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ABSTRACT

THE LONG-RUN DEMAND FOR MONEY FUNCTIONS IN TAIWAN (1961:4-1997:3): COINTEGRATION AND STABILITY

By

Ching-yi Chiang

This dissertation compares the performance of the long-run demand functions for alternative new definitions of money (M1A, M1B, M1BP, and M2) in Taiwan for the same sample period (1961:4-1997:3). Various unit-root tests cannot reject that the variables chosen for each function (real balances, real GNP, and short-term interest rates) are characterized as I(1) processes. Using residual-based cointegration tests (Phillips and Ouliaris 1990, Gregory and Hansen 1996), we found cointegration evidence in all four demand functions for real balances. However, the parameter stability tests of Hansen (1992b) suggest that the most stable long-run relationship occurs in the demand for real M1B equation. M1B being stable is consistent with the transactions-demand theory which links real money balances demanded with a measure of the volume of real transactions (real GNP) and opportunity costs of holding money balances. Instability detected in the demand for broadly defined money (M1BP and M2) is explained by misspecification. Low

liquidity in the non-M1B components suggests that the demand for broader measures may be justified by portfolio or speculative-demand theory; hence, wealth instead of real GNP, or a measure of volatility of alternative assets' returns enters the equation.

Johansen's (1991) trace and max-λ test establish that there is only one cointegrating vector in the system. For the demand for real M1B function, the equilibrium income elasticity is around 1.50, and interest elasticity is -0.41, which are robust to four alternative estimators. While the price homogeneity is not rejected, the unitary income elasticity is rejected by the data. The rejection of weak exogeneity status of the variables invalidates the inference conducted in single equation error correction model in estimating the long-run demand for real M1B function in Taiwan.

To my Mother, Kwey-jen Deng Chiang

and my Father, Ching-jen Chiang

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CHAPTER 1

INTRODUCTION

The aggregate demand for money function, which links money balances demanded with a scale variable and a measure of the opportunity cost of holding money balances, is an essential ingredient in macroeconomic analysis and its importance is associated with efficacy of monetary policy. As Keynes (1936) postulated long before, while a stable money demand relation under different state of nature does not guarantee the potence of the monetary policy, its instability certainly undermines the usefulness of manipulation of quantity of money in an economy as a central stabilizing device. Many macroeconomic models are built on a notion that the money demand relation remains stable to analyze the effect of monetary policy on economic activity, such as Monetarist (Friedman, 1956, 1974), New Classical model (Sargent and Wallace, 1975), New Keynesian model (Mankiw, 1991), and the empirical real business cycle model (King, Plosser, Stock, and Watson, 1991). Finding an empirical stable money demand relation provides these models an empirical content.

Numerous attempts to estimate money demand relations in Taiwan have been undertaken. Studies on Taiwanese money demand function concentrate on the Goldfeld-type (Goldfeld, 1973, 1976) partial adjustment model. Empirical evidence on the issue of stability based on Taiwanese data is mixed, depending on the functional form, variables used, sample period studied, and the choice of estimating and testing procedure, see Hwang and Wu (1994) and Lin (1997) for a review. Using Engle-Granger (1987) ADF cointegration tests, Chang (1989) finds that there is no evidence of cointegration among real balances, real GNP, one-month time deposit rate for two monetary aggregates (real M1A and real M1B) but cointegration occurs in real M2 equation (using the same scale variable, and interest rate variable). The evidence from the error-correction model shows that the short-run specification for the demand for real M2 remains stable over the sample period (1964:1-1987:4). Lee (1994)[1] examines the equilibrium relationship between real M1B, real GNP, and one-month time deposit rate over the sample period (1961:3-1992:4) using Johansen (1991) and Hansen and Johansen (1993) testing procedures. He finds that the demand for real M1B function shifted downward in the fourth quarter of 1982, while leaving the income and interest elasticity unchanged.

[1] Ph.D dissertation, C.N. Lee, 1994, Department of Economics , Michigan State University. Also see the published version of Lee's Ph.D dissertation (in Chinese) (1996).

Conventionally, studies on Taiwanese money demand use bank deposit rates as a measure of opportunity cost of holding money balances. However, the characteristics of the financial structure in Taiwan - flourishing unorganized money markets (U.M.M.) - have been overlooked in the Taiwanese money demand literature. In the study, we argue that the interest rate on U.M.M., which has been recorded since 1947 but omitted in most of the early studies, is a relevant variable entering the money demand relationships and this interest rate may also play an important role in the macroeconomic analysis in Taiwan.

