationary. In general, linear combinations of \(I(1)\) variables are also \(I(1)\), but if they are \(I(0)\), then the variables are said to be cointegrated.\(^\text{10}\) Although the estimated relationship is given a "feedback" interpretation, the term "error correction" is not used in the paper. The term is first introduced in Hendry (1980).
Engle and Granger (1987) proposed a very popular procedure for testing cointegration, in which residuals from a static regression of integrated variables are tested for having a unit root. The static regression is interpreted as a cointegrating relation if the hypothesis of a unit root in the residuals is rejected, where tests for a unit root are typically based on the augmented Dickey-Fuller (1981) statistic. Engle and Granger show that if the variables are cointegrated then there exists an error correction representation. Conversely, if there is an error correction representation for the series, then they are cointegrated. While intuitive and easy to implement, the Engle-Granger procedure often has little power to detect cointegration, and the long-run coefficient estimates from the static regression can be badly biased in finite samples (see Banerjee et al., 1986, and Kremers et al., 1992).

Johansen (1988, 1995) developed an asymptotically fully efficient, maximum likelihood systems estimation procedure for determining the number of cointegrating vectors and for estimating and conducting inference about the cointegrating vectors. The procedure is based on the so-called reduced rank regression method and presents some advantages over the Engle and Granger two-step approach (see Johansen and Juselius (1990) for money demand applications). First, it relaxes the assumption that the cointegrating vector is unique, and second, it takes into account that the cointegrating vector is unique, and second, it takes into

\[ A(L)(1 - L)x_t = -\gamma e_{t-1} + \varepsilon_t \]

where \( A(L) \) is a polynomial in \( L \) of the form \( (\alpha_0 + \alpha_1 L + \alpha_2 L^2 + \ldots) \), \( \varepsilon_t \) is a stationary multivariate disturbance. It is assumed that \( A(0) = I \), \( A(1) \) has all elements finite and \( \gamma \neq 0 \). The cointegrating vector is \( \alpha \), where \( \varepsilon_t = \alpha' x_t = 0 \), thus \( \varepsilon_t \) is interpreted as a measure of the error or deviation from equilibrium. In other words, if the variables are cointegrated, the residuals can be treated as the equilibrium correction term in subsequent specification of a dynamic model for the variables involved.
account the short-run dynamics of the system when estimating the cointegrating vectors.

The test of the number of cointegrating vectors ($r$) can be conducted using either of two statistics: the trace statistic, which tests the null hypothesis that the number of cointegrating vectors is less than or equal to $r$ against a general alternative, or the maximum eigenvalue statistic, which tests a null of $r$ cointegrating vectors against a specific alternative of $r + 1$. Asymptotic critical values for testing the number of cointegrating vectors appear in Osterwald-Lenum (1992) *inter alia*.12

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12 Dolado *et al.* (2001, p. 643) gives the following example to illustrate the underlying intuition behind Johansen's testing procedure. Assume $y_t$ represents a vector of $n$ I(1) variables and has a vector autoregressive representation of order 1 (VAR(1)) such that $A(L)y_t = \epsilon_t$ with $A(L) = I_n - A_1 L$ (note that Johansen deals with the more general case where $y_t$ follows a VAR($p$) process). This process can be reparameterised in the vector error correction model (VECM) representation as

$$\Delta y_t = Dy_{t-1} + \epsilon_t$$

where $D = (A1 - I_n) = -A(1) = -B\Gamma$, and $\epsilon_t$ is white noise. To estimate $B$ and $\Gamma$, $D$ is estimated using maximum likelihood estimation, subject to some identification restrictions since otherwise $B$ and $\Gamma$ could not be separately identified. If $\text{rank}(D) = 0$, then $y_t$ is I(1) and there are no cointegrating relationships ($r = 0$), whereas if $\text{rank}(D) = n$, there are $n$ cointegrating relationships among the $n$ series and hence $y_t$ is I(0). Thus testing the null hypothesis that there are $r$ cointegrating vectors is equivalent to testing whether $\text{rank}(D) = r$. 

2.3 Financial liberalisation, stationarity and structural stability of MDFs

Model stability is necessary for prediction and econometric inference. Since a parametric econometric model is completely described by its parameters, model stability is equivalent to parameter stability (Hansen, 1992). While model instability makes it difficult to interpret regression results, it is of particular importance in policy analysis to know if econometric models are invariant to possible policy interventions. A necessary condition for making a conditional money demand model immune from the Lucas (1976) critique is within sample parameter constancy (Ericsson, 1998).

Model instability may be caused by regime shifts/structural breaks or by the omission of an important variable. Both regime shifts and a missing variable in the money demand function have been linked to financial liberalisation and widely blamed for the instability of empirical money demand models (Hendry, 1979).

A broad definition of financial liberalisation involves institutional and policy changes as well as technological progress in transactions, which one usually interprets as financial innovation. Abiad and Mody (2005) identify three sources of financial liberalisation. First, reforms may be triggered by discrete events, or shocks, that change the balance of decision making such as various types of crisis, and external influences such as the leverage exercised by international financial institutions. Second, learning may foster reforms by revealing information that

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13 Regime shifts/structural breaks are structural changes in a time series sample. Although the exact definition of structural changes has not been given in the literature, it is usually interpreted as changes of regression parameters (Maddala and Kim, 1998).
causes reassessment of the costs and benefits of the policy regime. Third, reforms may be conditioned by political ideology of the ruling government and structural features such as openness to trade, legal system and form of government.

These sources of changes have different implications for the timing of liberalisation. Shocks leads to immediate policy change while learning allows for sustained changes. Existing measures of liberalisation can refer to a one-time change in rules (i.e. episodes of liberalisation), or to continuous proxies such as the level of financial development in an economy. Liberalisation is therefore "a mix of the episodic and the ongoing" (Abiad and Mody, 2005). While the timing, direction and size of policy changes may vary, it is important to observe that financial liberalisation is most likely to have permanent effects on the demand for money. Empirically, this phenomenon can be identified with all permanent shifts in money demand unrelated to the behaviour of the explanatory variables that appear most commonly in the literature.

2.3.1 Financial liberalisation and the random walk hypothesis

From the univariate point of view, structural breaks in nonstationary series can be viewed as permanent large shifts occurring intermittently, as against permanent small shifts occurring frequently and generating I(1) effects. The two forms of nonstationarity due to unit roots and regime shifts are closely related (Rappoport and Reichelin, 1989), and can be hard to discriminate empirically (Perron, 1989; Hendry and Neale, 1991). Stochastic process with short memory or stationarity that exhibit structural changes can display similar characteristics as the ones ob-
served in long memory or long range dependent processes. When one observes stationary series with structural breaks, the few occasional events or shocks with long-duration effects would produce certain persistence or symptoms of nonstationarity (Jimeno et al., 2006).

The distinction between random walk or broken trend stationarity of the series in the money demand function is important. Perron (1989) argues that, if a series is stationary around a deterministic trend that has undergone a permanent shift sometimes during the period of consideration, failure to take account of this change in the slope will be mistaken by the standard unit root test as a persistent innovation to a stochastic (nonstationary) trend. Hence a unit root test that does not take account of the break in the series will have very low power.\footnote{There is a similar loss of power if there has been a change in the intercept, possibly in conjunction with a shift in the slope of the deterministic trend.} If one assumes that the breakdate of the trend function is exogenous and chosen independently of the data, the unit root test can be adjusted by including composite dummy variables to ensure there are as many deterministic regressors as there are deterministic components in the data generating process.\footnote{These dummy variables take a value of \((0, 1)\) to allow for shifts in the intercept or are multiplied by a time trend to take into account any change in the slope of the deterministic trend.} Tests that are valid in the presence of such a break at a known point in time have been developed by Perron (1989; 1990) and Perron and Vogelsang (1993a; 1992b).

However, it is unlikely that the exact breakdate will be known \textit{a priori}. Furthermore, as argued by Christiano (1992) \textit{inter alia}, tests that treat the breakdate as exogenous are not appropriate in circumstances where the date of the break is
selected by reference to the data, for instance by looking at the plots of the series. Thus, one should use tests that endogenise the breakdate. Such tests have been proposed by Banerjee et al. (1992), Zivot and Andrews (1992), Perron (1997), Perron and Vogelsang (1992a) and Vogelsang and Perron (1998) inter alia.

From the multivariate point of view, conventional cointegration tests could yield misleading results in the presence of structural breaks (Leybourne and Newbold, 2003). Gregory et al. (1996) study the sensitivity of the augmented Dickey-Fuller (ADF) test for cointegration in the presence of a single break. Their Monte Carlo results show that the rejection frequency of the ADF test decreases substantially. Thus, in the presence of a break, one tends to underreject the null of no cointegration. The underrejection is similar to the underrejection of the unit root null hypothesis in the univariate case. However, in this case, the underrejection of the null indicates correctly that the constant parameter cointegrated relationship is not appropriate.

In a similar development to the extension of unit root test by deterministic components, cointegration tests have also been augmented by dummy variables, and made robust again structural breaks. Gregory and Hansen (1996) propose to test the null of no cointegration against the alternative of cointegration in the presence of a regime shift. The shift considered is a single break of unknown timing in the intercept and/or slope coefficients in the cointegrated relationship.
2.3.2 Financial liberalisation and the missing variable hypothesis

It is difficult to distinguish between genuine structural breaks that even a correctly
specified money demand function using appropriate modelling methodology would
break down under, and an omitted variable, which would also induce parameter
instability when a misspecified regression based on an inappropriate modelling
strategy is used.

To illustrate this point consider the theoretical money demand function of Ar­
rau et al. (1995) (equation (2.8)), which has the following econometric specification
(in log-linear form)

\[ m_t = \eta_t + \beta' x_t + v_t \] (2.16)

where \( m_t \) is money, \( x_t = [y_t, R_t]' \), \( y_t \) is the scale variable, \( R_t \) is the vector of
variables measuring the opportunity cost of money, \( \eta_t = \log(\theta) \), \( \theta \) is financial
liberalisation, and \( v_t \) is a stationary error term.

If the parameter capturing financial liberalisation in equation (2.16), \( \eta_t \), was
a stationary variable, the money demand function could be estimated without
controlling for financial liberalisation, since the process of liberalisation would be
part of a stationary error term. However, in so far as financial liberalisation has
permanent effects on the demand for money, equation (2.16) with an intercept
instead of \( \eta_t \) would be misspecified and cointegration of the money demand function
would not be obtained. Absence of cointegration would indicate that while the scale
and opportunity cost variables may be necessary for "pinning down" the steady-
state demand for money, these are not sufficient (Arrau et al., 1995). Intuitively, when compared to the previous hypothesis, this view of financial liberalisation suggests that the latter, while also having large permanent effects on the demand for money, is best characterised as an ongoing process of structural change rather than one inducing episodic regime shifts.
CHAPTER 3
AN OVERVIEW OF FINANCIAL LIBERALISATION IN INDONESIA

In the early 1980s, Indonesia was a high growth, low-income country heavily dependent on oil, with a financial system that was typical of most developing countries (Booth, 1998). The fall in oil revenues in 1982 prompted the Indonesian government to reform the financial system to make it more effective at mobilising and allocating savings to maintain investment. Financial reforms in Indonesia were implemented through major policy packages, most notably in 1983, 1988, 1990 and 1991. Financial liberalisation included the lifting of interest rate controls, the removal of directed credit programmes, the deregulation of banking activities and the opening of the capital market to foreign investors. The process, however, suffered a few set backs, most notably in 1997 when the country was hit by a severe economic crisis.

Although unique, the Indonesian experience is far from isolated. In fact, in the last quarter of the twentieth century, financial sector liberalisation was high on the agenda of policymakers and the trend worldwide was towards more liberalised financial systems. In two seminal papers, McKinnon (1973) and Shaw (1973) argue that financial repression is an obstacle to economic growth. Financial sectors, once they have been liberalised, can provide financial resources as cheaply and efficiently as possible which in turn stimulates the growth of the real economy. The experiences of financial liberalisation in Indonesia and elsewhere, however,
have also raised concerns about the possible link between financial liberalisation and economic crises.¹

The purpose of this chapter is to provide an overview of the financial reforms carried out in Indonesia since 1983 and discuss significant changes in monetary policy instruments and direction implemented in parallel with these reforms.

3.1 Banking system reforms

Throughout the 1970s, the Indonesian banking system exhibited many of the characteristics of financial repression described by McKinnon and Shaw. As described by McLeod (1999a, pp. 259-60)

the banking sector was dominated by five State commercial banks...

Regulations effectively prohibited the establishment of new private sector banks (domestic and foreign-owned), and heavily constrained the expansion of existing branch networks... State-owned banks accounted for roughly 80 per cent of total commercial bank assets throughout this period.

State-owned enterprises (SOEs) were required to hold their deposit accounts at the State banks and were also the main borrowers of State banks' loans. Both State and private banks did not compete for deposits because their lending was limited by credit ceilings. Under the directed subsidised credit schemes, known as liquidity credit schemes, State banks provided loans to targeted borrowers at

¹ The reader is referred to Levine (1997) for a review of the literature on financial liberalisation and economic growth and Demirguc-Kunt and Detragiache (2001) for a discussion of the causality from financial liberalisation to economic crises.
relatively low interest rates, with the loans eligible for refinancing from the central bank at subsidised rates.\textsuperscript{2} State banks had, therefore, little incentive to improve the level of service to borrowers and depositors alike, to offer attractive interest rates, or to improve their loan appraisal skills and their capacity to minimise loan losses (McLeod, 1999a).

By the early 1980s, the credit policies had led to a repressed financial system, with subsidised credit largely benefiting the State entreprises, which could access low-cost liquidity through their links to the State banks. In 1982, the central bank dominated the financial system, with about 45 percent of gross banking assets (22 percent of GDP). Deposit mobilisation was limited by the credit ceilings, limits on State banks lending and deposit interest rates, and branching. From 1975 to 1981, financial depth, as measured by the ratio of broad money, or M2, to nominal GDP, stagnated at about 16 percent, with an usually high fraction (60 percent) of M1 (currency and demand deposits) in broad money. The one-year deposit rate at the State banks, in real terms, averaged $-2.5$ percent from 1975 to 1982 and $-4.8$ percent from 1979 to 1982 (Hanna, 1994).

The fall in oil revenues in 1982, which led to a deterioration in Indonesia's current account deficit and a fall in GDP growth, prompted the Indonesian government to reform the financial system.

\textsuperscript{2} According to Hanna (1994, p. 5), direct credit allocation accounted for 48 percent of all bank lending by 1982. State banks granted loans at rates of 6 to 12 percent while the rediscount rates varied from one-third to one half of the subsidised rate paid by the borrower. The portion of a loan eligible for refinancing typically varied from 20 to 100 percent.
3.1.1 The June 1983 reforms

In June 1983 the government introduced a reform package known as PAKJUN, a major aspect of which was the curtailment of the subsidised loan programmes. The evidence had shown that the experience with the directed credit schemes did not have much impact on the course of structural change of the economy (Patten and Rosengard, 1991). Lending rates were freed and controls on all time deposit interest rates were removed to force the State banks to continue to lend based on deposit inflows alone. By the end of 1983, most subsidised loan programmes had been discontinued.

State banks responded to these policies by moving their time deposit and lending rates up to levels slightly below the prevailing rates at the private banks. Bank deposits increased very rapidly, with the growth of time deposits, those of the State banks in particular, far exceeding that of demand deposits. In contrast, the growth of bank lending was much more rapid in the private banks than in the State banks. As a result most State banks found themselves with an excess of funds which they were unable to lend, and became "chronic lenders" to the interbank market, while many private banks became "chronic borrowers" (Binhadi and Meek, 1992).

Although the reforms increased the flexibility of existing banks in pricing and allocating credit, they did not lower entry barriers, either among banks or between banks and other financial institutions. No new foreign banks were allowed entry.

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3 Private banks had not been subject to the interest rate controls, but had been under the credit ceilings.
and the opening of new branches continued to be severely constrained, giving State banks a clear advantage (Hanna, 1994).

3.1.2 The October 1988 reforms and further refinements

The government introduced in October 1988 a further deregulation package known as PAKTO (Government of Indonesia, 1988). The reforms contained in PAKTO removed impediments to competition by opening up possibilities for establishing new banks and new branch offices of existing banks (Binhadi, 1995). New domestic banks could be established, subject only to a small minimum capital requirement. Foreign banks were allowed to set up joint ventures with existing domestic banking partners. In addition, State enterprises were permitted to place up to half of their deposit funds with private banks. Another aspect of PAKTO was the reduction of the required reserve ratios from a multiple set of nominal rates that averaged about 11 percent for all banks together to a uniform level of 2 percent of all third party liabilities, thereby reducing the intermediation costs for banks and increasing the potential monetary expansion from any given increase in reserve money (Cole and Slade, 1996). Further refinements of the new rules came into supplementary packages introduced in March 1989 (Government of Indonesia, 1989) and January 1990 (Government of Indonesia, 1990), with the last significant (pre-crisis) deregulation reform contained in the banking law introduced in February 1992.

The March 1989 reform package known as PAKMAR removed controls on

---

4 Domestic banks were permitted to open new branches wherever they wanted, without limit, although foreign and foreign joint-venture banks were still only permitted no more than six offices each in only six major cities (McLeod, 1999a).
banks' offshore borrowing and introduced a prudential limit on banks' net open positions (NOP). The package announced on January 1990 known as PAKJAN abolished the remaining subsidised loan programmes, and moved interest rates on refinance from the central bank closer to market levels (Chant and Pangestu, 1994). The reform contained in the banking law enacted in February 1992 allowed foreign entities to acquire shares in domestic banks listed on the stock exchanges, and permitted the State banks to be listed (McLeod, 1992).6

Figure 3.1: Interest and inflation rates, 1980-2000

5 This term refers to the gap between banks' liabilities and assets denominated in foreign currencies, relative to their capital. Under the NOP constraint, banks may bring in as much foreign funds as they like provided they on-lend them predominantly in the same currencies.

6 Foreign shareholding, however, was not permitted to exceed 49 percent of the total, and the government was required to majority shareholding in any state bank whose shares were offered to the public.
3.1.3 The February 1991 and May 1993 reforms

In 1990, the private sector Bank Duta verged on bankruptcy as a result of its involvement in large-scale currency speculation (for a discussion, see Soesastro and Drysdale, 1990). Bank Duta had acquired a very large NOP, and subsequent exchange rate changes caused it to lose about US$420 million (McLeod, 1999a). This episode prompted the government to introduce in February 1991 a package of new prudential regulations known as PAKFEB (Government of Indonesia, 1991). The new prudential measures imposed various ratios for evaluating bank soundness and restricted bank activity in connection with the capital market (Hendrobudiyanto, 1994; Habir, 1994). The most important measures included new risk-adjusted capital adequacy requirements (CARs) (in line with the Basel Agreement), new loan/deposit ratio standards, and new restrictions on foreign exchange NOP for authorized foreign exchange banks. Further measures were also taken in 1992 including new laws on banking, insurance, and pensions.

In May 1993, a reform package known as PAKMEI relaxed and adjusted some of the PAKFEB bank prudential restraints in order to encourage new lending. Among other things, the CAR was marginally eased, the risk weighting of unused credit lines and loans to government enterprises was reduced from 100 percent to 50 percent, and capital as well as deposits was counted in calculating the loan/deposit ratio.

A number of previous deregulation measures were also reversed. BI reimposed controls on foreign borrowings by banks in 1991 (Muir, 1991).\footnote{Bank Indonesia Directors Decree No. 24/52/KEP/DIR, dated 29 November 1991.} In the mid-1990s,
BI adopted a target for the rate of growth of bank lending. The central bank also closed entry to almost all of the non-bank financial sector, which comprises leasing, factoring and consumer finance companies (with the exception of insurance and pension funds) in December 1995. Regulations on the issuing of commercial papers were also imposed in 1995. Under these regulations, any commercial paper issued or traded by banks had to be rated "investment grade" by a domestic agency established by the government (PT Pefindo) and was limited to a maximum 270 days' maturity.

3.1.4 Impact of the reforms on the banking system

The liberalisation of interest rates, together with the substantial progress achieved in reducing inflation, resulted in positive real interest rates during the 1980s and most of the 1990s (Figure 3.1). By contrast, prior to 1983, real interest rates had tended to be negative. The immediate impact of liberalisation was to move rates upwards.

Positive real interest rates in the 1980s contributed to financial deepening (or monetisation). For the period 1983-1988, the ratio of M2 to GDP increased from 17 to 30 percent in 1988 and further to about 60 percent at the end of 1997 (Figure 3.2). In contrast, the ratio of M1 to GDP fluctuated within a range of 8-12 percent over the same period.

During this period total assets of commercial banks increased 10-fold to Rp715.2

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8 Bank Indonesia Directors' Decree No. 28/52/KEP/DIR, dated 11 August 1995.
9 The real interest rate can be measured by the nominal interest rate minus the change in the consumer price index in the last 12 months. This assumes that inflation expectations can be approximated by actual inflation over the past 12 months.
trillion, exceeding for the first time the size of the country's GDP; nonbank financial institutions (NBFIs) also flourished with total assets growing at an annual compound rate of 56 percent in 1991-1996, reaching Rp80 trillion by end-1996 (Chou, 1999, p. 36). The banking system, nevertheless, still held about 65 percent of the financial sector's total assets in 1997.

![Figure 3.2: Monetary ratios as percentage of GDP, 1983-2000](image)

Changes in the ratios of time and savings deposits to GDP and demand deposits to GDP (Figure 3.3) also offer a good perspective on the growth of the banking system. Time and savings deposits, having been stagnant relative to output prior to 1983, suddenly began to increase rapidly. The impact of liberalisation on demand deposits, however, is less clear cut.

The 1988 reforms also stimulated the physical expansion of the banking industry. Some ninety-four new private domestic banks were established in the period...
Figure 3.3: Impact of deregulation on bank deposits (outstanding balances as % of nominal GDP), 1980-2000

up to June 1997, plus thirty-two joint foreign ventures; some 3,674 new branches were added to the private banks' network, while only 611 new State bank branches appeared; the total number of branches increased from 1,728 to 6,337 during this period (McLeod, 1999a, p. 275).

By opening up the financial sector to new entrants, most of which were small and joint venture banks, the reforms intensified the competition for funds. In 1983-1988, the seven State banks' share of total outstanding credit averaged around 65 percent. In contrast, after three years of liberalisation their share dropped to 56 percent in 1991 and further to 40 percent by end-1997. Private banks experienced an annual growth rate in their assets and credits of 60 and 39.5 percent in the four years after PAKTO, compared with 22 and 21 percent for the State banks (Chou, 1999, p. 36).
3.2 Capital market reforms

The capital market includes both equity and bond markets. The equity market in Indonesia is made of three Stock Exchanges: the Jakarta Stock Exchange (JSX), the Surabaya Stock Exchange (SSX), and an over-the-counter (OTC) market known as Bursa Paralel (created in 1988 and merged with the SSX in 1994).\(^1\)

While the bond market expanded rapidly between 1992 and 1997, it remains small and bank finance is still the dominant form of nonequity capital (Wells, 1999).\(^1\)

Originally opened in 1912 under the Dutch colonial government, the JSX reopened on 10 August 1977 with only one company listed. A year before, the Capital Market Executive Agency (Badan Pelaksana Pasar Modal, or BAPEPAM), an agency which answers to the Ministry of Finance, had been established as capital market regulator. In 1978, the Ministry of Finance issued decrees authorising the establishment of trustees and guarantors in the capital market.\(^1\) The following year, the government opened up the right for all Indonesian financial institutions, including pension funds, to buy securities on the Stock Exchange.\(^1\) In 1981, a further decree was announced allowing continuous trading at the Stock Exchange.\(^1\)

\(^1\) The SSX (and Bursa Paralel) established in 1989 is facilitating three markets: stock trading for small and medium scale companies, bond trading/reporting for corporate and government bonds, and derivatives trading for stock index futures.

\(^1\) The bond market is, therefore, omitted from this discussion. The interested reader is referred to Wells (1999).


\(^1\) Ministry of Finance Decree No. 313/KMK.011/1979.

\(^1\) Ministry of Finance Decree No. 161/KMK.011/1981.
3.2.1 The December 1987 and December 1988 reforms

On 23 December 1987, the Ministry of Finance issued a package of deregulation measures aimed at the capital market known as PAKDES I. The measures softened requirements for licensing underwriters, brokers, dealers, trustees, and guarantors; outlined requirements and procedures for public offerings of both shares and bonds, requirements for the prospectus, financial statements, and reporting and public recording of these documents. PAKDES I also authorised foreign portfolio investment in the JSX. Foreign investors were allowed to purchase shares up to an amount allowed by the provisions of the Investment Coordination Board (Badan Koordinasi Penanaman Modal, or BKPM) on foreign capital participation. Finally, the package also eliminated the limitations on daily price movements on the JSX, thereby freeing it from direct price controls (Cole and Slade, 1996).

A second package, known as PAKDES II, was issued in December 1988. The package focused on stimulating the capital market as well as non-bank financial institutions (NBFIs) and the insurance sector.¹⁵ The capital market measures included provisions authorising the establishment of a private securities exchange in Indonesia, including prohibitions against insider trading. New regulations also opened the market to foreign investors by allowing up to 85 percent foreign share ownership in securities companies. Another important measure was to subject domestic deposits to a 15 percent withholding tax, the same tax levied on dividend

¹⁵ NBFIs are categorised into two: those engaged primarily in capital market activities and the rest. The former are regulated and supervised by BAPEPAM, while the rest-finance companies, venture capital companies, insurance companies, and pension funds-are supervised by the Ministry of Finance, with assistance from BI in some cases (Chou, 1999).
payments. The aim was to move towards a more level playing field for financial instruments. PAKDES II also contained new regulations covering the establishment of multi-finance companies allowed to engage in leasing, factoring, venture capital, credit card operations and consumer credit. The same activities were made available to banks (Hanna, 1994).

3.2.2 The December 1990 reforms and 1992 Amendments to the Income Tax Law
A third package, known as PAKDES III, was adopted in December 1990. The package provided for the privatisation of the JSX (under the ownership of PT Jakarta Stock Exchange), which took place on 13 July 1992. The package also provided for the conversion of the Capital Market Executive Agency into the Capital Market Supervisory Agency (under the same acronym, BAPEPAM). New requirements for securities companies (including brokers and underwriters) were also adopted: the company listing requirement, the requirement for cash dividends, and the delisting criteria. The first ruling meant that all shares had to be traded on an exchange when they were sold, and that a larger percentage had to be available to foreigners; the second ruling was that a company could be delisted if it did not pay a cash dividend for three consecutive years; the third ruling was that a company could be delisted on the basis of one or more of the following criteria: less than 100 shareholders, no trades for six consecutive months, no cash dividend and/or losses of 50 percent for three consecutive years, and equity less than Rp3 billion (Cole and Slade, 1996).

The 1992 Amendments to the Income Tax Law set out the principles to re-
solve outstanding tax issues affecting the development of the capital market. The existing Income Tax Law of 1984 contained three discriminatory taxes: a tax discrimination between interest income from securities and from bank deposits, the triple taxation of investment funds and the discriminatory taxation of dividends of subsidiaries of banks and other financial institutions. The amendments introduced in 1992 imposed full taxation on the interest on corporate time deposits, allowed tax exemption on the income from investment funds distributed to unitholders, and permitted dividends from subsidiaries of financial institutions to receive the same treatment as all other dividends.

3.2.3 Corporate and capital market laws

The Corporate Law of 1995 governs the establishment and operation of limited liability companies in Indonesia. The law provides a number of rights to or protection of shareholders including access to regular and reliable information free of charge, one share one vote, mandatory shareholders' approval of interested transactions, mandatory shareholders' approval of major transactions, mandatory disclosure of connected interests, independence of auditing, mandatory independent board committee, and severe penalties for insider trading.

The Capital Market Law of 1995 regulates companies listed in the stock exchange, delineating the tasks and responsibilities of the Capital Market Supervisory Agency. It regulates the requirements of investment companies, securities companies, underwriters, brokers, investment managers, investment advisors, and other supporting agencies. It also regulates reporting and auditing procedures, trans-
parenity requirements, insider trading investigation, and administrative and legal punishment.

3.2.4 Impact of the reforms on the capital market

The capital market has grown significantly following the three deregulation packages introduced in the late 1980s. The number of equity share-issuers at the JSX, 132 in 1990 rose to reach 379 at the end of 2001 and the value of equity issues outstanding went from just over Rp8 trillion to almost Rp71 trillion during this period (Table 3.1).

The role of foreign capital has been crucial in driving the market in the early part of the 1990s. Foreign investors have dominated the market up until the end of 1997 according to the percentage of shares they owned at the end of each year (Table 3.2).

The increasing importance of the capital market as a source of funds for businesses and as an investment alternative for the public is also illustrated by the growth in the number and value of average yearly traded shares on the JSX as well as total market capitalisation figures (Table 3.3). For the period 1992-1997 the total value of stock market transactions rose from Rp2 trillion to Rp120 trillion and the volume of transactions increased every year from 1990 (except in 2000). Total share market capitalisation grew from Rp112.1 billion in 1987 to Rp482 at the end of 1988, and then to just over Rp215 trillion at the end of 1996.

The composite share price index rose from 82 in December 1987 to 305 in December 1988 and stayed in the 400s in both 1989 and 1990 (Table 3.4). The
<table>
<thead>
<tr>
<th>Year</th>
<th>Number of issuers</th>
<th>Number of outstanding issues</th>
<th>Value (Rp billion)</th>
</tr>
</thead>
<tbody>
<tr>
<td>1990</td>
<td>132</td>
<td>965,394,556</td>
<td>8,009.4</td>
</tr>
<tr>
<td>1991</td>
<td>145</td>
<td>1,178,465,725</td>
<td>8,976.1</td>
</tr>
<tr>
<td>1992</td>
<td>162</td>
<td>1,761,393,686</td>
<td>11,161.8</td>
</tr>
<tr>
<td>1993</td>
<td>181</td>
<td>3,338,513,735</td>
<td>16,065.0</td>
</tr>
<tr>
<td>1994</td>
<td>231</td>
<td>6,401,933,047</td>
<td>26,528.0</td>
</tr>
<tr>
<td>1995</td>
<td>248</td>
<td>11,110,964,641</td>
<td>35,395.0</td>
</tr>
<tr>
<td>1996</td>
<td>267</td>
<td>25,343,423,026</td>
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</tr>
<tr>
<td>1997</td>
<td>306</td>
<td>51,459,413,729</td>
<td>70,879.6</td>
</tr>
<tr>
<td>1998</td>
<td>309</td>
<td>62,719,348,831</td>
<td>75,947.0</td>
</tr>
<tr>
<td>1999</td>
<td>321</td>
<td>714,460,834,556</td>
<td>206,686.8</td>
</tr>
<tr>
<td>2000</td>
<td>347</td>
<td>811,675,983,545</td>
<td>226,057.3</td>
</tr>
<tr>
<td>2001</td>
<td>379</td>
<td>826,770,663,455</td>
<td>231,342.1</td>
</tr>
</tbody>
</table>

Table 3.1: Share issues by listed companies, 1990-2001

<table>
<thead>
<tr>
<th>Year</th>
<th>Domestic investors</th>
<th></th>
<th>Foreign investors</th>
<th></th>
<th>Total</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Value (Rp billion)</td>
<td>%</td>
<td>Value (Rp billion)</td>
<td>%</td>
<td>Value (Rp billion)</td>
</tr>
<tr>
<td>1992</td>
<td>857.5</td>
<td>41.02</td>
<td>1,237.8</td>
<td>59.08</td>
<td>2,095.2</td>
</tr>
<tr>
<td>1993</td>
<td>7,538.6</td>
<td>39.50</td>
<td>11,547.6</td>
<td>60.50</td>
<td>19,086.2</td>
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<td>1994</td>
<td>7,601.7</td>
<td>29.83</td>
<td>17,881.1</td>
<td>70.17</td>
<td>25,482.8</td>
</tr>
<tr>
<td>1995</td>
<td>10,667.0</td>
<td>32.97</td>
<td>21,690.5</td>
<td>67.03</td>
<td>32,357.5</td>
</tr>
<tr>
<td>1996</td>
<td>30,036.3</td>
<td>39.66</td>
<td>45,693.6</td>
<td>60.34</td>
<td>75,729.9</td>
</tr>
<tr>
<td>1997</td>
<td>57,584.0</td>
<td>47.82</td>
<td>62,801.1</td>
<td>52.18</td>
<td>120,385.2</td>
</tr>
<tr>
<td>1998</td>
<td>58,128.1</td>
<td>58.31</td>
<td>42,556.6</td>
<td>41.69</td>
<td>99,684.7</td>
</tr>
<tr>
<td>1999</td>
<td>96,152.6</td>
<td>65.02</td>
<td>51,727.4</td>
<td>34.98</td>
<td>147,880.0</td>
</tr>
<tr>
<td>2000</td>
<td>98,090.7</td>
<td>79.89</td>
<td>24,684.0</td>
<td>20.11</td>
<td>122,774.8</td>
</tr>
<tr>
<td>2001</td>
<td>87,005.8</td>
<td>89.22</td>
<td>10,517.0</td>
<td>10.78</td>
<td>97,522.8</td>
</tr>
</tbody>
</table>

Table 3.2: Trading value of listed shares on the JSX by type of investor, 1992-2001

<table>
<thead>
<tr>
<th>Year</th>
<th>Traded shares</th>
<th>Market capitalisation</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Number (million)</td>
<td>Value (Rp billion)</td>
</tr>
<tr>
<td>1983</td>
<td>3.5</td>
<td>10.1</td>
</tr>
<tr>
<td>1984</td>
<td>1.2</td>
<td>2.1</td>
</tr>
<tr>
<td>1985</td>
<td>1.9</td>
<td>3.0</td>
</tr>
<tr>
<td>1986</td>
<td>1.4</td>
<td>1.8</td>
</tr>
<tr>
<td>1987</td>
<td>2.5</td>
<td>5.2</td>
</tr>
<tr>
<td>1988</td>
<td>6.9</td>
<td>30.6</td>
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<tr>
<td>1989</td>
<td>95.8</td>
<td>964.3</td>
</tr>
<tr>
<td>1990</td>
<td>702.6</td>
<td>7,311.3</td>
</tr>
<tr>
<td>1991</td>
<td>1,007.9</td>
<td>5,778.2</td>
</tr>
<tr>
<td>1992</td>
<td>1,706.3</td>
<td>7,953.3</td>
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<td>1993</td>
<td>3,844.0</td>
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<td>1994</td>
<td>5,292.5</td>
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<td>1995</td>
<td>10,646.4</td>
<td>32,357.5</td>
</tr>
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<td>1996</td>
<td>29,527.7</td>
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<td>1997</td>
<td>76,599.1</td>
<td>120,385.2</td>
</tr>
<tr>
<td>1998</td>
<td>90,620.5</td>
<td>99,684.7</td>
</tr>
<tr>
<td>1999</td>
<td>178,486.6</td>
<td>147,880.0</td>
</tr>
<tr>
<td>2000</td>
<td>134,531.3</td>
<td>122,774.8</td>
</tr>
<tr>
<td>2001</td>
<td>148,381.3</td>
<td>97,522.8</td>
</tr>
</tbody>
</table>

Table 3.3: JSX, yearly statistics, 1983-2001

index fell back to 247 at the end of 1991 before taking off again in 1992 and reaching a record 637 at the end of 1996.

3.3 Changes in monetary policy

Monetary policy focused until 1983 on credit controls (i.e. credit ceilings and interest rate controls), which covered both private and State banks. These controls, however, proved ineffective in constraining growth of the money supply or inflation. This was epitomised by famous "incidents" (so-called "Pertamina Crisis" in 1975, reduction of the required reserve ratio from 30 percent to 15 percent in December 1977). Furthermore, the continuously positive balance of payments had an impact on reserve, or base money and was a major factor in the expansion of money supply (McLeod, 1993). When the oil boom ended in 1982, in parallel with the series of market-oriented financial reforms, the approach of monetary policy began to shift from direct to increasingly indirect instruments of intervention.

3.3.1 Policy objectives and intermediate targets

For the period 1983 to May 1999 (when the new central bank law was enacted), the basic objectives of monetary policy in Indonesia, as stipulated in Act No. 13 of 1968 concerning the central bank, were to assist the government in (a) managing, safeguarding and maintaining the stability of the Rupiah, and (b) facilitating the production and development with the aim of promoting employment creation and improving the living standards of the population. Accordingly, BI was responsible for formulating and implementing monetary policy which was directed towards
<table>
<thead>
<tr>
<th>Year</th>
<th>JSX</th>
</tr>
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<tbody>
<tr>
<td>1983</td>
<td>80.37</td>
</tr>
<tr>
<td>1984</td>
<td>63.53</td>
</tr>
<tr>
<td>1985</td>
<td>66.53</td>
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<tr>
<td>1986</td>
<td>69.69</td>
</tr>
<tr>
<td>1987</td>
<td>82.43</td>
</tr>
<tr>
<td>1988</td>
<td>305.12</td>
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<td>1989</td>
<td>399.69</td>
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<tr>
<td>1990</td>
<td>417.79</td>
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<td>1991</td>
<td>247.39</td>
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<td>1992</td>
<td>274.33</td>
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<td>1993</td>
<td>588.76</td>
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<td>1994</td>
<td>469.64</td>
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<tr>
<td>1995</td>
<td>513.84</td>
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<td>1996</td>
<td>637.43</td>
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<tr>
<td>1997</td>
<td>401.71</td>
</tr>
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<td>1998</td>
<td>398.03</td>
</tr>
<tr>
<td>1999</td>
<td>676.91</td>
</tr>
<tr>
<td>2000</td>
<td>416.32</td>
</tr>
<tr>
<td>2001</td>
<td>392.04</td>
</tr>
</tbody>
</table>

Table 3.4: JSX, yearly (end-period) composite share price index, 1983-2001

achieving several objectives, namely controlling inflation, avoiding serious loss of foreign exchange reserves, supporting investment in general, or in certain priority areas of economic activity, and achieving relatively high economic growth so as to increase per capita income (Iljas, 1998). BI sought to achieve these multiple objectives primarily through control of monetary aggregates (M1, M2) at levels that are adequate to support the targeted rate of economic growth. Under this framework, monetary targets are the intermediate target to be influenced in turn by controlling the amount of reserve money (M0—the operational target) through open market operations. Monetary policy thus relied on the linkages between M0 and M1/M2 and between M1/M2 and the ultimate objectives.\footnote{Monetary aggregates in Indonesia are classified as follows: (1) Coins and currency in circulation; (2) M0: base money (cash plus the reserves of the commercial banks); (3) M1 (cash plus the non-interest bearing demand deposits at the commercial banks); and (4) M2 (M1 plus the interest bearing time deposits at the commercial banks).}

Although BI never published its monetary targets, McLeod (1999b) reports that targets were set for growth of M1 and M2 of the order of 19-20 percent per year in 1995, and it was assumed that there were reliable multipliers which would enable these targets to be met through appropriate manipulation of base money. The multipliers were 1.9 (mm$_1$) and 7.8 (mm$_2$) at March 1995 (Bank Indonesia, 1995).\footnote{For comparison, the multipliers were about 1.4 (mm$_1$) and 6.5 (mm$_2$) at May 2004 (Bank Indonesia, 2004, p. 11, Graph. 22).} The required changes in base money were derived from

\[
\Delta M_n = mm_n \Delta B, \quad n = 1, 2
\]  

(3.1)

where $mm$ is the money multiplier which shows the changes in money (M1 or M2), $\Delta M_n$, for a given change in base money, $\Delta B$. 

Hence the framework required stability of the money demand function, income velocity and money multipliers.