The objective of this study is to investigate the stability of the long-run demand for money functions for four alternative new definitions of money by incorporating the characteristics of the financial structure in Taiwan. We estimate the equations using quarterly data from 1961:4 through 1997:3. The sample period covers several regime shifts (two oil crises 1973, 1979, shifts from a fixed exchange rate regime to a floating exchange rate regime 1978, and financial deregulation since 1980). Since we do not have theories suggesting the appropriate definition of money for monetary analysis and particular empirical measures as clearly superior to others which yield more stable money demand relationships, we estimate the money demand functions for four measures of money (M1A, M1B, M1BP, and M2) and then compare their performance over the same sample period using the same econometric techniques to reveal which measure may serve a better guide to monetary policy.

This dissertation is organized as follows. In Section I of CHAPTER 2, we briefly review the theories of the demand for money in the literature. In Section II, we derive a transactions demand for money model from a multiperiod utility maximization framework without imposing a cash-in-advance constraint. This model has solutions similar to those in Lucas (1988) and possesses some properties which help understand some empirical findings. In Section III the importance of a stable aggregate money demand function in macroeconomic analysis is discussed. In Section IV we clarify some questions regarding the appropriate interest rate variable in the Taiwanese money demand functions a priori and describe the data series used in the study.

In CHAPTER 3 the statistical properties of the data series are investigated in order to choose appropriate econometric techniques to study the money demand relationships. In Section I there is a brief discussion of how the time series analysis of models with unit roots has had a major impact on our understanding of the response of economic system to shocks. In Section II we apply various univarite unit root tests (Hylleberg, Engle, Granger, and

Yoo 1990, Fuller 1976, Dickey and Fuller 1979, 1981, Kwiatkowski, Phillips, Schmidt, and Shin 1992) incorporating seasonality in the data to establish the order of integration of the individual time series.

In CHAPTER 4 the empirical evidence of the long-run money demand relationships for four measures of money (M1A, M1B, M1BP, and M2) is explored and their stability during 1961:4-1997:3 is examined. In Section I equilibrium concepts in terms of cointegration are sketched. In Section II the econometric methods for analyzing relationship among I(1) variables are introduced and the empirical cointegration evidence in the four money demand functions is examined. Parameter stability in these four functions is investigated in Section III. In Section IV the order of cointegration rank in the demand for real M1B data is first examined and linear restrictions on the long-run parameters are estimated and tested using Johansen (1991)'s general deterministic trend model. The robustness of the equilibrium income elasticity and interest elasticity is checked using four alternative estimators, (OLS, Stock and Watson 1993, Phillips and Hansen 1990, Johansen 1991).

In CHAPTER 5 the conclusions are presented.

CHAPTER 2

THE DEMAND FOR MONEY: THEORY

I: THEORY REVIEW

1 *The Motives of the Demand for Money*

Keynes (1936) articulates three motives for liquidity-preference, namely, the transactions-motive, the precautionary-motive, and the speculative-motive. These three sources of liquidity-preference generate the demand for money. Baumol (1952), Tobin (1956), Miller and Orr (1966)[1], by emphasizing the choice-making behavior of individuals and giving the transactions demand for money a microfoundaton, go further to investigate the optimal transactions cash holdings for the economic agent. They find that even the transactions demand for cash can be interest-elastic due to the existence of transactions costs between cash and interest-bearing assets. This contrasts with the

[1] These models emphasize that the necessity of holding cash is to bridge the gap between receipts and expenditures. Under the assumptions that transactions are perfectly foreseen and occur in a steady stream, Baumol's and Tobin's models apply reasonably well to salary-earning households. In constrast to the deterministic Baumol/Tobin model, Miller and Orr make the opposite assumption that the net cash flows are completely stochastic, building a money demand model for business firms.

demand for transactions money developed from the quantity theory of money Fisher (1911) and the Cambridge approach (Marshall, Pigou), or Keynes (1936) who postulated that the effect of interest rates on the transactions money is negligible. Tobin (1958) explains that the speculative money demand comes from the risk averse behavior of the agent. The agent is uncertain of the future rate of interest on variable market yield monetary asset so he holds money in his investment balances to diversify the risk from the portfolio to avoid a loss from holding the non-money asset. Other than for the transactions purpose, economic agents hold idle balances in money even when money is dominated by many other monetary assets as a store of value[2].

2 A Missing Chapter in the Theoretical Money Demand Literature: The Time (Energy, Resources) Saving Aspect of the Use cf Money

The distinguishing feature between a barter economy and a money economy is the use of a medium of exchange (generally acceptable media of payment, money) in conducting transactions. In a barter economy, economic agents exchange goods and services for goods and services, whereas in a money economy, economic agents sell goods and services for

[2] It is possible that a risk averter would choose to hold all his investment balances in non-money asset and not to hold money at all. If the agent is risk lover, he certainly plunges his entire investment balances into variable yield monetary assets (see Tobin 1958).

money and use money to purchase goods and services[3]. However money is defined empirically[4], or whatever form it is held in, strictly speaking, by money we mean that money serves four functions: (1) as a medium of exchange, (2) as a standard of unit of account, (3) as a store of value (a temporary abode of purchasing power), and (4) as a final discharge of debts.

The use of money as a medium of exchange and a store of value allows purchases and sales to be conducted independently of one another so it effectively eliminates the requirement of a "double coincidence of wants" in the exchange search process of a barter economy. As a unit of account and a final discharge of debts, it also reduces the number of rates of exchange (prices) between different items quoted in the market and simplifies the information in any transaction. The use of money greatly reduces the time and resource devoted to exchange and therefore facilitates transactions.

Without recognizing the time saving aspect of the use of money, it is hard to explain why money is used at all in the economy. The literature cited above recognizes that money is indispensable in conducting transactions and the

[3] Some transactions, of course, may be conducted through barter in a money-economy.

transactions purpose is the reason why it is included in the portfolio even when other assets with higher yields are available. The analyses presume that money is used in the transactions so the demand for money is always positive, which amounts to imposing a cash-in-advance (Clower) constraint in the model. However the fundamental reason for the use of money is unstated in these models, there is a missing chapter in the theoretical literature in explaining the money demand behavior.

Saving (1971) fills this missing chapter in the money demand literature by emphasizing the medium-of-exchange role of money and time saving aspect of money. In his sophisticated model, it is demonstrated that without imposing a cash-in-advance restriction, we can still obtain a positive transactions demand for money.

In the spirit of Saving (1971) and McCallum (Chapter 3, 1989) models, we build a transactions demand for money model from a utility maximization framework. The model tries to capture the notion in Friedman's (1956) restatement of the quantity theory of money that money is an asset held for the services which it provides and that these services and the low transactions cost associated with exchange of money for other assets (including physical goods) explain the

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[4] Empirical money is defined according to the relative ease (liquidity) with which an asset may be converted into "money", a means of payment.

existence a positive demand for money. Without a cash-in-advance restriction, the solutions of the model have similar properties to those in Lucas (1988), so our model provides a possible explanation for the cash-in-advance constraint in his model. The resulting money demand function also has some properties which may help understand some empirical findings in the money demand literature.

II: TRANSACATIONS DEMAND FOR MONEY MODEL: A SHOPPING TIME MODEL

1 *Assumptions*

We assume complete certainty in order to abstract from precautionary and speculative demand for money and to concentrate on the transaction motive for holding money.

(1) *Preferences*

The agent is assumed to live for T periods. He gets utility out of real consumption of goods and services (c) and leisure time (L) and he prefers to smooth out the consumption over the lifetime by borrowing and lending. Thus, at any point in time, e.g. period 1, preferences are described by the explicit multiperiod utility function as,

$$U = U(c_1, c_2, \dots, c_T; L_1, L_2, \dots, L_T)$$

$$= \sum_{t=1}^{T} \beta^{t-1} \ln c_t + \alpha \sum_{t=1}^{T} \omega^{t-1} \ln L_t \qquad 0 < \alpha, \beta, \omega < 1 \tag{1.1}$$

where β and ω are the intertemporal elasticities of consumption and leisure, respectively. α is the contemporaneous elasticity between consumption and leisure. U has properties of positive marginal utility, and diminishing marginal utility with respect to each c and L.