3.3.2 Policy instruments

The PAKJUN policies allowed the State banks to raise their deposit and loan interest rates. The resulting imbalance between the deposit mobilising and lending capabilities of the State and private banks stimulated the growth of the interbank money market. In such context the monetary authorities undertook the development of an indirect instrument to influence the total supply of reserve money in the banking system. In the absence of any government debt instruments, BI introduced a new money market instrument in February 1984, called a Sertifikat Bank Indonesia (SBI), which is a short-term liability of the Bank used to reduce banking system liquidity (Binhadi and Meek, 1992). The SBIs have maturities of 30 and 90 days and are sold each week through an auction. In early 1985, BI introduced another money market instrument to increase banking system liquidity called the surat berharga pasar uang (SBPU), which is a money market security issued or endorsed by banks that BI is prepared to purchase at a discount from the banks. The rate of discount is set by BI.

The two new instruments formed the new basis for monetary policy to replace direct bank lending controls, "a practical approximation to open market operations in other countries where government debt is bought and sold by the central bank" (McLeod, 1999a). In effect, BI takes part in the interbank market by buying up any excess reserves from the banks through sales to them of SBIs and then supplying
reserves through discounting of their SBPUs. BI sets no limits on the amount of SBIs it sells to the State banks, but it sets ceilings on the amount of SBPUs it buys from each bank as a means of controlling the growth of reserve money. SBPUs are, therefore, used primarily as quantitative instruments of monetary policy, to control the supply of reserve money while the SBI is a fixed interest rate instrument used for liquidity adjustment, primarily by the State banks (Cole and Slade, 1996).

3.3.3 Reverting to direct controls: the Sumarlin shocks

The sharp decline in world oil prices in 1986-1987 and the combination of steady, heavy losses on foreign exchange reserves and high interbank interest rates throughout this period eventually led the government to intervene in June 1987. The authorities implemented a dramatic contraction of domestic liquidity, which came to be labeled "Sumarlin shock", after Sumarlin, the then Minister of Finance.

The liquidity contraction was accomplished by two specific measures. The first was an instruction to four large State-owned enterprises to transfer part of their time deposit balances at the State banks into holdings of new SBIs. The second was a reduction in the ceilings on SBPUs discounted at the central bank to zero, in effect forcing all the commercial banks to buy back their SBPUs from Bank Indonesia. These measures wiped out, almost overnight, the equivalent of all the available legal reserves of the banking system (Cole and Slade, 1996). As a result, banks could only meet their reserve requirements by selling foreign exchange to BI, or borrowing through the discount window. The use of the central bank facility, however, was discouraged by an increase in the discount rate from 20 percent to
30 percent.

On 27 February 1991, a number of large State-owned enterprises were again instructed to transfer their time deposits in banks into holdings of SBIs. The amount transferred on 1 March 1991 was Rp8 trillion, equal to approximately one-third of M1 or three-quarters of reserve money. BI offered to purchase SBPUs from the affected banks to offset 80 percent of their loss of funds. The net contractionary effect on reserve money from this sale of SBIs and purchase of SBPUs was about Rp2 trillion. This exceeded the total reserves of the banking system averaging about Rp1.6 trillion. As a result, both SBI and SBPU markets were reactivated, allowing the central bank to have larger stocks of these instruments outstanding with which to carry out money market operations. On the negative side, SBI rates as well as bank deposit and loan rates were held at high levels throughout 1991. Growth of total bank credit, which had averaged 51 percent per year in 1989-1990, dropped to 16 percent in 1991 and 9 percent in 1992; private national bank credit only increased by 1 percent in 1992, compared with average annual increases of more than 80 percent in 1989-1990 (Cole and Slade, 1996, p. 61).

3.4 Crisis-driven reforms and policy changes

The Indonesian economic crisis started as a contagion of the Thai currency crisis of July 1997. The complex chain of events that led to the crisis has been discussed in a number of papers (see for instance Enoch et al., 2003; Grenville, 2004). The following discussion provides a brief overview of monetary policy approach during the crisis as well as crisis-driven changes to banking regulations.
3.4.1 Floating of the Rupiah and third Sumarlin shock

Critical and far-reaching decisions were taken in the initial phase of the crisis, before the involvement of the IMF towards the end of October 1997. The exchange rate band was widened in July 1997 from 8 percent to 12 percent, to allow the rate to move without requiring BI to use reserves in its defence. On 14 August 1997, the Rupiah was floated. A few days later, a major contractionary open market operation took place, which preceded was the earlier Sumarlin shocks. SOEs were once again instructed to withdraw their bank deposits and use them to buy SBIs, thereby raising concerns that banks would not have enough liquidity to meet either their small reserve requirements or their daily cheque clearing obligations in the payments system.

On 31 October 1997, a letter of intent from the Indonesian government to the IMF set out a new IMF-endorsed monetary strategy. The main element of the
strategy was to set targets for the slow growth of base money (7.7 percent according to Fane, 2000, p. 50) and its components—net domestic credit and net foreign assets. The aim was to restore confidence in the Rupiah and cause a reversal of the depreciation. Accordingly, BI followed a high interest rate policy and raised SBI rates (the benchmark rates) several times (Figure 3.4). The rates increased from the average 14 percent pre-crisis to 30 percent in December 1997, and 80 percent in August 1998. As a result, the amount of interbank call money transactions surged by 400 percent, from Rp0.5 trillion in 1996 to Rp2 trillion in 1997. The sale of SBI also rose by four times in value, while SBPU rose by around 80 percent over the same period (Halim, 2000). SBI rates were subsequently reduced to 40 percent in March 1999 before returning gradually to their pre-crisis level.

Despite the base money targets and high interest rate policy, the exchange rate of the Rupiah against the US dollar depreciated by almost 30 percent in 1998 from its 1997 value (by more than 80 percent in the first months of 1998), and stabilised at a much depreciated value in the first half of 1999. Between September 1997 and September 1998, base money grew by 95 percent (Fane, 2000). On the supply side the closure of 16 private banks in November 1997 caused runs on individual banks and forced BI to provide bank liquidity support (Bantuan Likuiditas Bank Indonesia or BLBI) amounting to around Rp150 trillion.\(^{18}\) On the demand side there was an important increase in the demand for currency from the public, motivated by the loss of confidence in the banking system and rapidly

\(^{18}\) These liquidity injections were partly offset by net issue of SBIs and foreign currency intervention amounting to around Rp10 trillion (Halim, 2000).
rising prices. Currency in circulation doubled in the year to mid 1998, accounting for 85 percent of the increase in base money (Grenville, 2000).

3.4.2 Reforms to banking laws and regulations

The following discusses amendments to the 1992 banking law enacted in November 1997, the new central bank law enacted in May 1999, and various changes to the banking prudential regulations introduced during 1998, drawing from McLeod (1999b).19

Amendments to the Banking Law introduced in November 1997 lifted the previous requirement that the government retain majority ownership of State banks; abolished regulations discriminating between domestic and foreign joint venture bank (in particular, in relation to the expansion of their branch networks); and allowed foreign banks as well as foreign non-bank financial institutions and corporations to acquire up to 100 percent of the shares in existing banks (including the State banks). New clauses also provided the legal basis for the establishment and operations of the Indonesian Bank Restructuring Agency (IBRA) and for the establishment of a deposit insurance institution.

The central bank law enacted in May 1999 established the stability of the Rupiah, in terms of both its purchasing power and its rate of exchange for other currencies as the sole objective of BI. The law also imposed new constraints on BI including the requirement that its monetary liabilities do not exceed its capital by more than a factor of 10; the interdiction to acquire equity in other entities

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without the approval of parliament; the interdiction to lend to the government, or
to purchase government bonds except in the secondary market; and the limitation
to a maximum of 90 days for the provision of credit to the banks as lender of last
resort. The new law also called for the transfer of the task of prudential supervision
of banks from BI to a new, independent institution, by the end of 2002.

BI decrees issued in November 1998 softened the existing prudential regula­
tions. The decrees focused on capital adequacy (No. 31/146/KEP/DIR),\textsuperscript{20} the
classification of assets by quality (No. 131/47/KEP/DIR); the setting aside of
reserves against possible reductions in asset values (No. 31/148/KEP/DIR) as
well as the procedures to be followed when restructuring borrowers’ debts (No.
31/150/KEP/DIR).

3.4.3 Impact of the crisis on the financial sector

The crisis had a significant impact on the banking system, which as mentioned
above held as much as 65 percent of the financial sector’s total assets in 1997.
The income of the banking system declined by 300 percent in 1997 (Halim, 2000).
Banks became incapable of performing their financial intermediary functions. The
sharp devaluation of the Rupiah against the US dollar caused investors, who had
borrowed in US dollars (about 30 percent of all bank loans) to default, as the
value of their collateral was now very much below the value of their loans.\textsuperscript{21} High

\textsuperscript{20} The minimum required capital adequacy ratio was reduced from 8 percent before the crisis
to 4 percent. Moreover, banks were permitted to include part of their loan loss provisions in this
figure.

\textsuperscript{21} In July 1998, the Asset Management Unit (AMU) was established as part of IBRA to handle
non-performing loans from banks.
interest rates also encouraged defaults. November 1997 saw the liquidation of 16 banks. In February 1998, IBRA intervened in 54 weak banks, whose BI emergency borrowings exceeded 200 percent of their capital and whose capital was less than 5 percent of their assets. In April 1998 IBRA suspended 7 banks whose borrowings exceeded 500 percent of their capital, and more than 75 percent of their assets. In March 1999 38 banks were closed, 9 banks recapitalised and 7 banks were taken over by IBRA.

The crisis also had a significant impact on the capital market. There was a major decline in foreign and domestic portfolio investment (Table 3.2). The share of transactions by foreign investors fell from 52 percent in 1997 to 42 percent the next year and was just under 11 percent in 2001. The total value of stock market transactions fell from Rp120 trillion in 1997 to Rp99 trillion the next year. Market capitalisation fell by about 35 percent in 1997 compared to 1996 (Table 3.3). Based on the Price Earning Ratios, the average stock market price was halved, from 21.6 to 10.5, with many share prices below Rp500, about two cents at Rp8,500 per US dollar (Halim, 2000). As share prices fell, the composite share price index plunged from 637 points at the end of 1996 to only 402 points at the end of 1997 and just under 400 points at the end of 1998 (Table 3.4). Despite a recovery to around 678 points at the end of 1999, it plunged again to 416 points at the end of 2000 and reached 392 points at the end of 2001.

22 The Price Earning Ratio (PER) is the ratio of the regular closing price over the earnings per share (EPS). The latter is derived by dividing net profit (after taxation) by the number of shares issued.
CHAPTER 4
FINANCIAL LIBERALISATION AND STRUCTURAL BREAKS IN
MONETARY AGGREGATES AND THE DETERMINANTS OF MONEY
DEMAND IN INDONESIA

4.1 Introduction

Many experts claim that the process of financial liberalisation experienced by Indonesia since 1983 has caused instability in the money demand function somewhere in the 1990s (see for instance Alamsyah et al., 2001). However, there has been no systematic attempt at corroborating such claim. In addition, the issue of whether any potential instability in the money demand relationship is a temporary or a permanent phenomenon has been largely ignored.

In this chapter we want to investigate this claim. We start from the observation that there are not many results on the time series properties of monetary aggregates and the determinants of money demand in Indonesia. Previous studies including inter alia Tseng and Corker (1991); Price and Insukindro (1994); Dekle and Pradhan (1997; 1999); and Chaisrisawatsuk et al. (2004) model these series as unit root processes. This practice, however, suffers from two important caveats. First, neglecting the possible presence of structural breaks might bias the results of unit root tests towards a failure to reject the unit root null hypothesis. Perron (1989); Rappoport and Reichlin (1989); and Hendry and Neale (1991)
have shown that (i) structural change/regime shifts can mimic unit roots in stationary autoregressive time series; (ii) such shifts are very hard to detect using conventional parameter constancy tests. Second, applying the conventional integration/cointegration approach if the series are actually stationary around broken deterministic trends can give misleading results (Leybourne and Newbold, 2003).

Since the detection of structural break in economic variables crucially depends on the maintained hypothesis about the dynamics of the variables into question, we apply the univariate method proposed by Vogelsang and Perron (1998) designed to detect structural breaks on nine series in a standard specification for money demand using Indonesian data spanning from 1983:1-2001:4. This statistical method can detect single breaks of unknown timing in both the intercept and slope of the trend function of the series, assuming each of them follows a simple auto-regressive (AR) process.1

Finally, unlike previous studies of the demand for money in Indonesia, the sample period covers the four critical quarters—the last two quarters of 1997 and the first two quarters of 1998—often dubbed the "crisis period" in the Indonesian economic crisis literature (Grenville, 2000; Fane, 2000).

Structural change would be quite perilous to ignore. Inferences about the money demand relationship could go astray, forecasts could be inaccurate and policy recommendations could be misleading. We find evidence in favour of a

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1 While this approach is the most advanced method to endogenously detect structural breaks, it is unable to identify multiple structural breaks. As Ben-David and Papell (1998) note, tests that allow for multiple structural breaks, such as Bai and Perron (1998, 2003), are restricted to stationary and non-trending data, which is not the case for the variables under investigation in this study.
structural break in all the series except the three monetary aggregates (M0, M1, and M2). Moreover, we find that the break for most series occurs around the end of 1997 and coincides with the crisis and a major government intervention in the money market for which the precedent is the Sumarlin shock of 1991 (discussed in the previous chapter).

The rest of the chapter is organised as follows. Section 4.2 presents the Vogelsang-Perron methodology. Section 4.3 discusses the empirical results. Section 4.4 offers some concluding remarks.

4.2 Methodology

To detect structural breaks in our series, we apply the procedure outlined in Perron (1989), Zivot and Andrews (1992), and Vogelsang and Perron (1998) (VP hereafter). These authors have shown that the presence of structural breaks, if ignored, bias standard unit root tests towards the acceptance of the unit root hypothesis. VP have developed unit root tests which allow for the possibility of a structural break in the deterministic trend and in which the breakdates are endogenously determined. The null hypothesis is that of a unit root against the alternative of stationarity with a break in the intercept and slope of the trend function.

Throughout, $T_b$ denotes the time at which the change in the time function occurs. The first model allows only a shift in the intercept under both the null and alternative hypotheses (crash model). Under the second model, both a shift in intercept and slope are allowed at time $T_b$ (crash/changing growth model). The
third model allows a smooth shift in the slope by requiring the end points of the two segments of the broken trend to be joined (changing growth model).

The innovational outlier (IO) models the break as evolving slowly over time while the additive outlier (AO) models the break as occurring suddenly. The choice of AO versus IO depends on the view one is taking as to the dynamics of the transition path following a break.

Under the IO model, the change is assumed to occur gradually. The test of the null hypothesis of a unit root is performed using the $t$-statistics for testing $\alpha = 1$ in the regressions:

$$y_t = \mu + \beta t + dD(T_b)t + \theta DU_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + u_t, \quad (4.1)$$

$$y_t = \mu + \beta t + dD(T_b)t + \theta DU_t + \gamma DT_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + u_t. \quad (4.2)$$

$$y_t = \mu + \beta t + \gamma DT_t + \alpha y_{t-1} + \sum_{i=1}^{k} c_i \Delta y_{t-i} + u_t. \quad (4.3)$$

where $DU_t = 1(t > T_b), DT_t = 1(t > T_b)(t - T_b), D(T_b)t = 1(t = T_b + 1), 1(.)$ is the indicator function, and $u_t$ is the residual term assumed white noise.

For the additive outlier model, VP follow a two-step procedure. First, the series are detrended using the estimates of the following OLS regressions where $DU_t = 1(t > T_b), DT_t = 1(t > T_b)(t - T_b)$ and $1(.)$ is the indicator function:

$$y_t = \mu + \beta t + \theta DU_t + \tilde{y}_t^{1}, \quad (4.4)$$

$$y_t = \mu + \beta t + \theta DU_t + \gamma DT_t + \tilde{y}_t^{2}, \quad (4.5)$$
\[ y_t = \mu + \beta t + \gamma DT_t + \tilde{y}_t^3, \quad (4.6) \]

The test of the null hypothesis of a unit root is then performed using the t-statistic for testing \( \alpha = 1 \) in the regressions:

\[ \tilde{y}_t^j = \sum_{i=0}^{k} \omega_i D(T_b)_{t-i} + \alpha \tilde{y}_{t-1}^j + \sum_{i=1}^{k} c_t \Delta \tilde{y}_{t-1}^j + u_t, \quad (j = 1, 2) \quad (4.7) \]

\[ \tilde{y}_t^3 = \alpha \tilde{y}_{t-1}^3 + \sum_{i=1}^{k} c_t \Delta \tilde{y}_{t-1}^3 + u_t, \quad (4.8) \]

where \( D(T_b)_t = 1(t = T_b + 1) \) and \( u_t \) is the residual term assumed white noise.\(^2\)

In all the regressions \( T_b \) (the breakdate used), and \( k \) (the lag length of the autoregression) are treated as unknown.

For all the statistics, the appropriate values of the breakdate and the truncation lag parameter are unknown. Various data-dependent methods to select these values endogenously. VP propose two approaches for choosing \( T_b \). The first approach involves choosing \( T_b \) such that the statistics \( t_\alpha \) is minimised. This choice of \( T_b \) corresponds to the breakdate that is more likely to reject the unit-root hypothesis. The second approach involves choosing \( T_b \) to maximise some statistics that tests the significance of one or more of the break parameters (\( \theta, \gamma \)). For the AO and IO Models 1, \( T_b \) is chosen using the maxima of \( t_\alpha \) and \( |t_\alpha| \). The maximum of \( |t_\theta| \) is used when the direction of the break is unknown while the maximum of \( t_\theta \) is used when the direction of the break is known a priori to be positive. If the direction of the break is known to be negative, then the minimum of \( t_\theta \) is used.

\(^2\) The one-time \( k+1 \) dummy variables \( D(T_b)_{t-i} \) (\( i = 0, \ldots, k \)) in (4.7) are not needed in Model 3 as the t-statistic on \( \alpha \) is asymptotically invariant to the correlation structure of the data with the appropriate choice of \( k \) (see Perron and Vogelsang (1993a) for details).
When appropriate the assumption that the direction of the test is known can lead to more powerful tests (see Perron (1997), and Perron and Vogelsang (1992)). For the AO and IO Models 2, $T_b$ is chosen as the argmax of $t_\gamma$, $|t_\gamma|$ and $F_{\hat{\theta}, \hat{\gamma}}$ (the argmin of $t_\gamma$ if the break is negative). $F_{\hat{\theta}, \hat{\gamma}}$ is the statistics for testing the joint hypothesis that $\theta = \gamma = 0$. For the AO and IO Models 3, $T_b$ is chosen as the argmax of $t_\gamma$ and $|t_\gamma|$ (or argmin of $t_\gamma$ for negative breaks).

The truncation lag parameter $k$ can be chosen using data-dependent and data-independent methods. Following Perron (1989; 1997) and Perron and Vogelsang (1992a), $k$ may be chosen so that, for any given value of $T_b$, the coefficient on the last included lagged first difference in high-order autoregressions is insignificant up to some a priori specified maximum lag length denoted by $k_{\text{max}}$. Asymptotic normality of the $t$-statistic on these coefficients is used to carry out the inference. This procedure is denoted $k(t - \text{sig})$ in VP. In the second data-dependent procedure, the truncation lag is chosen using either the Akaike (AIC) or Schwartz (SC) information criterion where $k$ ranges from 0 to $k_{\text{max}}$. The choice of $k$ based on these criteria are denoted by $k(\text{aic})$ and $k(\text{bic})$ in VP (see Ng and Perron (1995) for theoretical justifications of these procedures).