(2) Time constraint

The agent allocates each period of time (normalized as 1) into leisure (L_t), work (N_t) and shopping (transaction time, S_t). The agent is assumed to be locked in a labor contract, supplying labor services inelastically ($\bar{N}_t$) at real wage rate w_t, which is equivalent to assuming that the household receives real income y_t (= $w_t \bar{N}_t$), the amount of which unaffected by the household's choice. So the agent faces the time constraint

$$L_t + \bar{N}_t + S_t = 1 \tag{1.2}$$

Since barter transactions take time, the transaction time S is assumed as an increasing function of the volume of transactions undertaken (consumption). For a given volume of transactions, the amount of time (and energy) spent is reduced by additional money holding so S is a decreasing function of real money holdings[5]. Therefore, $S_t = g(c_t, m_{t-1})$, where $m_{t-1} = M_{t-1}/p_t$ and g possesses partial derivatives $g_1 > 0$, and $g_2 < 0$, and diminishing marginal effects, $g_{11} < 0$ and $g_{22} > 0$. Since the greater the time spent in shopping, the smaller the amount left over for leisure, it implies that leisure time is a decreasing function in c_t and increasing in m_{t-1}, $\frac{\partial L_t}{\partial c_t} < 0$ and $\frac{\partial L_t}{\partial m_{t-1}} > 0$; $\frac{\partial^2 L_t}{\partial c_t^2} > 0$ and $\frac{\partial^2 L_t}{\partial m_{t-1}^2} < 0$. Specifically, it is assumed to take the form of homogeneous function:

$$L_t = c_t^{-b} m_{t-1}^{a}, \qquad t = 1, 2, \cdots\cdots, T \qquad (1.3)$$

$$0 < a,\ b < 1$$

where a and b stand for transaction technology parameters or transactions costs variables.

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[5] Money holdings facilitate transactions. It is the real quantity of money that matters. With higher prices, greater nominal amounts of money are needed for given real consumption quantities.

The parameter "a" characterizes the time saving aspect of the use of money and may represent the revolution of money and payments technology[6]. For different definitions of money, the parameter "a" differs. For example, we may reason that the use of fiat money is less time consuming than the use of commodity money, in this case $a_f > a_c$. Hence, when there is a monetary reform (e.g. from metallic money to paper money), we may observe changes in the parameter "a" to reflect alterations in service flows from a given real money holdings. We may also infer that a broader measure of money has smaller "a" than a narrow measure because the former has a large proportion of less liquid assets than the latter. In other words, whenever the agent needs to make payment, he has to transfer the less liquid components of his transactions balances into a means of payment first, consequently, this incurs financial transactions costs (costs like in the Baumol-Tobin model), and then reduces the convenience of the use of money. Similarly, for the same definition of money, changes in "a"

[6] Saving (1971) deals with the model in a disaggregate manner, deriving demand function for multiple "moneys" (currency, demand deposits, etc.). In this case, the service flow (being a means of payment) from different "money" is not the same because liquidity in various assets differs and this should be reflected in the transaction time function. That is, different parameter "a" for each "money" may appear in L. Whereas in our model we consider "money" as an aggregate so the parameter "a" may be thought of an index of average service flows from various components of money whatever it is defined.

may reflect some regulatory changes or innovations in payments system which alter the liquidity of money over time and in turn simplify (complicate) the payment process[7]. In other words, for a given volume of transactions, the improvement (deterioration) in payments technology saves even more (less) energy and time by additional money holding.

The parameter "*b*" represents the transactions technology, being determined by geographical factors, or the nature of transactions. For example, we may expect that exchange of goods and services is easier in an economy where the transportation system is highly developed than in an economy where the transportation system is poor. In this sense, we reason that the parameter "*b*" may change over time due to the development of the transportation conditions in an economy.

[7] For example, the automatic transfer from savings to demand accounts (ATS) authorized by depository institutions significantly increase the liquidity of savings deposits, making these deposits serve as transactions balances, consequently, the transaction costs involved in the transfer of less liquidity asset to a medium-of-exchange are reduced.

(3) The multiperiod budget constraint

To specify the multiperiod budget constraint, we first describe the trading process. Following Lucas (1988), we assume that the agent alternates between portfolio transactions and goods trading in lockstep fashion.