Said and Dickey (1984) consider a third procedure to select the truncation lag parameter. This method is based on testing whether additional lags are jointly significant using a $F$-test on the estimated coefficients.\(^3\) Alternatively, $k$ may be

\(^3\) First $k_{\text{max}}$ is specified, then the autoregressions with $k_{\text{max}}$ and $(k_{\text{max}} - 1)$ lags are estimated. An $F$-test is used to assess whether the coefficient on the $k_{\text{max}}$th lag is significant and if so, the value of $k$ chosen is this maximum value. If not, the model is estimated with $(k_{\text{max}} - 2)$ lags. The lag $(k_{\text{max}} - 1)$ is deemed significant if either the $F$-test for $(k_{\text{max}} - 2)$ versus $(k_{\text{max}} - 1)$ lags or the $F$-test for $(k_{\text{max}} - 2)$ versus $k_{\text{max}}$ lags are significant. This is repeated by lowering $k$ until a rejection that additional lags are insignificant occurs or some lower
determined \textit{a priori}, by treating it as a fixed function of the sample size $T$, thus specifying a choice of the parameter independent of the data. However, as argued in Perron (1997), there is some evidence that using data-dependent methods to select the truncation lag parameter leads to test statistics having better properties (size and higher power) than if a fixed $k$ is chosen \textit{a priori} (unless of course the value of $k$ which is best happens to be selected).

4.3 Empirical results

It seems reasonable to allow a structural change to take a period of time to take effect. In what follows we focus on the Vogelsang and Perron IO Model 2 (Zivot and Andrews Model C) in which the break occurs gradually. Model 2 (crash/changing growth) is chosen because it encompasses Model 1 (crash) and Model 3 (changing growth). The series considered are RM0 (real M0), RM1 (real M1), RM2 (real M2), RGDP (real GDP), MMR (money market rate), TDR (3 months time deposit rate), USTBR (3 months US T-bill rate), INF (inflation rate), NCDR (currency depreciation rate in nominal terms), RCDR (currency depreciation rate in real terms). These series have been used in the conventional specification of the money demand function for Indonesia. All variables are measured in levels and converted to logs (see Appendix A for details). The series are shown in Figure 4.1. For completeness, additional results for Model 2 when the break is modelled as an additive outlier (i.e. occurs suddenly) are provided in Appendix B.\footnote{Results for the AO Models 1 and 3 are also available upon request.} It is worth

\footnote{Results for the AO Models 1 and 3 are also available upon request.}
Figure 4.1: Plots of the time series
noting that the results obtained under the AO assumption are by and large similar to those obtained under the IO assumption.

We estimated the IO Model 2 equation described in the previous section over all possible breakdates, $T_b$:

$$ y_t = \mu + \beta t + dD_t + \theta DU_t + \gamma DT_t + \alpha y_{t-1} + \sum_{i=1}^{k} \phi_i \Delta y_{t-i} + u_t \quad (4.9) $$

where $DU_t = 1$ for $t > T_b$, 0 otherwise, $D_t = 1$, $t = T_b$, 0 otherwise, $DT_t = (t - T_b)$ for $t > T_b$, 0 otherwise, and $u_t$ is the residual term assumed white noise. $k$ is the optimal lag length, and following Perron (1989; 1997), and Perron and Vogelsang (1992), $k$ is selected using the significant $t$ method. Starting with maximum lag length of 9, Equation (4.9) is estimated for each breakdate. Using the appropriate lag length, Equation (4.9) is estimated sequentially over all possible breakdate $T_b$, $T_b = k_0, ..., T - k_0$, where $k_0 = \delta T$, $T$ is the sample size and $\delta$ is the trimming parameter adopted. Here, $T = 76$ for each series, and we set $\delta = 0.20$ so that $T_b = 1986:3, ..., 1998:2$. The date of the break is chosen in three different ways, using the argmin of $t_\gamma$, the argmax of $|t_\gamma|$, and the argmax of $F_{\delta,l}$. After identifying the appropriate breakdate, the unit root hypothesis is tested using the $t$-statistic for $\alpha = 1$. Perron (1997) reports finite sample critical $t$-values for testing $\alpha = 1$ when the breakdate is chosen using the argmin of $t_\gamma$ and argmax of $|t_\gamma|$, and Vogelsang

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5 The choice of the maximum lag length is somewhat arbitrary. As Perron (1997) argued, on the one hand, one would like a large value to have as unrestricted a procedure as possible. On the other hand, a large value yields problems of multicollinearity in the data and also a substantial loss of power.

6 Some trimming is required because in the presence of the endpoints of the sample, the asymptotic distribution of the statistics diverges to infinity. Consequently, the procedures cannot pick up any breaks in the series occurring before the third quarter of 1986, which could possibly be related to the June 1983 reform package discussed in the previous section.
and Perron (1998) report finite sample critical values when the breakdate is chosen using the argmax of $F_{\theta, \gamma}$.

The results for each series are reported in Tables 4.1-4.3.\(^7\) Note first that the null hypothesis of a unit root in RM0, RM1, RM2, and USTBR cannot be rejected, even at the 10% level, irrespective of the method chosen to select the breakdate. In contrast, we can reject the unit root null for NCDR and RCDR in all three cases. When the breakdate is chosen using the argmin of $t_4$, the unit root null is only rejected for two series, NCDR and RCDR. Note that this is not surprising as the selection procedure imposes the a priori restriction that the direction of the break is negative, which is unlikely with MMR, TDR, and INF. The breakdates for RCDR coincide and show a break located in the third quarter of 1997.

The estimated breakdates for NCDR are located in the second and third quarter of 1997.\(^8\) Consider now the case where the procedures to select the breakdate do not impose the restriction of a one-sided change, that is, when the break is chosen by maximising $|t_4|$ (as shown in Table 4.2) or by maximising $F_{\theta, \gamma}$ (as shown in Table 4.3).

The unit root null hypothesis is now rejected for three more series, MMR, TDR, and INF. The breakdate for MMR is located in the second quarter of 1997.

\(^7\) All unit root tests are performed using GAUSS 6.0 for Windows, kernel review 6.0.8, GUI review 6.0.3 and the procedures included in Blini version 1.03 by Charemza and Makarova (2002).

\(^8\) While the second quarter of 1997 does not exactly correspond to the start of the Asian crisis (i.e. third quarter of 1997), the economic interpretation remains the same. The selection of 1997:2 is due to the presence of the dummy variable $D_t$ in (4.9). Hence, 1997:2 may be chosen because the dummy variable takes value 1 in 1997:3 and offers some additional fit to the 1997:3 crash over what the change in the intercept can do alone.
Table 4.1: 10 Model 2 results

Choosing the breakdate minimising $t_\gamma$

Notes: † denotes rejection of the unit root null hypothesis at the 0.01 nominal test size. Critical values of the $t_\alpha$ statistics are from Perron (1997), Table 1 (using $T=100$).
<table>
<thead>
<tr>
<th></th>
<th>$t_\delta$</th>
<th>$t_\gamma$</th>
<th>$F_{\delta,\gamma}$</th>
<th>$t_\alpha$</th>
<th>Break date</th>
</tr>
</thead>
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<tr>
<td>RM0</td>
<td>0.94</td>
<td>3.22</td>
<td>4.89</td>
<td>2.00</td>
<td>1989:1</td>
</tr>
<tr>
<td>RM1</td>
<td>1.94</td>
<td>4.08</td>
<td>8.27</td>
<td>4.23</td>
<td>1996:1</td>
</tr>
<tr>
<td>RM2</td>
<td>1.04</td>
<td>3.76</td>
<td>6.73</td>
<td>3.31</td>
<td>1996:1</td>
</tr>
<tr>
<td>RGDP</td>
<td>1.30</td>
<td>3.44</td>
<td>5.87</td>
<td>3.89</td>
<td>1994:3</td>
</tr>
<tr>
<td>MMR</td>
<td>11.20</td>
<td>6.40</td>
<td>58.87</td>
<td>9.31↑</td>
<td>1997:2</td>
</tr>
<tr>
<td>TDR</td>
<td>5.89</td>
<td>4.72</td>
<td>16.48</td>
<td>5.12↑↑↑</td>
<td>1998:1</td>
</tr>
<tr>
<td>USTBR</td>
<td>1.79</td>
<td>2.19</td>
<td>2.46</td>
<td>1.88</td>
<td>1998:1</td>
</tr>
<tr>
<td>INF</td>
<td>12.27</td>
<td>10.45</td>
<td>70.91</td>
<td>12.97↑</td>
<td>1997:3</td>
</tr>
<tr>
<td>NCDR</td>
<td>7.11</td>
<td>6.67</td>
<td>24.02</td>
<td>8.52↑</td>
<td>1997:3</td>
</tr>
<tr>
<td>RCDR</td>
<td>8.28</td>
<td>7.68</td>
<td>32.40</td>
<td>9.57↑↓</td>
<td>1997:3</td>
</tr>
</tbody>
</table>

Table 4.2: IO Model 2 results (continued)

Choosing the breakdate maximising $|t_\gamma|$

Notes: ↑ and ↑↑↑ denote rejection of the unit root null hypothesis at the 0.01 and 0.10 nominal test size respectively. Critical values of the $t_\alpha$ statistics are from Perron (1997), Table 1 (using $T=100$).
while the breakdate for TDR is located in the first quarter of 1998.\textsuperscript{9} The null of a unit root in RGDP is also rejected when the break is chosen by maximising $F_{\theta,4}$, with the estimated breakdate located in the fourth quarter of 1997.

Overall, the results show that the null hypothesis of a unit root can be rejected for six out the nine Indonesian series under investigation. When the unit root hypothesis is rejected, the date of the break suggests in most cases a shift located in one of the four quarters of the crisis.

For comparison only, the test proposed by Leybourne (1995) is also used. The results are shown in Table 4.4. The test is based on forward and reverse augmented Dickey-Fuller (ADF) regressions. Critical values are obtained from Leybourne (1995, p. 565). Leybourne showed that this test offers a 15 percent increase in power relative to the standard ADF test. The unit root hypothesis is rejected for MMR, TDR, INF, NCDR, and RCDR, but not for RGDP.

\section*{4.4 Conclusions}

Previous studies of money demand in Indonesia have worked under the maintained assumption that the trend in the series which enter the money demand function could be characterised as a random walk. The hypothesis of a random walk trend implies that the sum of the autoregressive coefficients equals one. In other words there is a unit root in the autoregressive polynomial. The implication of this hypothesis is that the trend is moved by random shocks and then stays at the new level until disturbed by another random shock.

\footnote{See previous footnote.}
### Table 4.3: IO Model 2 results (continued)

<table>
<thead>
<tr>
<th></th>
<th>( t_{\hat{\theta}} )</th>
<th>( t_{\gamma} )</th>
<th>( F_{\hat{\theta},\gamma} )</th>
<th>( t_{\delta} )</th>
<th>Break date</th>
</tr>
</thead>
<tbody>
<tr>
<td>RM0</td>
<td>0.60</td>
<td>3.01</td>
<td>4.94</td>
<td>-3.03</td>
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</tr>
<tr>
<td>RM1</td>
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<td>-0.66</td>
<td>17.71</td>
<td>-3.60</td>
<td>1998:1</td>
</tr>
<tr>
<td>RM2</td>
<td>-2.70</td>
<td>-1.48</td>
<td>7.50</td>
<td>-2.58</td>
<td>1998:1</td>
</tr>
<tr>
<td>RGDP</td>
<td>13.21</td>
<td>-1.71</td>
<td>81.96</td>
<td>-7.18†</td>
<td>1997:4</td>
</tr>
<tr>
<td>MMR</td>
<td>11.20</td>
<td>-6.40</td>
<td>58.87</td>
<td>-9.31†</td>
<td>1997:2</td>
</tr>
<tr>
<td>TDR</td>
<td>5.89</td>
<td>-4.72</td>
<td>16.48</td>
<td>-5.12†††</td>
<td>1998:1</td>
</tr>
<tr>
<td>USTBR</td>
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<td>4.82</td>
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<td>1996:1</td>
</tr>
<tr>
<td>INF</td>
<td>12.27</td>
<td>-10.45</td>
<td>70.91</td>
<td>-12.97†</td>
<td>1997:3</td>
</tr>
<tr>
<td>RCDR</td>
<td>-8.28</td>
<td>7.68</td>
<td>32.40</td>
<td>-9.57†</td>
<td>1997:3</td>
</tr>
</tbody>
</table>

**Choosing the breakdate maximising \( F_{\hat{\theta},\gamma} \)**

Notes: † and ††† denote rejection of the unit root null hypothesis at the 0.01 and 0.10 nominal test size respectively. Critical values of the \( t_{\delta} \) statistics are from Vogelsang and Perron (1998), Table 3 (using \( T=100 \)).

---

**Choosing the breakdate maximising \( F_{\hat{\theta},\gamma} \)**

Notes: † and ††† denote rejection of the unit root null hypothesis at the 0.01 and 0.10 nominal test size respectively. Critical values of the \( t_{\delta} \) statistics are from Vogelsang and Perron (1998), Table 3 (using \( T=100 \)).
<table>
<thead>
<tr>
<th></th>
<th>( ADF_{\text{max}} )</th>
</tr>
</thead>
<tbody>
<tr>
<td>RM0</td>
<td>-1.68</td>
</tr>
<tr>
<td>RM1</td>
<td>-1.48</td>
</tr>
<tr>
<td>RM2</td>
<td>-0.29</td>
</tr>
<tr>
<td>RGDP</td>
<td>-1.64</td>
</tr>
<tr>
<td>MMR</td>
<td>-2.49(^{††})</td>
</tr>
<tr>
<td>TDR</td>
<td>-2.64(^{††})</td>
</tr>
<tr>
<td>USTBR</td>
<td>-0.69</td>
</tr>
<tr>
<td>INF</td>
<td>-2.76(^{††})</td>
</tr>
<tr>
<td>NCDR</td>
<td>-2.45(^{††})</td>
</tr>
<tr>
<td>RCDR</td>
<td>-3.14(^{†})</td>
</tr>
</tbody>
</table>

**Table 4.4: Leybourne (1995) unit root tests**

Notes: An intercept and trend are included in the Dickey-Fuller regressions for RM0, RM1, RM2, RGDP. An intercept only is included in the Dickey-Fuller regressions for MMR, TDR, USTBR, INF, NCDR, and RCDR. \(^{†}\) and \(^{††}\) denote rejection of the unit root null hypothesis at the 0.01 and 0.05 nominal test size respectively (using \( T = 100 \)). Critical values of the \( ADF_{\text{max}} \) statistics are obtained from Leybourne (1995).
Our results challenge this assumption by providing empirical evidence that for most of the series, the movement of the trend can be explained by a single structural break in an otherwise constant linear trend. This explanation is plausible since a trend break produces serial correlation properties that are similar to those of a random walk. In effect, the results reverse inference for six out of the nine Indonesian series under investigation. For these series, the timing of the break typically occurs at the end of 1997 and coincides with the crisis and a major government intervention in the money market for which the precedent is the Sumarlin shock of 1991. The results do not reverse inference for the real money stock variables (M0, M1, and M2). One important implication is that the series, with the exception of the monetary aggregates, should not be differenced to achieve stationarity. Instead, they should be detrended using the estimated breakdates. Another implication concerns the significance of the coefficients, which can be tested using the standard Student t distribution. Finally, the finding that the series are stationary after accounting for structural change possibly simplifies the query regarding the predictability of the variables entering the money demand function by removing some of the problems associated with forecasting from nonstationary series. The results also suggest that the cointegration analyses of previous studies of Indonesian money demand, which treated all the variables as unit root processes, could be based on regressions containing mixtures of (trend-break) stationary and nonstationary variables. The findings of such studies should therefore be taken with some caution.
For each series the sample period is 1983:1-2001:4. The series are all seasonally adjusted (with the exception of the three interest rate series).¹⁰

(1) RM0, RM1, and RM2 are the natural logarithms of the MO, M1 and M2 definitions of the money stock (in billion rupiahs), each deflated by the consumer price index (1995=100). Source: Bank Indonesia (unpublished).


(3) MMR is the nominal money market rate, TDR is the nominal time deposit rate (3 months), USTBR is the nominal US T-bill rate (3 months). The quarterly rate is computed as:

$$0.25 \times \ln[1 + (R_t/100)]$$,

where $R_t$ is either one of the interest rates above. Source: IMFs International Financial Statistics. INF is the inflation rate, computed as:

$$\ln(CPI_t) - \ln(CPI_{t-1})$$,

where $CPI_t$ is the consumer price index (1995=100). Source: IMFs International Financial Statistics. NCDR is the currency depreciation rate in nominal

¹⁰ The seasonal adjustment is performed using the Stamp package (see Koopman, Harvey, Doornik and Shephard, 1995). We adopted the Stamp manual’s recommended version (p. 88) of the basic structural model of a stochastic trend with a stochastic slope, a trigonometric seasonal and an irregular component. A cyclical component was not included in the adjustment procedure. It is worth noting the Stamp manual’s comment (p. 88) that in practice seasonal components seem to be insensitive to the specification of the trend and the inclusion of a cycle.
terms, computed as:

$$l(E_t) - l(E_{t-1}),$$

where $E_t$ is the nominal Rupiah exchange rate defined as the foreign price of domestic currency (indirect quotation, a rise represents an Indonesian currency appreciation). Source: IMFs International Financial Statistics. RCDR is the currency depreciation rate in real terms, computed as:

$$[l(WPI_t) - l(WPI^*_t) + l(E_t)] - [l(WPI_{t-1}) - l(WPI^*_t) + l(E_{t-1})],$$

where $WPI_t$ is the domestic wholesale price index (1995=100), and $WPI^*_t$ is a weighted average of foreign wholesale price indices (1995=100) computed as:

$$WPI^*_t = \sum_{j=1}^{m_r} w_j WPI_{jt},$$

where $w_j$ are fixed weights and $m_r = 4$.

The four chosen countries account for half of Indonesia's trade over the period 1997-2001. Each of the weights is computed as the average of annual trade with the country divided by annual trade with all four countries over the period 1997-2001. The countries and weights in brackets are: Japan (0.4025), the United States (0.2642), Singapore (0.2058) and Korea (0.1275). Sources: IMFs International Financial Statistics and United Nations Yearbook 2001.
Appendix B: Additional results

Additional results for Model 2 when the break is modelled as an additive outlier are reported in Tables 4.5-4.7 below.
<table>
<thead>
<tr>
<th></th>
<th>$t_r$</th>
<th>$t_\alpha$</th>
<th>Break date</th>
</tr>
</thead>
<tbody>
<tr>
<td>RM0</td>
<td>10.81</td>
<td>-2.83</td>
<td>1994:1</td>
</tr>
<tr>
<td>RM1</td>
<td>-5.99</td>
<td>-4.02</td>
<td>1995:4</td>
</tr>
<tr>
<td>RM2</td>
<td>-12.68</td>
<td>-3.75</td>
<td>1997:1</td>
</tr>
<tr>
<td>RGDP</td>
<td>7.87</td>
<td>-5.44</td>
<td>1996:2</td>
</tr>
<tr>
<td>MMR</td>
<td>-2.17</td>
<td>-8.32$t$</td>
<td>1996:2</td>
</tr>
<tr>
<td>TDR</td>
<td>-2.45</td>
<td>-5.89$t$</td>
<td>1996:3</td>
</tr>
<tr>
<td>USTBR</td>
<td>1.69</td>
<td>-2.50</td>
<td>1986:3</td>
</tr>
<tr>
<td>INF</td>
<td>-0.80</td>
<td>-9.68$t$</td>
<td>1996:1</td>
</tr>
<tr>
<td>NCDR</td>
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<td>1995:4</td>
</tr>
<tr>
<td>RCDR</td>
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<td>-10.25$t$</td>
<td>1995:4</td>
</tr>
</tbody>
</table>

Table 4.5: AO Model 2 results

Choosing the breakdate minimising $t_\alpha$

Notes: $t$ denotes rejection of the unit root null hypothesis at the 0.01 nominal test size. Critical values for $t_\alpha$ are from Vogelsang and Perron (1998), Table 2 (using $T=100$).
<table>
<thead>
<tr>
<th></th>
<th>$t_\gamma$</th>
<th>$t_\delta$</th>
<th>Break date</th>
</tr>
</thead>
<tbody>
<tr>
<td>RM0</td>
<td>13.16</td>
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<td>1992:2</td>
</tr>
<tr>
<td>RM1</td>
<td>1.65</td>
<td>-2.08</td>
<td>1989:1</td>
</tr>
<tr>
<td>RM2</td>
<td>-0.35</td>
<td>-1.87</td>
<td>1985:3</td>
</tr>
<tr>
<td>RGDP</td>
<td>11.35</td>
<td>-3.59</td>
<td>1994:1</td>
</tr>
<tr>
<td>MMR</td>
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<td>1992:4</td>
</tr>
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<td>TDR</td>
<td>0.46</td>
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<td>1993:4</td>
</tr>
<tr>
<td>USTBR</td>
<td>4.89</td>
<td>-1.73</td>
<td>1992:4</td>
</tr>
<tr>
<td>INF</td>
<td>1.02</td>
<td>-3.54</td>
<td>1986:2</td>
</tr>
<tr>
<td>NCDR</td>
<td>2.66</td>
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<td>1997:2</td>
</tr>
<tr>
<td>RCDR</td>
<td>3.21</td>
<td>-7.27†</td>
<td>1997:2</td>
</tr>
</tbody>
</table>

**Table 4.6: AO Model 2 results (continued)**

**Choosing the breakdate maximizing $t_\gamma$**

Notes: † denotes rejection of the unit root hypothesis at the 0.01 nominal test size. Critical values for $t_\delta$ are from Vogelsang and Perron (1998), Table 2 (using $T=100$).
**Table 4.7**: AO Model 2 results (continued)

Choosing the breakdate maximising $|t_\gamma|$  

<table>
<thead>
<tr>
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<th>$t_\gamma$</th>
<th>$t_\alpha$</th>
<th>Break date</th>
</tr>
</thead>
<tbody>
<tr>
<td>RM0</td>
<td>13.16</td>
<td>2.51</td>
<td>1992:2</td>
</tr>
<tr>
<td>RM1</td>
<td>6.05</td>
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<td>1996:2</td>
</tr>
<tr>
<td>RM2</td>
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<td>1996:3</td>
</tr>
<tr>
<td>RGDP</td>
<td>11.35</td>
<td>3.59</td>
<td>1994:1</td>
</tr>
<tr>
<td>MMR</td>
<td>9.74</td>
<td>4.69</td>
<td>1997:3</td>
</tr>
<tr>
<td>TDR</td>
<td>8.41</td>
<td>2.68</td>
<td>1998:2</td>
</tr>
<tr>
<td>USTBR</td>
<td>4.89</td>
<td>1.73</td>
<td>1992:4</td>
</tr>
<tr>
<td>INF</td>
<td>5.53</td>
<td>6.90$^\dagger$</td>
<td>1997:4</td>
</tr>
<tr>
<td>NCDR</td>
<td>2.66</td>
<td>7.23$^\dagger$</td>
<td>1997:2</td>
</tr>
<tr>
<td>RCDR</td>
<td>3.21</td>
<td>7.27$^\dagger$</td>
<td>1997:2</td>
</tr>
</tbody>
</table>

Notes: $^\dagger$ denotes rejection of the unit root hypothesis at the 0.01 nominal test size. Critical values for $t_\alpha$ are from Vogelsang and Perron (1998), Table 2 (using $T=100$).
CHAPTER 5

FINANCIAL LIBERALISATION AND REGIME SHIFTS IN INDONESIAN MONEY DEMAND

5.1 Introduction

Previous money demand studies for Indonesia (e.g. Dekle and Pradhan, 1999; Agrawal, 2001; Chaisrisawatsuk et al., 2004) that used cointegration analysis are somewhat inconclusive. These studies, however, have not considered the possibility that financial liberalisation may be associated with structural change, in which case standard cointegration tests could yield misleading results (Leybourne and Newbold, 2003).

Several tests have been recently developed in the statistical literature for testing cointegration in the presence of breaks but there is no consensus yet about which is more powerful and offers more robust results (Jimeno et al., 2006). To detect common structural breaks we apply the residual-based cointegration test with regime shifts of Gregory and Hansen (1996). This test has been used inter alia by Siddiki and Morrissey (2006) to investigate the stability of the demand for money in three South Asian countries (Bangladesh, India and Pakistan).

The test has two main advantages: (i) it is more powerful than standard residual-based tests for cointegration, which presume that the coefficients are time-invariant under the alternative hypothesis; and (ii) the timing of the break is endogenous, which prevents informal data analysis from contaminating the choice
of breakpoint. We follow a three-step strategy. First, we test for cointegration using a cointegrated vector autoregression (VAR) framework, assuming that the full system can be partitioned into a vector of endogenous variables and a vector of weakly exogenous variables. Second, we test for weak exogeneity of the conditioning variables in the long-run real money demand function. We show that the weak exogeneity hypothesis cannot be rejected and reduce the full system model to a conditional model, which increases efficiency and can be given a structural interpretation (Johansen, 1992a). Finally, we test for cointegration between the variables in the real money demand function allowing for the possibility of a shift of unknown timing in the cointegration vector using the test of Gregory and Hansen (1996). Unlike previous studies of Indonesian money demand, the sample period covers the critical quarters of the 1997-98 economic crisis.

The results show that the money demand function is cointegrated in the presence of structural breaks occurring (i) in the second quarter of 1991 and coinciding with the Sumarlin shock and (ii) in the last quarter of 1997 and coinciding with the crisis and a government intervention in the money market for which the Sumarlin shock of 1991 is the precedent (all discussed in chapter 3).

The rest of the chapter is organised as follows. Section 5.2 describes the model and methodology. Section 5.3 provides empirical results. Section 5.4 offers some conclusions.
5.2 Methodology

We start the following conventional theoretical money demand function (see for instance Goldfeld, 1992)

\[ m_t = f_m(p_t, y_t, R_t, \Delta p_t, e_t) \]  
\[(5.1)\]

where \( m_t \) is the money supply, \( y_t \) is the scaling variable, the vector of interest rates, \( R_t \), measures the yields on financial assets, the overall inflation rate, \( \Delta p_t \), measures the return to holding goods, and the real exchange rate, \( e_t \), captures the attractiveness of holding foreign assets.

Since money demand is not directly observable, it is assumed that the money market is always in equilibrium and that the money supply is exogenous. Thus, one can identify the money demand function and use the quantity of money to measure the demand for money (see for instance Laidler (1993) for a discussion of the identification problem of the money demand function).

The money demand function in (5.1) has the following semi-logarithmic specification

\[(m - p)_t = \beta_p y_t + \beta_R R_t + \beta_{\Delta p} \Delta p_t + \beta_e e_t + \text{intercept} \]  
\[(5.2)\]

Defining the vector \( z_t \) containing the \( n \) potentially endogenous variables discussed above as

\[ z_t = [(m - p)_t, y_t, R_t, \Delta p_t, e_t]' = [(m - p)_t, x_t]' \]  
\[(5.3)\]
Assuming that the time series properties of the variables included in $z_t$ can be well approximated by a log-linear unrestricted vector autoregression (VAR) involving up to $k$ lags of $z_t$ and augmented with intercepts and trends gives:

$$z_t = A_1 z_{t-1} + ... + A_k z_{t-k} + \Psi D_t + u_t$$  \hspace{1cm} (5.4)

where $u_t \sim \text{IN}(0, \Sigma)$, $z_t$ is $(n \times 1)$, and each of the $A_i$ is an $(n \times n)$ matrix of parameters.

The system is in reduced-form with each variable in $z_t$ regressed on only lagged values of both itself and all the other variables in the system. Thus, ordinary least-squares (OLS) is an efficient way to estimate each equation comprising (5.4) since the right-hand side of each equation in the system comprises a common set of lagged regressors.

Equation (5.4) can be formulated into a vector equilibrium correction model (VECM) form:

$$\Delta z_t = \Gamma_1 \Delta z_{t-1} + ... + \Gamma_k \Delta z_{t-k+1} + \Pi z_{t-k} + \Psi D_t + u_t$$  \hspace{1cm} (5.5)

where $\Gamma_i = -(I - A_1 - ... - A_i)$ ($i = 1, ..., k - 1$), $\Pi = -(I - A_1 - ... - A_k)$ and $u_t \sim \text{IN}(0, \Sigma)$.

This way of specifying the system contains information on both the short-run and long-run adjustment to changes in $z_t$, via the estimates of $\Gamma$ and $\Pi$, respectively, where $\Pi = \alpha \beta'$, $\alpha$ represents the speed of adjustment to disequilibrium and $\beta$ is a matrix of long-run coefficients such that the term $\beta' z_{t-k}$ embedded in (5.5) represents up to $r \leq (n - 1)$ cointegration relationships, which ensures that the $z_t$ converge with their long-run steady-state solutions.
Estimates of $\alpha$ and $\beta$ can be obtained using the Johansen method of reduced rank regression. Furthermore, Johansen (1992b) shows that if the variables in $x_t$ are weakly exogenous with respect to the parameters of the model (i.e., the $\Gamma$, $\Pi$, $\alpha$, $\beta$), there will be no loss of information in modelling the change in real money balances $\Delta(m_t - p_t)$ in the context of a conditional single-equation model of the type:

$$\Delta(m - p)_t = \sum_{i=1}^{k-1} \delta_i \Delta(m - p)_{t-i} + \sum_{i=0}^{k-1} \gamma_i \Delta x_{t-i} + \alpha((m - p)_{t-1} - \beta' x_{t-1}) + \epsilon_t, \quad (5.6)$$

where $\epsilon_t \sim \text{IID}(0, \sigma^2)$.

The parameter $\alpha$ captures a feedback effect on the change in money holdings, $\Delta(m - p)_t$, from the lagged deviation from the long-run target money holdings, $((m - p) - (m - p)^*)_t$. The target $(m - p)^*$ is defined as a linear function of the forcing variables $x_t$, that is, $(m - p)^* = \beta' x_t$.

It is hypothesised that the long-run money demand relationship embedded in (5.6) may have shifted at some unknown point in the sample.

The above proposition cannot be tested with standard residual-based tests for cointegration since they presume that the cointegration relationship is time invariant under the alternative hypothesis. Gregory and Hansen (1996)—hereafter GH—propose a class of residual-based tests allowing for a break in the intercept and/or the slope vector in the cointegration relationship. These tests are non-informative with respect to the timing of the regime shift, thus preventing informal data analysis (such as the visual examination of the time series plots) from
contaminating the choice of breakpoint.

Following GH, the long-run (static) money-demand function embedded in (5.6) is rewritten with the $\beta$ vector separated into the intercept and slope parameters $(\theta, \beta)$ as:

$$(m - p)_t = \theta_1 + \theta_2 \varphi_{tk} + \lambda t + \beta'_1 x_t + \varphi_{tk} \beta'_2 x_t + n_t$$

where $n_t$ is a stationary error term.

Structural change is included through the dummy variable:

$$\varphi_{tk} = \begin{cases} 
0 & \text{if } t \leq k \\
1 & \text{if } t > k 
\end{cases}$$

where $k$ is the unknown date of the (potential) structural break. Equation (5.7) allows for three alternatives if $(m - p)_t$ and $x_t$ cointegrate: (i) a change in the intercept but no change in the slope vector ($\lambda = \beta'_2 = 0$); (ii) a change in the intercept allowing for a deterministic trend ($\beta'_2 = 0$); and (iii) a change in the slope vector as well as a change in the intercept (with $\lambda = 0$).

Since $k$ is unknown, both the augmented Dickey-Fuller (ADF) and the Phillips-Perron (PP) tests involving $n_t$ are computed for each date within the sample (i.e., $k \in T$) with the largest negative values of the ADF $t$-statistic and the PP $Z$-statistics across all possible breakpoints taken as the relevant statistics for testing the null hypothesis. Critical values are provided by GH.
<table>
<thead>
<tr>
<th>variable</th>
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</thead>
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<tr>
<td></td>
<td>C/T</td>
<td>C</td>
</tr>
<tr>
<td>RM1</td>
<td>-2.10</td>
<td>-1.14</td>
</tr>
<tr>
<td></td>
<td>(0)</td>
<td>(0)</td>
</tr>
<tr>
<td>RM2</td>
<td>-0.49</td>
<td>-1.70</td>
</tr>
<tr>
<td></td>
<td>(4)</td>
<td>(4)</td>
</tr>
<tr>
<td>RGDP</td>
<td>-2.20</td>
<td>0.86</td>
</tr>
<tr>
<td></td>
<td>(4)</td>
<td>(4)</td>
</tr>
<tr>
<td>MMR</td>
<td>-1.91</td>
<td>-1.91</td>
</tr>
<tr>
<td></td>
<td>(0)</td>
<td>(0)</td>
</tr>
<tr>
<td>INF</td>
<td>-3.64**</td>
<td>-3.55**</td>
</tr>
<tr>
<td></td>
<td>(0)</td>
<td>(0)</td>
</tr>
<tr>
<td>USTBR</td>
<td>-2.60</td>
<td>-0.89</td>
</tr>
<tr>
<td></td>
<td>(1)</td>
<td>(8)</td>
</tr>
<tr>
<td>RER</td>
<td>-1.95</td>
<td>-0.18</td>
</tr>
<tr>
<td></td>
<td>(1)</td>
<td>(1)</td>
</tr>
</tbody>
</table>


Notes: The lag length was selected using the Modified Akaike Information Criterion (MAIC). The maximum lag length was chosen using the formula reported in Schwert (1989, p. 151): \( k_{max} = int\left(12(T/100)^{1/4}\right) \). The bandwidth was selected using the Newey-West (1994) method based on the Bartlett kernel. * denotes rejection of the null hypothesis at least at the 10% level. ** denotes rejection of the null hypothesis at least at the 5% level.
<table>
<thead>
<tr>
<th>variable</th>
<th>ADF</th>
<th>PP</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>C/T</td>
<td>C</td>
</tr>
<tr>
<td>(and lag length/bandwidth)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>ΔRM1</td>
<td>-4.81**</td>
<td>-4.82**</td>
</tr>
<tr>
<td></td>
<td>(1)</td>
<td>(1)</td>
</tr>
<tr>
<td>ΔRM2</td>
<td>-9.14**</td>
<td>-1.77</td>
</tr>
<tr>
<td></td>
<td>(0)</td>
<td>(7)</td>
</tr>
<tr>
<td>ΔRGDP</td>
<td>-3.54**</td>
<td>-3.53**</td>
</tr>
<tr>
<td></td>
<td>(1)</td>
<td>(1)</td>
</tr>
<tr>
<td>ΔMMR</td>
<td>-3.74**</td>
<td>-3.77**</td>
</tr>
<tr>
<td></td>
<td>(1)</td>
<td>(1)</td>
</tr>
<tr>
<td>ΔINF</td>
<td>-5.16**</td>
<td>-5.20**</td>
</tr>
<tr>
<td></td>
<td>(1)</td>
<td>(1)</td>
</tr>
<tr>
<td>ΔUSTBR</td>
<td>-1.77</td>
<td>-2.12</td>
</tr>
<tr>
<td></td>
<td>(4)</td>
<td>(2)</td>
</tr>
<tr>
<td>ΔRER</td>
<td>-5.45**</td>
<td>-5.44**</td>
</tr>
<tr>
<td></td>
<td>(0)</td>
<td>(0)</td>
</tr>
</tbody>
</table>

Table 5.2: Augmented Dickey-Fuller and Phillips-Perron tests (continued)

See notes Table 5.1.
5.3 Empirical results

Consider the model set out above using Indonesian quarterly data from 1989:1 to 2001:4.\(^1\) \((m - p)_t\) is proxied throughout by the real M1 money balances denoted \(RM_1\), comprising cash plus the non-interest bearing demand deposits at the commercial banks, or alternatively, by real M2 money balances denoted \(RM_2\), comprising M1 plus the interest bearing demand deposits at the commercial banks. The real income variable \(y_t\) is proxied by real GDP denoted \(RGDP\). The interest rate variable \(R_t\) is proxied by the money market rate denoted \(MMR\); \(\Delta p_t\), is proxied by the change in the consumer price index, which measures inflation and is denoted \(INF\). In addition, the effects of capital account openness and currency substitution on the demand for money are introduced via a foreign interest rate variable proxied by the US treasury bill rate denoted \(USTBR\), and a real exchange rate variable denoted \(RER\) (see the appendix for the exact definitions).

Tables 5.1 and 5.2 report Augmented Dickey-Fuller (ADF) and Phillips and Perron (PP) unit root tests. On the basis of the ADF and PP tests applied to each series on levels, \(RM_1\), \(RM_2\), \(RGDP\), \(MMR\), \(USTBR\) and \(RER\) appear to be \(I(1)\) since one fails to reject the null of nonstationarity. In contrast, \(INF\) appears to be \(I(0)\). When taking the first difference of the series, the results tend to confirm the hypothesis that each series is \(I(1)\), as differencing removes

\(^1\) *EViews 4.0* (Quantitative Micro Software, LLC) is used to perform the unit root tests. *PcGive 10.3* is used throughout to carry out the estimations (see Doornik and Hendry, 2001). The residual-based tests for cointegration in models with regime shifts are performed with the GAUSS programme SHIFTS written by Bruce E. Hansen using GAUSS 6.0 for Windows, kernel review 6.0.8, GUI review 6.0.3.
<table>
<thead>
<tr>
<th>variable</th>
<th>DF-GLS</th>
<th>NP</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(and lag length)</td>
<td>C/T</td>
</tr>
<tr>
<td>RM1</td>
<td>-2.14</td>
<td>0.59</td>
</tr>
<tr>
<td></td>
<td>(0)</td>
<td>(0)</td>
</tr>
<tr>
<td>RM2</td>
<td>-1.01</td>
<td>-0.12</td>
</tr>
<tr>
<td></td>
<td>(4)</td>
<td>(8)</td>
</tr>
<tr>
<td>RGDP</td>
<td>-1.43</td>
<td>0.22</td>
</tr>
<tr>
<td></td>
<td>(4)</td>
<td>(8)</td>
</tr>
<tr>
<td>MMR</td>
<td>-1.96</td>
<td>-1.92*</td>
</tr>
<tr>
<td></td>
<td>(0)</td>
<td>(6)</td>
</tr>
<tr>
<td>INF</td>
<td>-3.70**</td>
<td>-3.48**</td>
</tr>
<tr>
<td></td>
<td>(0)</td>
<td>(0)</td>
</tr>
<tr>
<td>USTBR</td>
<td>-2.78</td>
<td>0.06</td>
</tr>
<tr>
<td></td>
<td>(1)</td>
<td>(10)</td>
</tr>
<tr>
<td>RER</td>
<td>-1.97</td>
<td>0.73</td>
</tr>
<tr>
<td></td>
<td>(1)</td>
<td>(1)</td>
</tr>
</tbody>
</table>

Table 5.3: Dickey-Fuller GLS de-trended and Ng-Perron tests: Indonesian money demand data (1986:1-2001:4), seasonaly adjusted

Notes: DF-GLS refers to the ADF test with GLS de-trending. Critical values are from Table 1 in Elliott, Rothenberg, and Stock (1996). The lag length was selected using the Modified Akaike Information Criterion (MAIC). NP refers to the NG-Perron test based on the modified PP Z statistic with de-trending. Critical values are from Table 1 in Ng-Perron (2001). The lag length was chosen using the AR spectral GLS-detrended estimator based on the MAIC. The maximum lag length was chosen using the Schwert formula (see notes Table 5.1).* denotes rejection of the null hypothesis at least at the 10% level.** denotes rejection of the null hypothesis at least at the 5% level.
<table>
<thead>
<tr>
<th>variable</th>
<th>DF-GLS</th>
<th>NP</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>C/T</td>
<td>C</td>
</tr>
<tr>
<td>ΔRM1</td>
<td>-4.87**</td>
<td>-4.84**</td>
</tr>
<tr>
<td></td>
<td>(1)</td>
<td>(1)</td>
</tr>
<tr>
<td>ΔRM2</td>
<td>-1.64</td>
<td>1.41</td>
</tr>
<tr>
<td></td>
<td>(7)</td>
<td>(7)</td>
</tr>
<tr>
<td>ΔRGDP</td>
<td>-2.14</td>
<td>-1.00</td>
</tr>
<tr>
<td></td>
<td>(7)</td>
<td>(7)</td>
</tr>
<tr>
<td>ΔMMR</td>
<td>-3.76**</td>
<td>-2.36**</td>
</tr>
<tr>
<td></td>
<td>(1)</td>
<td>(8)</td>
</tr>
<tr>
<td>ΔINF</td>
<td>-4.99**</td>
<td>-0.95</td>
</tr>
<tr>
<td></td>
<td>(1)</td>
<td>(10)</td>
</tr>
<tr>
<td>ΔUSTBR</td>
<td>-1.28</td>
<td>-1.03</td>
</tr>
<tr>
<td></td>
<td>(4)</td>
<td>(4)</td>
</tr>
<tr>
<td>ΔRER</td>
<td>-1.57</td>
<td>-0.67</td>
</tr>
<tr>
<td></td>
<td>(10)</td>
<td>(10)</td>
</tr>
</tbody>
</table>

Table 5.4: Dickey-Fuller GLS de-trended and Ng-Perron tests (continued)

See Notes Table 5.3.
the unit-root, although there is some evidence that \( RM2 \) and \( USTBR \) may not be stationary after differencing, suggesting these series may contain two unit roots. ADF and PP tests are usually considered to have poor power. For improvements in power, Elliott, Rothenberg and Stock (1996) have shown that the power of the ADF test can be optimised using a form of de-trending known as generalised least squares (GLS) de-trending. Ng and Perron (2001) have also suggested to incorporate a new information criterion for setting the lag length along with GLS de-trending. Thus, both Dickey-Fuller de-trended (DF-GLS) and Ng and Perron (NP) tests are employed as additional tests. The results are presented in Tables 5.3 and 5.4. These results mainly confirm that each series is \( I(1) \), except for \( RM2 \), \( INF \) and \( USTBR \). Consequently, these variables are excluded from the subsequent analysis. Arguably, and recalling the findings of the previous chapter, some of the series under investigation could well be broken trend stationary. Nonetheless, since the emphasis of this chapter is on the detection of a possible shift in the money demand, we proceed under the maintained assumption that the series which enter the money demand function are all \( I(1) \).