At the beginning of each period (t), the agent allocates his nominal wealth (W_{t-1}) carried from the end of last period (t-1) into one-period risk free bonds (B_{t-1}) and money (M_{t-1}), by trading bonds and money in a single centralized market. That is,

$$W_{t-1} = B_{t-1} + M_{t-1}, \quad t = 1,2,\cdots,T \tag{1.4}$$

The bond market is perfectly competitive and the agent is free to borrow or lend at a same rate of r_t. Other than bonds, money is the only asset functioning as a store of value. The direct cost of holding money is δ_t. If $\delta_t > 0$, the own rate of money is positive; for example, in the modern fiat-money world, deposits in the banking system pay positive interest. $\delta_t < 0$ may occur in a commodity-money economy in which perishable goods are used as money and its real (and nominal) value diminishes over time, other things being the same. We assume that the agent enters period 1 with a given initial exogenous nominal wealth W_0[8].

When security trading is concluded, all agents disperse the money acquired in the portfolio transactions to produce or purchase consumption goods at a nominal price p_t.

Hence, at the end of period 1, the agent's nominal wealth W_1 is

$$W_1 = (1+r_1)B_0 + (1+\delta_1)M_0 + p_1 y_1 - p_1 c_1 \tag{1.5}$$

From (1.4) using $W_0 = B_0 + M_0$, and $W_1 = B_1 + M_1$, we can rewrite (1.5) as

$$B_1 + M_1 = (1+r_1)W_0 - (r_1 - \delta_1)M_0 + p_1 y_1 - p_1 c_1 \tag{1.6}$$

Similarly, we generalize the case

$$\begin{aligned} W_t &= (1+r_t)B_{t-1} + (1+\delta_t)M_{t-1} + p_t y_t - p_t c_t \\ &= B_t + M_t, \qquad t = 1,2,\cdots,T \end{aligned} \tag{1.7}$$

Thus, the bond holding at the beginning of each period up to $T+1$ is

[8] This initial condition allows us to solve the model.

$$B_t = (1+r_t)\, B_{t-1} + (1+\delta_t)\, M_{t-1} + p_t\, y_t - p_t\, c_t - M_t$$

$$t = 1, 2, \cdots\cdots, T \qquad (1.8)$$

By recursive substitution, we obtain the bond holdings at the beginning of period $T+1$ (B_T):

$$\begin{aligned}
B_T = {} & (1+r_1)(1+r_2)\cdots\cdots(1+r_{T-1})(1+r_T)\, W_0 \\
& + (1+r_2)(1+r_3)\cdots\cdots(1+r_{T-1})(1+r_T)\, p_1\,(y_1 - c_1) \\
& + (1+r_3)\cdots\cdots(1+r_{T-1})(1+r_T)\, p_2\,(y_2 - c_2) \\
& + \cdots\cdots + \\
& + (1+r_{T-1})(1+r_T)\, p_{T-2}\,(y_{T-2} - c_{T-2}) \\
& + (1+r_T)\, p_{T-1}\,(y_{T-1} - c_{T-1}) \\
& + p_T\,(y_T - c_T) \\
& - (r_1 - \delta_1)(1+r_2)(1+r_3)\cdots\cdots(1+r_{T-1})(1+r_T)\, M_0 \\
& - (r_2 - \delta_2)(1+r_3)\cdots\cdots(1+r_{T-1})(1+r_T)\, M_1 - \cdots\cdots - \\
& - (r_{T-2} - \delta_{T-2})(1+r_{T-1})(1+r_T)\, M_{T-3} \\
& - (r_{T-1} - \delta_{T-1})(1+r_T)\, M_{T-2} \\
& - (r_T - \delta_T)\, M_{T-1} \\
& - M_T \qquad (1.9)
\end{aligned}$$

To obtain maximum utility, the agent has to consume all his nominal wealth at the end of his life T, that is, the terminal constraint (B_T = 0 and M_T = 0). Thus, the agent's multiperiod budget constraint at period 1 is obtained by dividing both sides of (1.9) by $(1+r_1)(1+r_2)\cdots\cdots(1+r_{T-1})(1+r_T)$

$$0 = W_0 + \frac{p_1 y_1}{(1+r_1)} + \frac{p_2 y_2}{(1+r_1)(1+r_2)} + \cdots\cdots + \frac{p_T y_T}{(1+r_1)(1+r_2)\cdots\cdots(1+r_T)}$$

$$- \frac{p_1 c_1}{(1+r_1)} - \frac{p_2 c_2}{(1+r_1)(1+r_2)} - \cdots\cdots - \frac{p_T c_T}{(1+r_1)(1+r_2)\cdots\cdots(1+r_T)}$$

$$- \frac{(r_1-\delta_1)m_0 p_1}{(1+r_1)} - \frac{(r_2-\delta_2)m_1 p_2}{(1+r_1)(1+r_2)} - \cdots\cdots - \frac{(r_T-\delta_T)m_{T-1}p_T}{(1+r_1)(1+r_2)\cdots\cdots(1+r_T)} \tag{1.10}$$

(1.10) can be interpreted as follows.