Leaving the cointegration rank unrestricted as \( r = 4 \), OLS are used to estimate the system denoted by (5.5), restricting \( D_t \) to include only an intercept. Setting \( k = 2 \) produces the output in Table 5.5.\(^2\) The diagnostic tests involve \( F \)-tests for

\(^2\) The choice of \( k \) based on the Akaike information criterion (AIC) would in fact result in \( k = 4 \), whereas in contrast the Schwartz criterion (SC) and Hannan-Quinn (HQ) criterion suggest \( k = 1 \) and \( k = 3 \), respectively. When information criteria suggest different values of \( k \), Johansen et al. (2000) note that it is common practice to prefer the HQ criterion. However, like others, we have set \( k = 2 \) in our subsequent analysis mainly because setting \( k \) at different values results in implausible estimates of the cointegration vectors.
<table>
<thead>
<tr>
<th>Statistic</th>
<th>RM1</th>
<th>RGDP</th>
<th>MMR</th>
<th>RER</th>
</tr>
</thead>
<tbody>
<tr>
<td>$F_{k=1}(4,40)$</td>
<td>5.59**</td>
<td>4.39**</td>
<td>4.36**</td>
<td>10.82**</td>
</tr>
<tr>
<td>$F_{k=2}(4,40)$</td>
<td>1.25**</td>
<td>2.85**</td>
<td>0.81</td>
<td>1.28</td>
</tr>
<tr>
<td>$\sigma$</td>
<td>4.90%</td>
<td>2.98%</td>
<td>4.92%</td>
<td>15.15%</td>
</tr>
<tr>
<td>$F_{ar}(4,39)$</td>
<td>6.18**</td>
<td>1.51</td>
<td>1.15</td>
<td>0.82</td>
</tr>
<tr>
<td>$F_{arch}(4,35)$</td>
<td>0.56</td>
<td>0.04</td>
<td>0.11</td>
<td>2.30</td>
</tr>
<tr>
<td>$F_{het}(20,26)$</td>
<td>0.68</td>
<td>2.36*</td>
<td>0.68</td>
<td>4.84**</td>
</tr>
<tr>
<td>$\chi^2_{nd}(2)$</td>
<td>3.08</td>
<td>17.30**</td>
<td>67.51**</td>
<td>10.45**</td>
</tr>
</tbody>
</table>

Multivariate tests: $F_{ar}(64, 96) = 3.72**; F_{het}(160, 164) = 1.37**; \chi^2_{nd}(8) = 123.69**;
$F_{uv}(32, 149) = 136.767**; \text{AIC} = -23.2173; \text{SC} = -21.8664; \text{HQ} = -22.6994

Table 5.5: Indonesian money demand model evaluation diagnostics, 1989:1-2001:4

Notes: * denotes rejection of the null hypothesis at 5% significance level. ** denotes rejection of the null hypothesis at 1% significance level.
<table>
<thead>
<tr>
<th>Statistic</th>
<th>RM1</th>
<th>RGDP</th>
<th>MMR</th>
<th>RER</th>
</tr>
</thead>
<tbody>
<tr>
<td>$F_{k=1}(4, 36)$</td>
<td>7.91**</td>
<td>6.89**</td>
<td>4.83**</td>
<td>7.82**</td>
</tr>
<tr>
<td>$F_{k=2}(4, 36)$</td>
<td>0.96</td>
<td>1.79</td>
<td>5.68**</td>
<td>1.23</td>
</tr>
<tr>
<td>$F_{k=3}(4, 36)$</td>
<td>4.53**</td>
<td>2.25</td>
<td>11.96**</td>
<td>1.50</td>
</tr>
<tr>
<td>$\sigma$</td>
<td>3.96%</td>
<td>2.17%</td>
<td>4.84%</td>
<td>13.88%</td>
</tr>
<tr>
<td>$F_{ar}(4, 35)$</td>
<td>1.79</td>
<td>1.83</td>
<td>0.61</td>
<td>0.64</td>
</tr>
<tr>
<td>$F_{arch}(4, 31)$</td>
<td>1.37</td>
<td>0.68</td>
<td>0.02</td>
<td>3.10*</td>
</tr>
<tr>
<td>$F_{het}(24, 14)$</td>
<td>0.87</td>
<td>0.73</td>
<td>0.51</td>
<td>3.56**</td>
</tr>
<tr>
<td>$\chi^2_{rd}(2)$</td>
<td>1.04</td>
<td>0.15</td>
<td>54.37**</td>
<td>1.72</td>
</tr>
</tbody>
</table>

Multivariate tests: $F_{ar}(64, 80) = 2.34**$; $F_{het}(240, 70) = 0.60$; $\chi^2_{rd}(8) = 88.11^{**}$; $F_{ur}(48, 140) = 105.178^{**}$; AIC = $-23.9129$; SC = $-21.9616$; HQ = $-23.1648$

Table 5.6: Indonesian money demand model evaluation diagnostics (continued)

See Notes Table 5.5
Table 5.7: Indonesian money demand model evaluation diagnostics when dummy variables are included

<table>
<thead>
<tr>
<th>Statistic</th>
<th>RM1</th>
<th>RGDP</th>
<th>MMR</th>
<th>RER</th>
</tr>
</thead>
<tbody>
<tr>
<td>Lag length = 2; Intercept, Dum97:3, Dum98:1, and Dum98:2</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$F_{k=1}(4, 37)$</td>
<td>4.68**</td>
<td>9.99**</td>
<td>22.33**</td>
<td>26.75**</td>
</tr>
<tr>
<td>$F_{k=2}(4, 37)$</td>
<td>1.55</td>
<td>3.09*</td>
<td>6.03**</td>
<td>7.11**</td>
</tr>
<tr>
<td>$\sigma$</td>
<td>4.24%</td>
<td>2.15%</td>
<td>2.02%</td>
<td>7.71%</td>
</tr>
<tr>
<td>$F_{ar}(4, 36)$</td>
<td>3.33*</td>
<td>1.36</td>
<td>0.44</td>
<td>2.57</td>
</tr>
<tr>
<td>$F_{arch}(4, 32)$</td>
<td>0.35</td>
<td>0.24</td>
<td>1.08</td>
<td>2.71*</td>
</tr>
<tr>
<td>$F_{het}(16, 23)$</td>
<td>0.58</td>
<td>0.43</td>
<td>0.97</td>
<td>7.30**</td>
</tr>
<tr>
<td>$\chi^2_{nd}(2)$</td>
<td>1.94</td>
<td>0.63</td>
<td>6.76*</td>
<td>3.80</td>
</tr>
</tbody>
</table>

Multivariate tests: $F_{ar}(64, 84) = 2.56**; F_{het}(160, 138) = 0.97; \chi^2_{nd}(8) = 14.491;$

$F_{ar}(32, 138) = 245.839**; \text{AIC} = -26.4517; \text{SC} = -24.6505; \text{HQ} = -25.7612$

Notes: * denotes rejection of the null hypothesis at 5% significance level. ** denotes rejection of the null hypothesis at 1% significance level.
the hypotheses: that the $i$-period lag ($F_{k=i}$) is zero; that there is no serial correlation ($F_{ar}$, against fourth-order autoregression); that there is no autoregressive conditional heteroskedasticity (ARCH) ($F_{arch}$, against fourth order); that there is no heteroskedasticity ($F_{het}$); and lastly a $\chi^2$-test for normality ($\chi^2_{ind}$). Analogous system (vector) tests are also reported, with the last test $F_{ur}$ representing the test against the significance of the regressors in $D_t$.

Figure 5.1: Residuals (scaled) and residual densities for RM1, RGDP, MMR, and RER, lag length=2

The results indicate that the second period lag is significant in at least two of the equations in the model. The null hypothesis of no serial correlation is rejected in the univariate case for the real money supply (cf. the $F_{ar}$ statistics against fourth-order autoregression) and for the system as a whole. There is also some evidence that the residuals from the real output, interest rate and exchange rate equations are non-normally distributed. The impact of the outlier observations is
seen more clearly in Figure 5.1. The residual plots indicate that the outlier problem is largely associated with the third quarter of 1997 and the first and second quarters of 1998.

Figure 5.2: Residuals (scaled) and residuals densities for RM1, RGDP, MMR, and RER, lag length=2, including *Dum97:3, Dum98:1, Dum98:2*

Increasing the lag length to $k = 3$ produces the results in Table 5.6; the additional lags are only significant in the real money and interest rate equations, serial correlation is no longer a problem, although the test for ARCH is significant for the exchange rate equation. Non-normal residuals are now only a problem in the equation determining *MMR*. However, the null hypothesis of normality in the system as a whole is rejected. In contrast, keeping $k = 2$ and including three dummy variables (labelled *Dum97:3, Dum98:1, and Dum98:2*, taking the value of 1 for each of these quarters) improves the stochastic properties of the model without the risk of overparametrising. The results are shown in Table 5.7. Although
the residuals from the interest rate equation are still non-normally distributed, normality is no longer rejected in the other equations and in the system as a whole. This is seen more clearly in Figure 5.2 which gives the plots of the residuals and residual densities when the three dummy variables are included. This points to a potential trade-off in terms of setting the value of $k$ and introducing $I(0)$ variables into $D_t$.

![Figure 5.3: Plots of the vectors](image)

Alternatively, if real output, interest rate, and exchange rate prove to be weakly exogenous, then non-normality is less of a problem since one can condition on the weakly exogenous variables and improve the stochastic properties of the model.\(^3\) It would therefore be appropriate to condition on these three variables instead of adding extra dummy variables.

In what follows, the Johansen reduced rank approach is applied to the model

\(^3\) The test for weak exogeneity of the above variables is conducted at a later stage.
Figure 5.4: Plots of the vectors, corrected for short-run dynamics

Table 5.8: Test of the cointegration rank of the Indonesian money demand model, unrestricted intercepts and no trends, 1989:1-2001:4

<table>
<thead>
<tr>
<th>$H_0: r$</th>
<th>$\lambda_i$</th>
<th>$\lambda_{max}$</th>
<th>adj. c.v.</th>
<th>adj.</th>
<th>$\lambda_{trace}$</th>
<th>adj. c.v.</th>
<th>adj.</th>
</tr>
</thead>
<tbody>
<tr>
<td>0</td>
<td>0.466</td>
<td>32.67*</td>
<td>27.65*</td>
<td>27.42</td>
<td>32.35</td>
<td>71.76*</td>
<td>60.72**</td>
</tr>
<tr>
<td>1</td>
<td>0.412</td>
<td>27.67*</td>
<td>23.42*</td>
<td>21.12</td>
<td>24.92</td>
<td>39.08*</td>
<td>33.07*</td>
</tr>
<tr>
<td>2</td>
<td>0.149</td>
<td>8.41</td>
<td>7.12</td>
<td>14.88</td>
<td>17.55</td>
<td>11.41</td>
<td>9.66</td>
</tr>
<tr>
<td>3</td>
<td>0.059</td>
<td>3.00</td>
<td>2.54</td>
<td>8.07</td>
<td>9.52</td>
<td>3.00</td>
<td>2.54</td>
</tr>
</tbody>
</table>

Notes: * and ** denote rejection of the null hypothesis at 1% and 5% significance level respectively. Critical values are from Table 6c in Pesaran et al. (2000).
Table 5.9: Determining cointegration rank and the model for the deterministic components using the trace test, Indonesian money demand data, 1989:1-2001:4

Notes: * Cannot reject the null hypothesis at the 5% significance level. Critical values are from Pesaran et al. (2000).

<table>
<thead>
<tr>
<th>$H_0: \tau$</th>
<th>rest. int.-no trends</th>
<th>unrest. int.-no trends</th>
<th>unrest. int.-rest. trends</th>
</tr>
</thead>
<tbody>
<tr>
<td>0</td>
<td>114.26</td>
<td>71.76</td>
<td>105.11</td>
</tr>
<tr>
<td>1</td>
<td>65.20</td>
<td>39.08</td>
<td>54.63</td>
</tr>
<tr>
<td>2</td>
<td>34.42</td>
<td>11.41*</td>
<td>25.61</td>
</tr>
<tr>
<td>3</td>
<td>7.47</td>
<td>3.00</td>
<td>7.88</td>
</tr>
</tbody>
</table>

Table 5.10: Normalised cointegration coefficients of the Indonesian money demand model, 1989:1-2001:4

Notes: Model with unrestricted intercepts and no trends. Based on 1 cointegration vector. t-values in parenthesis.
\[ \chi^2(3) = 7.40, \text{ p-value} = 0.059 \]

<table>
<thead>
<tr>
<th>RM1</th>
<th>RGDP</th>
<th>MMR</th>
<th>RER</th>
</tr>
</thead>
<tbody>
<tr>
<td>( \beta )</td>
<td>-1.000</td>
<td>0.793</td>
<td>-0.386</td>
</tr>
<tr>
<td>( \alpha )</td>
<td>0.524</td>
<td>0.000</td>
<td>0.000</td>
</tr>
<tr>
<td>(5.92)</td>
<td>(0.00)</td>
<td>(0.00)</td>
<td>(0.00)</td>
</tr>
</tbody>
</table>

Table 5.11: Test for weak exogeneity of RGDP, MMR, and RER

Notes: Based on LR test of restrictions on \( \alpha \). t-values in parenthesis.

discussed above (with intercept, no dummy variables, and \( k = 2 \)). Firstly, in discussing the formulation of the dynamic model, the questions of whether an intercept and trend should enter the short- and/or long-run model needs to be considered. Following Johansen (1992c), the joint hypothesis of both the rank order and the deterministic components is tested based on the so-called Pantula principle.\(^4\) The results from estimating the various alternatives and then applying the Pantula principle are presented in Table 5.9. The first time the null is not rejected is indicated by the *. Thus, one can accept that there are at most two cointegrating vectors and there are deterministic trends in the levels of the data.

The results of the trace and maximum eigenvalue tests are detailed in Table 5.8.

However, Reimers (1992) \textit{inter alia} argues that in small samples the Johansen procedure over-rejects when the null is true. Reimers suggests taking account of the

\(^4\) That is, all three models are estimated (restricted intercepts-no trends, unrestricted intercepts-no trends, and unrestricted intercepts-restricted trends). The results are presented from the most restrictive alternative (\( r = 0 \) and restricted intercepts-no trends) through to the least restrictive alternative (\( r = 3 \) and unrestricted intercepts-restricted trends). The test procedure is then to move through from the least restrictive model and at each stage to compare the trace statistic to its critical value and only stop the first time the null hypothesis is not rejected.
number of parameters to be estimated in the model and making an adjustment for degrees of freedom to obtain satisfactory size properties in finite samples. In contrast, Cheung and Lai (1993) suggest making finite-sample corrections by adjusting the critical values and not the tests statistics. Table 5.8 shows that using adjusted critical values/test statistics produces results similar to those obtained when no adjustment is made.

These results can be seen more clearly in Figure 5.3 which plots the four relations (associated with the four rows in $\beta$). The first two vectors correspond to the most stationary relations in the model, but there is evidence that both relationships are upward trending. The other two vectors are clearly nonstationary. The plots in Figure 5.4 confirm these findings while presenting a different version of the same relations, with all the short-run dynamics removed.

Doornik and Hendry (2001) point out that it is still unclear which correction should be preferred. Juselius (1995) suggests instead looking at the eigenvalues (i.e., the roots) of the companion matrix. There are 8 roots of the companion matrix. The moduli of the three largest roots are 0.989, 0.896 and 0.896, respectively,

---

5 This is accomplished by multiplying the tests statistics by a factor $(T - nk)/T$ where $T$ is the number of observations, $n$ is the number of variables in the VAR system and $k$ is the lag length of the VAR.

6 In addition, Cheung and Lai (1993) find that the trace test is more robust to non-normality in the residuals than the maximum eigenvalue test. Skewness in the residuals is found to have a statistically significant effect on the test sizes of both the Johansen tests, but less so for the trace test. The latter also appears to be more robust to excess kurtosis (in the residuals) than the maximum eigenvalue test.

7 See Doornik and Hendry (2001) for further details.

8 An alternative to making finite-sample corrections would be to simulate the exact distribution based on the d.g.p. underlying the model being considered, using bootstrap methods, but this is not without problems. The results in Harris and Judge (1998) suggest that the bootstrap test statistic has poor size properties.
AR Model estimated

\[ A(L)(RM1)_t = B_1(L)RGDP_t + B_2(L)MMR_t + B_3(L)RER_t + u_t \]

\[ R^2 = 0.979355 \quad F(7, 44) = 298.2[0.0000] \quad \sigma = 0.0404724 \quad DW = 2.12 \]

\[ RSS = 0.0720726522 \text{ for 8 variables and 52 observations} \]

Information Criteria: SC = -5.97344; HQ = -6.15855; FPE = 0.00189002

Diagnostic tests

AR 1-4 \[ F(4, 40) = 1.2004 \quad [0.3257] \]
ARCH 4 \[ F(4, 36) = 0.39198 \quad [0.8130] \]
Normality \[ \chi^2(2) = 0.26560 \quad [0.8756] \]
\[ \chi^2 = F(14, 29) = 0.26983 \quad [0.9938] \]
RESET \[ F(1, 43) = 1.8239 \quad [0.1839] \]

Solved Static Long Run equation

\[ RM1 = 2.80129 + 0.794656RGDP - 0.462968MMR + 0.301354RER \]

SE \( (2.678) \quad (0.0917) \quad (0.2410) \quad (0.0587) \)

WALD test \( \chi^2(4) = 210.784 \quad [0.0000]^{**} \)

Table 5.12: Results of the conditional model, 1989:1-2001:4
## Test on the significance of each variable

<table>
<thead>
<tr>
<th>Variable</th>
<th>F(num, denom)</th>
<th>Value</th>
<th>Probability</th>
</tr>
</thead>
<tbody>
<tr>
<td>RM1</td>
<td>F(1,44)</td>
<td>49.919</td>
<td>[0.0000]**</td>
</tr>
<tr>
<td>Constant</td>
<td>F(1,44)</td>
<td>0.94228</td>
<td>[0.3370]</td>
</tr>
<tr>
<td>RGDP</td>
<td>F(2,44)</td>
<td>7.8034</td>
<td>[0.0013]**</td>
</tr>
<tr>
<td>MMR</td>
<td>F(2,44)</td>
<td>4.3591</td>
<td>[0.0187]*</td>
</tr>
<tr>
<td>RER</td>
<td>F(2,44)</td>
<td>20.116</td>
<td>[0.0000]**</td>
</tr>
</tbody>
</table>

## Test on the significance of each lag

| Lag 1  | F(4,44)       | 49.902    | [0.0000]**          |

### COMFAC WALD test statistic table

<table>
<thead>
<tr>
<th>Order</th>
<th>Cumulative tests</th>
<th>Incremental tests</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>$\chi^2(3) = 29.976$ [0.0000]**</td>
<td>$\chi^2(3) = 29.976$ [0.0000]**</td>
</tr>
</tbody>
</table>

Table 5.13: Tests of significance of each variable/lag and COMFAC tests

indicates that all roots are inside the unit circle with the three largest close to unity. This suggests that $(n-r) = 3$. Thus it is possible to accept that there is one cointegration vector (and deterministic trends in the levels of the data), which is consistent with our theoretical prior on the existence of a long-run money demand function. The estimates of $\alpha$ and $\beta$ with their $t$-statistics (assuming $r = 1$ and normalising on $RM_{1t}$) are presented in Table 5.10.

Since one cointegration relation is identified, the weak exogeneity test are eval-
uated under the assumption of rank one. The null hypothesis that $RGDP_t$, $MMR_t$ and $RER_t$ are weakly exogenous is tested, thus imposing three row restrictions on $\alpha$. The results are presented in Table 5.11. Weak exogeneity of all three variables is not rejected at the 5% significance level. This indicates that it is valid to condition on these variables and use a single equation estimator of the cointegration vector. Furthermore, comparison of the Johansen multivariate estimator of $\beta_{full} = [1.0, 0.793, -0.504, 0.29]$ with the single equation estimator of $\beta_{part} = [1.0, 0.794, -0.462, 0.301]$ reported in Table 5.12 shows that the two approaches are equivalent.

The Wald test decisively rejects the null hypothesis that all the long-run coefficients (except the intercept term) are zero. The income elasticity is close to unity and the interest and real exchange rates both have coefficients which are quite well determined, correctly signed and highly significant. The income and interest rate response lies in the range found in the inventory-theoretic model of Miller and Orr (1966, 1968), and therefore could represent a structural response. Formal checks on whether the model is data coherent are provided by testing for residual autocorrelation (fourth-order), ARCH (fourth order), normality and functional form misspecification. The results reveal no problems.

Next, we turn to the analysis of the lag structure. The results are shown in Table 5.13: every $F$-test is significant.\textsuperscript{9} The $F$-test on the significance of the first-order lag shows that this lag matters greatly. Finally, tests of common factors

\footnote{The $F$-tests are tests of the joint significance of the associated variables, with all their lags, and provide tests for deleting each variable completely from the model.}
(COMFAC) in the lag polynomials also reported in Table 5.13 reject the hypothesis of four common factors (see Doornik and Hendry (2001) for an interpretation). Thus, we tentatively conclude that this model provides a cointegrated and data congruent representation of Indonesian money demand.

The model is next used to test whether there has been a shift in the long-run relationship amongst the variables in the money demand function. In order to test this proposition, the methodology of Gregory and Hansen (1996) described in the previous section is followed. The results are shown in Table 5.14. For all the statistics the maximum lag length is chosen using the formula reported in Schwert (1989, p. 151): \( k_{\text{max}} = \text{int}(12(T/100)^{1/4}) \). The selection of lag is based on the significant \( t \) method but other procedures based on information criteria (e.g., AIC, SC, HQ) have given identical results.

The conventional ADF test rejects the null hypothesis of no cointegration at the 5% significance level both in the model with intercept and in the model with intercept and trend. The ADF* tests which allow for a level shift (the C and C/T tests) also reject the null hypothesis at the 10% significance level with structural breaks occurring in the last quarter of 1997 and in the second quarter of 1991, respectively. These results can be seen more clearly in Figure 5.5 which plots the ADF* for level shift (C), level shift with trend (C/T) and regime shift (C/S).

Turning now to the PP statistics, the \( Z_t^* \) test which allows for a level shift is the only test which rejects the null hypothesis. Again, the structural break is identified with the last quarter of 1997.
### Tests of Cointegration with Regime Shifts

<table>
<thead>
<tr>
<th>Test statistics</th>
<th>Break point</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>ADF</strong></td>
<td></td>
</tr>
<tr>
<td>C</td>
<td>-5.17**</td>
</tr>
<tr>
<td>C/T</td>
<td>-5.34**</td>
</tr>
<tr>
<td>C/S</td>
<td>-5.71</td>
</tr>
<tr>
<td><strong>Z_t</strong></td>
<td></td>
</tr>
<tr>
<td>C</td>
<td>-5.20**</td>
</tr>
<tr>
<td>C/T</td>
<td>-5.24</td>
</tr>
<tr>
<td>C/S</td>
<td>-5.68</td>
</tr>
<tr>
<td><strong>Z_α</strong></td>
<td></td>
</tr>
<tr>
<td>C</td>
<td>-34.47</td>
</tr>
<tr>
<td>C/T</td>
<td>-35.20</td>
</tr>
<tr>
<td>C/S</td>
<td>-38.93</td>
</tr>
<tr>
<td><strong>ADF</strong></td>
<td></td>
</tr>
<tr>
<td>C</td>
<td>-4.33*</td>
</tr>
<tr>
<td>C/T</td>
<td>-4.26*</td>
</tr>
</tbody>
</table>

Table 5.14: Tests of cointegration with regime shifts

Note: * and ** denotes rejection of the null hypothesis of no cointegration at the 5% and 10% significance level respectively. Critical values for ADF*, Z_t*, and Z_α* are from Table 1 in Gregory and Hansen (1996).
5.4 Conclusions

Using a residual-based cointegration test with regime shift we find supporting evidence of a level shift (with trend) occurring in the second quarter of 1991 and coinciding with the Sumarlin shock. We also find evidence of a level shift in the money demand function occurring in the last quarter of 1997. This result corroborates the finding of the previous chapter: the timing of this break both in the univariate and multivariate contexts can be linked to the crisis and a government intervention in the money market for which the Sumarlin shock of 1991 is the precedent. These results suggest that the decision by the Indonesian monetary authorities in the aftermath of the crisis to move away from monetary growth targeting towards inflation targeting may be well-placed. In effect, one of the main advantages of inflation targeting over monetary targeting is that velocity shocks are irrelevant because the monetary policy strategy no longer relies on a stable money demand function (Mishkin, 1999).
Furthermore, the examination of the exogeneity property of the conditioning variables in the long-run money demand function has revealed that both the interest and real exchange rates variables are weakly exogenous. The result for the exchange rate is not surprising because for the sample period up to the third quarter of 1997 it has been more or less fixed (i.e. allowed to fluctuate within an intervention band widened several times). Hence, the exchange rate was not determined endogenously within the system. The fact that the interest rate is weakly exogenous indicates that it might not be possible to recoup any type of interest rate policy such as a Taylor rule.

The above findings lend support to the conventional Mundell-Fleming model which postulates that the trinity of open capital markets, pegged exchange rate regime and monetary policy autonomy is inconsistent. Since monetary policy is primarily directed to maintaining the peg between the Rupiah and the US dollar, monetary policy effects are potentially undermined and immobilised.
Appendix: Data definitions and sources

For each series the sample period is 1983:1-2001:4. The series are all seasonally adjusted (with the exception of the interest rate and real exchange rate series).\(^\text{10}\) The source is the IMFs International Financial Statistics (except where a different source is mentioned).

(1) \(RM1\), and \(RM2\) are the natural logarithms of the M1 and M2 definitions of the money stock (in billion rupiahs), each deflated by the consumer price index (1995=100). Source: Bank Indonesia (unpublished).

(2) \(RGDP\) is the natural logarithm of the Indonesian gross domestic product at 1995 market prices (in billion rupiahs). Source: Bank Indonesia (unpublished).

(3) \(MMR\) is the nominal money market rate (in level).

(4) \(INF\) is the inflation rate, computed as the change in the consumer price index (1995=100).

(5) \(USTBR\) is the nominal 3-months US T-bill rate (in level).

(6) \(RER\) is the Indonesian exchange rate in real terms, computed as:

\[
[l(WPI_t) - l(WPI^*_t) + l(E_t)]
\]

where \(E_t\) is the nominal Rupiah exchange rate defined as the foreign price of domestic currency (indirect quotation, a rise represents an Indonesian currency appreciation). Source: IMFs International Financial Statistics. \(WPI_t\) is the domestic wholesale price index (1995=100). \(WPI^*_t\) is a weighted average of foreign

\(^{10}\) See chapter 4, Appendix A.
wholesale price indices (1995=100) computed as:

\[ WPPI_t^* = \sum_{j=1}^{m_r} w_j^* WPPI_{jt}, \]

where \( w_j^* \) are fixed weights and \( m_r = 4.\)

\[ \text{See chapter 4, Appendix A.} \]
CHAPTER 6

ARDL APPROACH TO MODELLING INDONESIAN MONEY DEMAND
WITH FINANCIAL LIBERALISATION

6.1 Introduction

Traditional specifications of money demand applied to countries under financial liberalisation often yield parameter estimates that are not economically plausible, are subject to highly autocorrelated errors (Courakis, 1978), and frequently result in persistent overprediction of money demand, the so-called missing money episodes (Goldfeld, 1973; 1976). Inspection of the data reveals shifts, or continuing movements in holdings of money balances that are unrelated to the behaviour of the explanatory variables that appear most commonly in the literature. These shifts can be attributed, at least in part, to financial market developments. These developments cannot be considered as part of a stationary error term since their effect on money demand is permanent (Arrau et al., 1995). It then follows that, for countries under financial liberalisation, money demand functions that do not explicitly control for this process are likely to be misspecified.

In principle, financial liberalisation can be captured in the money demand function in various ways: by including dummy variables (Friedman and Schwartz, 1982), a deterministic trend (Moore et al., 1990; Arrau et al., 1995; Dekle and Pradhan, 1999), institutionally-related variables (Bordo and Jonung, 1981; 1987;
1990; Siklos, 1993; Klovland, 1983; Akhtar, 1983)\(^1\), or by adjusting monetary indices (Binner et al., 2004).

In this chapter, the empirical implications of financial liberalisation are explored by introducing a deterministic trend as an additional explanatory variable in the money demand function for Indonesia. As already argued in chapters 4 and 5, standard cointegration procedures, which have been widely used to estimate the long-run demand for money, may not be appropriate when applied to countries under financial liberalisation. This phenomenon is widely associated in the empirical literature with structural change. Instead, the procedure carried out in this chapter—the ARDL modelling approach of Pesaran et al. (2001)—is based on asymptotic critical value bounds for the Wald or \(F\)-statistics. It follows that one can investigate the existence of the long-run money demand function for Indonesia, in the presence of financial liberalisation, without the need to know if the variables included in the money demand function are integrated, stationary, or mutually cointegrated.

A similar strategy is used to estimate import demand functions for Cyprus (Pattichis, 1999) and Korea (Mah, 2000). Ghatak and Siddiki (2001) also use this approach to estimate virtual exchange rates in India. Tang (2001; 2002) use the procedure to model inflation behaviour and the demand for M3 in Malaysia.\(^2\)

\(^1\) Examples of such institutionally-related variables are bank offices per head of population, the proportion of the labor force employed outside of agriculture, the ratio of currency to total money stock, the ratio of non-bank to bank financial assets, etc...

\(^2\) These authors have used earlier unpublished versions of Pesaran et al. (2001) entitled: "Bounds Testing Approaches to the Analysis of Level Relationships", DAE Working Paper Series No. 9622 (1996) and No. 9907 (1999), Cambridge University, Department of Applied Economics.
The rest of the chapter is organised as follows. Section 6.2 discusses the ARDL methodology. The description of the data and the empirical results are presented in Section 6.3. Section 6.4 offers some concluding remarks.

6.2 Methodology

The econometric specification used in this chapter is similar to the one used in the previous chapter. Real money demand is a function of a scale variable (e.g., the level of real income), and an opportunity cost variable (e.g., the rate of interest on an alternative asset). The availability of foreign currency assets for investment implies that the variety of assets available for portfolio diversification is wider (Chowdury, 1997). The direct currency substitution literature, which suggests portfolio shifts between domestic and foreign money, focuses on the exchange rate variable, whereas the capital mobility literature focuses on the foreign interest rate variable (Cuddington, 1983; Leventakis, 1993; Tan, 1997). We include a foreign interest rate variable to obtain the open-economy version of the money demand function.

As argued in Lieberman (1977, 1979), measuring the impact of technological change on money demand is a difficult exercise. Ideally, one would like to model explicitly each source of improvement in cash management. Unfortunately, innovations in cash management include many new activities, instruments, institutions, and practices for which data are often sparse or non-existent. Even if it were feasible, modelling each source of innovation would quickly exhaust the degree of freedom available. Conversely, omitting any measure of technological change is
hardly costless. If technological change is ignored, the coefficients estimates will be biased. As a result, even crude proxies for the omitted variable, technological change, should be considered.

A crude proxy for technological change which is commonly used elsewhere in the literature is a deterministic trend. The inclusion of a trend in the estimation merely assumes that the implementation of new technologies or practices reduces money demand smoothly over time. Although the introduction of new technologies or cash management methods may be discrete or lumpy rather than continuous, the implementation of such methods throughout the economy will be distributed with a lag, thereby smoothing out the impact on money balances. A deterministic trend would measure the mean rate at which new cash management techniques have an effect on money balances, assuming technological change is exogenous to the financial sector (Lieberman, 1979).

Since money demand is not directly observable, it is assumed that the money market is always in equilibrium and that the money supply is exogenous. Thus, one can identify the money demand function described above and use the quantity of money to measure the demand for money.

Following the modelling approach developed in Pesaran, Shin and Smith (2001) (PSS hereafter) the section starts from the maintained assumption that the time series properties of the variables included in the money demand function can be well approximated by a log-linear VAR(p) model. Let:
\[ z_t = [m_t, y_t, R_t, R_t^*]' = [m_t, x_t]' \] (6.1)

where \( m_t \) is real money balances (in logarithms), \( y_t \) is real income (in logarithms), \( R_t \) is the nominal domestic interest rate (in levels), and \( R_t^* \) is the foreign interest rate (in levels).

The conditional (partial) model can be written as:

\[ \Delta m_t = c_0 + c_1 t + \pi_{mm} m_{t-1} + \pi_{mx.t} x_{t-1} + \sum_{i=1}^{p-1} \psi_i \Delta z_{t-i} + \omega' \Delta x_t + u_t \] (6.2)

where \( c_0 \) is an intercept, \( t \) is a deterministic trend, and \( u_t \sim \text{IID}(0, \sigma^2) \).

Under the assumption that lagged real money demand, \( m_{t-1} \), does not enter the sub-VAR model for \( x_t \), for equation (6.2) the above real money demand function is identified and estimated consistently by the OLS. PSS developed bounds tests to test the existence of a long-run relationship between the levels of \( m_t \) and \( x_t \), \( t = 1, 2, \ldots \). More specifically, the approach consists in testing for the absence of any long-run relationship between \( m_t \) and \( x_t \), \( t = 1, 2, \ldots \); that is, the exclusion of the lagged level variables \( m_{t-1} \) and \( x_{t-1} \) in (6.2). Hence, the null hypotheses are given by:

\[ H_{0}^{\pi_{mm}}: \pi_{mm} = 0, \quad H_{0}^{\pi_{mx.t}}: \pi_{mx.t} = 0' \]

and the alternative hypotheses by:

\[ H_{1}^{\pi_{mm}}: \pi_{mm} \neq 0, \quad H_{1}^{\pi_{mx.t}}: \pi_{mx.t} \neq 0'. \]
Since the asymptotic distributions of the $F$- and $t$- statistics are non-standard under the null hypothesis, irrespective of whether the forcing variables $\{x_t\}$ are $I(0)$ or $I(1)$, PSS provide two sets of critical values for these statistics covering various specifications of the deterministic terms; one set assuming that the forcing variables $\{x_t\}$ are $I(0)$ and the other assuming that $\{x_t\}$ are $I(1)$. These two sets provide lower and upper critical value bounds covering all possible classifications of $\{x_t\}$ into $I(0)$, $I(1)$ and mutually cointegrated processes. If the computed Wald or $F$- statistics fall outside the critical value bounds, a conclusive decision results without needing to know the cointegration rank $r$ of the $\{x_t\}$ process. If, however, the Wald or $F$- statistics fall within these bounds, inference would be inconclusive, and knowledge of the cointegration rank $r$ of the forcing variables $\{x_t\}$ is necessary to proceed further.

6.3 Empirical results

This study uses quarterly data from 1983:1 to 2000:4. The logarithm of the monetary aggregate M2 (broad money) is used to proxy $m$. In the Indonesian context, Dekle and Pradhan (1999) have argued that the behaviour of broad money velocity seems to be better captured by a deterministic trend (the chosen proxy for financial liberalisation) than the behaviour of narrow money velocity. Since the model requires money demand in real terms, the consumer price index (with the base 1993) is used as a deflator. The logarithm of the real gross domestic product is used as a proxy for $y$. The domestic interest rate, $R$, is proxied by the money market rate (a typical value is 0.10). $R^*$, the foreign interest rate, is proxied by
the 3-months US T-Bill rate (a typical value is 0.10).\(^3\)

Two impulse dummy variables are entered in the short-run part of the model to dummy out two outliers in the data. These dummy variables are defined by

\[ D_{\text{um90:4}} = 1 \text{ for the fourth quarter of 1990 and zero elsewhere}, \]

\[ D_{\text{um98:2}} = 1 \text{ for the second quarter of 1998 and zero elsewhere}. \]

It is worth mentioning that the asymptotic theory on which the bounds test is developed is not affected by the inclusion of such one-off dummy variables. The results of the estimation of (6.2) without including the two dummy variables are reported in Table 6.4. The estimated coefficients are very close to those obtained when the dummy variables are included in the model. However, the Bera-Jarque test suggests a problem of non-normality, giving further support to the inclusion of the dummy variables to dummy out the outliers.

To determine the appropriate lag length, \(p\), and whether it is important to add the proxy for financial liberalisation, the deterministic trend, to the income and interest rate variables, model (6.2) was estimated by the OLS with and without a deterministic trend, for \(p = 1,2\). Unfortunately, the size of the sample does not allow to estimate (6.2) for higher lag lengths as the degree of freedom needs to be saved. First-lagged changes of the income and domestic interest rate variables, \(\Delta y_{t-1}\) and \(\Delta R_{t-1}\), were found not significant (either singly or jointly) when a deterministic linear trend is included in the regression. First-lagged change of

\(^3\) Data on nominal M2 are from Bank Indonesia (unpublished). Data on real gross domestic product (at constant prices 1993) are from Biro Pusat Statistik (BPS, Central Bureau of Statistics, unpublished). Data on the consumer price index, the money market rate and the US 3-months treasury bill rate are from the IMF's International Financial Statistics (IFS).
the income variable and the first-difference of the foreign interest rate variable, \( \Delta y_{t-1} \) and \( \Delta R^*_t \), were also found not significant (either singly or jointly) when the regression did not contain a deterministic trend. Therefore, in both cases, for \( p = 2 \), parsimonious versions of (6.2) were re-estimated to avoid overparameterisation.

Table 6.1 gives the Akaike's Information, Schwarz's Bayesian Criteria and the Lagrange multiplier (LM) statistics for testing the hypothesis of serial correlation of order 1 and 4. These are denoted by (1) and (4), respectively. The AIC suggests the lag order, \( p \), to be 1 when the model contains a trend and 2 when it does not. The SC criterion, on the contrary, suggests \( p \) to be only 1, irrespective of whether a deterministic trend is included in the model or not. The \( \chi^2_{sc}(4) \) statistic suggests to select \( p \) to be either 1 or 2 if the model includes a deterministic trend, but only 2 if the model does not.

Table 6.2 gives the values of the \( F \)- and \( t \)-statistics for testing the existence of a long-run money demand function under three different cases depending on whether the model contains a linear trend and whether the trend coefficients are restricted. The \( F \)- and \( t \)-statistics in Table 6.2 are compared with the critical value bounds provided in Tables C1 and C2 in PSS.

First, consider the bounds \( F \)-test. For the model with a deterministic trend, \( F_v \) is the standard \( F \)-statistic for testing the restrictions \( \pi_{mm} = 0 \) and \( \pi_{mx.x} = 0 \), while \( F_{IV} \) is the standard \( F \)-statistic for testing \( \pi_{mm} = 0, \pi_{mx.x} = 0, \) and \( c_1 = 0 \) in (6.2). As argued in Pesaran, Shin and Smith (2000), the statistic \( F_{IV} \), which sets the trend coefficient to zero under the null of no long-run level relationship,
is more appropriate than $F_V$, which does not impose this restriction. The critical value bounds for the statistics $F_{IV}$ and $F_V$ are given in Tables C1.iv and C1.v, respectively, in PSS. Since the model contains three regressors, the 95% critical value bounds are $(4.23, 5.29)$ and $(4.01, 5.07)$ for $F_{IV}$ and $F_V$, respectively.

When a linear trend is included in the model, the test outcome critically depends on the choice of the lag order, $p$. For $p = 1$ (selected by the AIC), the hypothesis that there exists no long-run money demand function is rejected at the 95% level, irrespective of whether the regressors are $I(0)$ or $I(1)$. However, for $p = 2$, the statistics $F_{IV}$ and $F_V$ are inconclusive. When the bounds $F$-test is applied without a linear trend, the 95% critical value bounds of the relevant test statistic, the $F_{III}$, are $(3.23, 4.35)$. Irrespective of the lag order, the $F_{III}$ statistic rejects the null hypothesis of no long-run money demand function.

Table 6.2 also reports the results of the bounds $t$-test to the money demand functions. The two $t$-statistics, $t_V$ and $t_{III}$, are the $t$-ratios of the OLS estimates of $\pi_{mm}$ in (6.2), with and without the deterministic trend, respectively. The critical value bounds for these statistics are given in Tables C2.iii and C2.v in PSS. These statistics show that the null hypothesis is not rejected for $p = 2$, irrespective of whether the trend is included or not. However, for $p = 1$, the null hypothesis is rejected with or without the trend. The test results therefore support the existence of a long-run relationship when $p = 1$ and when the deterministic trend is included in the model. This specification is in accord with the evidence of the performance of the alternative models set out in Table 6.1.4 These results

---

4 It is likely that the results of the $F_{III}$ and $t_{III}$ statistics reported in Table 6.3 for $p = 1$
With Deterministic Trends | Without Deterministic Trends
---|---
P | AIC | SC | $\chi^2_{SC}$ (1) | $\chi^2_{SC}$ (4) | AIC | SC | $\chi^2_{SC}$ (1) | $\chi^2_{SC}$ (4)
1 | 116.04 | 105.86 | 4.59 | 6.41$^\dagger$ | 114.70 | 105.65 | 7.88 | 10.55
2 | 115.70 | 103.34 | 0.55$^\dagger$ | 1.30$^\dagger$ | 115.86 | 104.61 | 2.66$^{\dagger\dagger}$ | 4.24$^\dagger$

Table 6.1: Statistics for selecting the lag order of the Indonesian money demand equation

Notes: $p$ is the lag order of the conditional model (6.2), with zero restrictions on the coefficients of lagged changes in the $y$ and $R$ variables with deterministic trends and $p = 2$, and zero restrictions on the coefficients of lagged changes in the $y$ and $R^*$ variables without deterministic trends and $p = 2$. $AIC_p = LL_p - sp$ and $SC_p = LL_p - ln(T)$, are the Akaike and Schwarz Information Criteria, where $LL_p$ is the maximised log-likelihood value of the model, $p$ is the lag order, $sp$ is the number of freely estimated coefficients, and $T$ is the sample size. (1) and (4) are the LM statistics for testing the absence of residual serial correlations of order 1 and 4, respectively. The symbols $\dagger$, and $\dagger\dagger$ represent significance at 5% or less, and 10% or less, respectively.
show that the decision to include a deterministic trend in the model is not trivial. Since the trend captures financial liberalisation, it implies in turn that a long-run money demand relationship can only be obtained when the chosen proxy for financial liberalisation is included.

The remainder of this chapter focuses on a money demand model which has a lag order \( p = 1 \) and includes a deterministic trend. This specification yields the following estimated long-run money demand function:

\[
m_t = 1.526 y_t - 0.160 R_t - 1.021 R_t^* + 0.008 t - 10.377 + \bar{v}_t
\] (6.3)

where \( \bar{v}_t \) is the equilibrium correction term.

As argued in Dekle and Pradhan (1997), financial liberalisation may change the velocity of money, in principle, in either direction. Reforms that increase the number of banks, and spur institutional and technological advances such as credit cards, and electronic transfers of deposits or cash machines, can raise the velocity of broad and narrow money, as these developments make it easier to convert money into money substitutes. However, recall that Bordo and Jonung (1990) hypothesised that in many developing countries the velocity of broad money may decline over time because of the increasing monetisation of the economy or financial deepening.\(^5\) Furthermore, there can be shifts between the various categories of money.

---

\(^5\) Bordo and Jonung suggest that the institutional and financial factors that systematically influence the demand for money in an economy over the entire course of its development are of two types. First, the process of monetisation, defined as the growth of the commercial banking
### Table 6.2: F- and t-statistics for testing the existence of a long-run money demand equation for Indonesia

<table>
<thead>
<tr>
<th></th>
<th>With Deterministic Trends</th>
<th>Without Deterministic Trends</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$F_{IV}$</td>
<td>$F_V$</td>
</tr>
<tr>
<td>1</td>
<td>5.91$^c$</td>
<td>6.59$^c$</td>
</tr>
<tr>
<td>2</td>
<td>4.12$^b$</td>
<td>4.25$^b$</td>
</tr>
</tbody>
</table>

Notes: $p$ is the lag order of the underlying model. See also the notes to Table 6.1. $F_{IV}$ is the F-statistic for testing zero restrictions on the coefficients of the lagged level variables and the trend term in (6.2). $F_V$ is the F-statistic for testing zero restrictions on the coefficients of the lagged level variables in (6.2). $F_{III}$ is the F-statistic for testing zero restrictions on the coefficients of the lagged level variables in (6.2) without the trend term. $t_V$ and $t_{III}$ are the t-ratios of the coefficients of $m_{t-1}$ in (6.2) with and without a deterministic trend. $a$ denotes that the statistic lies below the 95% lower bound, $b$ denotes that it falls within the 95% bounds, and $c$ denotes that it falls outside the 95% upper bound.

As interest rates are liberalised on time deposits, private agents may shift their assets from currency and demand deposits to time deposits, raising the velocity of narrow money, but lowering the velocity of broad money. In the case of Indonesia, there has been a marked secular decline in $M_2$ velocity in the past two decades.

The coefficient estimates of the long-run interest rate semi-elasticities are reasonable and correctly signed. Since, as mentioned above, financial liberalisation can in principle change the demand (velocity) for money in either direction, the

---

_system in addition to the expansion of formal markets at the expense of barter and production for own use, can increase the demand for money as an economy grows. On the other hand, the emergence of a variety of nonbank financial intermediaries offering assets that potentially substitute for money and the invention of cash management techniques used to economise on real balances can have the opposite effect of lowering money demand. Bordo and Jonung’s hypothesis is that monetisation dominates early in the course of economic development while financial innovation dominates the later stages of growth. Hence, velocity will tend to trace out a U-shaped pattern over time._
coefficient of the deterministic trend can have either a positive or negative sign. In effect, the coefficient of the trend indicates a positive effect of financial liberalisation on broad money demand of about 3.2 percent per year. In comparison, Deckle and Pradhan (1999), using other measures of domestic and foreign interest rates and annual data over 1974-1995, report a positive effect of 7.4 percent per year.

The unrestricted error-correction regression associated with the above (level) long-run relationship is reported in Table 6.3. All contemporaneous changes in the income, and domestic and foreign interest rate variables are correctly signed and statistically significant, further justifying the choice of \( p = 1 \). The equilibrium correction coefficient is estimated to be \(-0.501\), which is reasonably large and suggests that the money demand process converges fairly rapidly towards its equilibrium given by (6.3).

The regression fits reasonably well, and the plot of actual and fitted values (see Figure 6.1) shows that the estimated model follows closely the path of the actual variables and captures almost all the turning points. The model also passes a number of diagnostic tests. The Bera-Jarque test for normality indicates that the normality assumption appears to be consistent with the data. The Koenker-Bassett test shows no evidence of heteroskedasticity. The Breusch-Godfrey Lagrange multiplier (LM) test for serial correlation indicates no evidence of serial correlation up to the fourth order. The model should not be given extra-credit for this result since it was chosen specifically to meet this test.\(^6\) However, the regression fails Ramsey's

\(^6\) The difference between the \( \chi^2 \) reported in Table 6.2 and the one reported in Table 6.1
<table>
<thead>
<tr>
<th>Regressor</th>
<th>Coefficient</th>
<th>Standard Error</th>
<th>p-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>$m_{t-1}$</td>
<td>-0.5013</td>
<td>N/A</td>
<td>N/A</td>
</tr>
<tr>
<td>$y_{t-1}$</td>
<td>0.7651</td>
<td>N/A</td>
<td>N/A</td>
</tr>
<tr>
<td>$R_{t-1}$</td>
<td>-0.0807</td>
<td>N/A</td>
<td>N/A</td>
</tr>
<tr>
<td>$R_{t-1}^*$</td>
<td>-0.5121</td>
<td>N/A</td>
<td>N/A</td>
</tr>
<tr>
<td>$\Delta y_t$</td>
<td>0.6835</td>
<td>0.1474</td>
<td>0.000</td>
</tr>
<tr>
<td>$\Delta R_t$</td>
<td>-0.1973</td>
<td>0.0744</td>
<td>0.010</td>
</tr>
<tr>
<td>$\Delta R_t^*$</td>
<td>-3.1053</td>
<td>1.0183</td>
<td>0.003</td>
</tr>
<tr>
<td>Intercept</td>
<td>-5.2029</td>
<td>0.9181</td>
<td>0.000</td>
</tr>
<tr>
<td>$t$</td>
<td>0.0044</td>
<td>0.0014</td>
<td>0.003</td>
</tr>
<tr>
<td>Dum90:4</td>
<td>0.1527</td>
<td>0.0378</td>
<td>0.000</td>
</tr>
<tr>
<td>Dum98:2</td>
<td>0.1923</td>
<td>0.0418</td>
<td>0.000</td>
</tr>
</tbody>
</table>

Adjusted-$R^2 = 0.5435$, Standard Error of Regression = 0.0356, $AIC = 130.91$, $SC = 118.46$, $\chi^2SC(4) = 1.95[0.743]$, $\chi^2FF(1) = 9.04[0.003]$, $\chi^2N(2) = 0.32[0.851]$, $\chi^2H(1) = 0.18[0.668]$

Table 6.3: Unrestricted error correction model of the Indonesian money demand equation

Notes: $\chi^2SC(4)$, $\chi^2FF(1)$, $\chi^2N(2)$, $\chi^2H(1)$ are the Chi-squared statistics for tests of residual serial correlation, functional form misspecification, non-normal errors and heteroskedasticity, respectively.
RESET general test of misspecification at the 5% level. It may be the consequence of the simple dynamic structure retained to save the degree of freedom. It could also reflect some non-linear effects or asymmetries in the adjustment of real broad money demand that the linear specification cannot take into account.

The CUSUM and CUSUMSQ tests, based on recursive estimation of the unrestricted error correction model, suggest that the regression coefficients are statistically stable over the sample period. The plots of the cumulative sum and cumulative sum of squares based on the recursive residuals are shown in Figures 6.2 and 6.3. They do not show evidence of statistically significant structural breaks.

![Figure 6.1: Plot of actual and fitted values](image)

is explained by the inclusion of the two impulse dummy variables in the regression.

These tests do not require the possible break point in the data to be known a priori. The CUSUM test detects systematic changes in the regression coefficients while the CUSUMSQ test is used to capture sudden departure from the constancy of regression coefficients. Note that the two impulse dummies are excluded from the model to allow the tests to be performed over the whole sample period.
<table>
<thead>
<tr>
<th>Regressor</th>
<th>Coefficient</th>
<th>Standard Error</th>
<th>p-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>$m_{t-1}$</td>
<td>-.4336</td>
<td>N/A</td>
<td>N/A</td>
</tr>
<tr>
<td>$y_{t-1}$</td>
<td>.6836</td>
<td>N/A</td>
<td>N/A</td>
</tr>
<tr>
<td>$R_{t-1}$</td>
<td>-.0330</td>
<td>N/A</td>
<td>N/A</td>
</tr>
<tr>
<td>$R_{t-1}^*$</td>
<td>-.2782</td>
<td>N/A</td>
<td>N/A</td>
</tr>
<tr>
<td>$\Delta y_t$</td>
<td>.4374</td>
<td>.1764</td>
<td>.016</td>
</tr>
<tr>
<td>$\Delta R_t$</td>
<td>-.1194</td>
<td>.0893</td>
<td>.186</td>
</tr>
<tr>
<td>$\Delta R_t^*$</td>
<td>-2.9904</td>
<td>1.2701</td>
<td>.022</td>
</tr>
<tr>
<td>Intercept</td>
<td>-4.7373</td>
<td>1.1376</td>
<td>.000</td>
</tr>
<tr>
<td>$t$</td>
<td>.0035</td>
<td>.0017</td>
<td>.044</td>
</tr>
</tbody>
</table>

Table 6.4: Unrestricted error correction model without impulse dummy variables

Adjusted-$R^2$ = .2896, Standard Error of Regression = .0444, $AIC = 116.04$, $SC = 105.86$, $\chi^2 SC(4) = 6.41[.170]$, $\chi^2 FF(1) = 6.91[.009]$, $\chi^2 N(2) = 12.79[.002]$, $\chi^2 H(1) = 0.02[.869]$
6.4 Conclusions

Previous research on financial liberalisation and money demand in Indonesia is fairly limited. Using cointegration and estimating money demand (jointly with a purchasing power parity equation) in the ASEAN-4 (Indonesia, Malaysia, Singapore and Thailand), Dekle and Pradhan (1999) found that the money function for Indonesia did not cointegrate when deterministic trend and step dummy variables—controlling for financial liberalisation—were excluded. The deterministic trend was found statistically significant but all three step dummy variables
were found insignificant.

Our findings support the inclusion in the money demand function for Indonesia of the deterministic trend as a crude proxy for financial liberalisation. In particular, we show that the existence of the long-run money demand function can only be firmly established when the trend is included in the model. Moreover, it is also found to be statistically significant. This result is obtained, however, without the need to include any exogenous step dummy variables. The model, however, does not pass the misspecification test. This may be the consequence of the simple dynamic structure retained to save the degree of freedom. It could also reflect some non-linear effects or asymmetries in the adjustment of real broad money demand that our linear specification cannot take into account.

From a policy perspective, the finding that Indonesian broad money was extremely sensitive to changes in the foreign (US) interest rate over the sample period suggests that—while M2 appeared to be a reliable indicator of liquidity conditions—controlling its growth was particularly difficult. Overall, the results suggest that financial liberalisation does play an important role in determining money demand, and its fluctuations, in Indonesia. This study provides a robust empirical basis for further research on the impact of financial liberalisation on money demand in countries which have experienced this phenomenon.
CHAPTER 7

CONCLUSION

In this thesis, we have shown, through a battery of statistical tests, both in univariate and multivariate contexts, that in effect, something structural seems to have happened to money demand in Indonesia. There is supporting evidence of a break in the money demand function occurring in the second quarter of 1991 and coinciding with a major government intervention in the money market known as the Sumarlin shock. We have also found evidence of a break, both in the univariate series and the multivariate money demand relation occurring in the last quarter of 1997 coinciding with a deep recession and a government intervention in the money market much similar to the earlier Sumarlin Shock of 1991.

The issue of distinguishing between stochastic processes with long memory and those with short memory or stationarity that exhibit structural changes (chapter 4) is relevant because the current shocks will have temporary effects of greater (integrated process) or lower duration (stationary process), and only occasional events or shocks associated to structural breaks will have permanent effects on the long-run level of the series. In effect, when stationary series with structural breaks are observed, these few occasional shocks with long-duration effects would produce certain persistence or symptoms of nonstationarity in the series. Since our results were obtained under the maintained assumption that there was at most one structural break, future research could give more attention to the issue of multiple
structural breaks.

When estimating breaks under common stochastic trends (chapter 5), it is important to signal that conventional cointegration tests could yield misleading results in the presence of structural breaks (Leybourne and Newbold, 2003). In addition, the study of cointegration relations in the presence of structural breaks is very complicated and even more, in case of more than two series, since the instability or break could be in all or several cointegration coefficients, in the error correction coefficient, in all long-run relations or in some of them, or in the short run dynamics, and the breaks could happen contemporaneously in all the series or not (Jimeno et al., 2006). Since our results were obtained under the maintained hypothesis of cointegration they should be taken with some caution. Instability in the number of cointegration relations or cointegration rank (probably due to the presence of structural breaks in the cointegration relation) may have occurred. Future research could apply the approaches of Hansen and Johansen (1999), Inoue (1999), and Johansen et al. (2000) to test the cointegration rank allowing for structural breaks.

Finally, we have shown how modelling financial liberalisation in the money demand function as a deterministic drift process can result in an improvement over the standard specifications in terms of yielding more constant and plausible values for the parameters of the money demand function (chapter 6). In particular, our findings support the existence of a stable long-run money demand function for Indonesia which, unlike previous studies, does not rely on the inclusion of
statistically insignificant exogenous step dummy variables. This result is based on a novel but robust econometric procedure which avoids pretesting of the order of integration of the variables in the money demand function.

From a policy perspective, the findings that M1 was sensitive to changes in the exchange rate, and that the demand for narrow money may have undergone a regime shift during the period under investigation suggest that a narrow money growth target was not suitable. Similarly, the finding that the demand for M2 was stable but extremely sensitive to the foreign (US) interest rate suggests that—while broad money appeared to be a reliable indicator of liquidity conditions throughout the sample period—controlling its growth was particularly difficult.

More generally, our results support the notion that financial liberalisation might have rendered the money supply endogenous, with respect to output. The Indonesian central bank would have had difficulty in controlling the money supply because even though it could manage to generate credit constraints in the Indonesian economy, Indonesian banks could easily borrow on international financial markets, which undermined the effectiveness of monetary policy.

The above findings support the Mundell-Fleming type argument that the trinity of open capital markets, pegged exchange rate regime and monetary policy is inconsistent. Before the 1997 crisis, during periods of upswing in the economy, rising aggregate demand was accompanied by both increased foreign borrowing and the liquidation of SBI by sale to the central bank, both of which resulted in increases in base money (given the quasi-fixed exchange rate regime). In such
context, controlling base money growth (and through it aggregate demand) was difficult and required extremely high interest rates and/or dramatic contractions of domestic liquidity of the Sumarlin shock type. Most significantly, monetary policy was immobilised in the period before the decision to float the Rupiah in 1997.

Overall, our findings should foster more research at identifying the underlying causes of money demand instability in Indonesia and other developing economies where similar programmes of financial liberalisation have been implemented.
REFERENCES