Multiplying both sides of (1.10) by $(1+r_1)\dfrac{1}{p_1}$ and rearranging terms, we obtain

$$(1+r_1)\frac{W_0}{p_1} + y_1 + \sum_t^T \prod_{i=2}^t (1+r_i)^{-1} \frac{p_t}{p_1} y_t$$

$$= c_1 + \sum_t^T \prod_{i=2}^t (1+r_i)^{-1} \frac{p_t}{p_1} c_t$$

$$+ (r_1-\delta_1)m_0 + \sum_t^T \prod_{i=2}^t (1+r_i)^{-1} \frac{p_t}{p_1}(r_t-\delta_t)m_{t-1}$$

Equivalently,

$$(1+r_1)\frac{W_0}{p_1} + y_1 + \sum_t^T \prod_{i=2}^t (1+r_i)^{-1}(1-\pi_i)^{-1} y_t$$

$$= c_1 + \sum_t^T \prod_{i=2}^t (1+r_i)^{-1}(1-\pi_i)^{-1} c_t$$

$$+ (r_1-\delta_1)\, m_0 + \sum_t^T \prod_{i=2}^t (1+r_i)^{-1}(1-\pi_i)^{-1}(r_t-\delta_t) m_{t-1}, \quad t = 2,3,\cdots,T$$

$$\text{where } \pi_t = \frac{p_t - p_{t-1}}{p_i} \tag{1.11}$$

The left hand side of (1.11) is the agent's life-time resources in terms of period 1 consumption at the end period 1, which is the sum of the real value of non-human wealth (W_0/p_1) and the present discounted value of current and future real labor income streams. The right hand side of (1.11) represents the present discounted values of consumption service flows and real service flow from holding nominal money balances. Note $(r_t-\delta_t)M_t$ is the opportunity cost of holding nominal money; $(1+r_t)(1-\pi_t)$ is real rate of interest over the course of period t.

2 The Constrained Multiperiod Utility Maximization Problem

The agent at period 1 is to choose $\{c_t\}_{t=1}^{T}$ and $\{m_{t-1}\}_{t=1}^{T}$ sequences, subject to the time constraint (1.2), transactions technology (1.3), and the multiperiod budget constraint (1.10), so as to maximize the multiperiod utility function (1.1). Therefore, the Lagrangian (ℓ) for the constrained maximization problem is

$$
\begin{aligned}
\ell = \max_{c_t, m_{t-1}, \lambda} U &= \ln(c_1) + \beta \ln(c_2) + \beta^2 \ln(c_3) + \cdots + \beta^{T-1}\ln(c_T) \\
&+ \alpha[- b\ln(c_1) - \omega b \ln(c_2) - \omega^2 b\ln(c_3) - \omega^{T-1} b\ln(c_T) \\
&\quad + a\ln(m_0) + \omega a \ln(m_1) + \omega^2 a\ln(m_2) + \omega^{T-1} a\ln(m_{T-1})] \\
&+ \lambda[W_0 + \frac{p_1 y_1}{(1+r_1)} + \frac{p_2 y_2}{(1+r_1)(1+r_2)} + \cdots\cdots + \frac{p_T y_T}{(1+r_1)(1+r_2)\cdots\cdot(1+r_T)} \\
&\quad - \frac{p_1 c_1}{(1+r_1)} - \frac{p_2 c_2}{(1+r_1)(1+r_2)} - \cdots\cdots - \frac{p_T c_T}{(1+r_1)(1+r_2)\cdots\cdot(1+r_T)} \\
&\quad - \frac{(r_1-\delta_1)m_0 p_1}{(1+r_1)} - \frac{(r_2-\delta_2)m_1 p_2}{(1+r_1)(1+r_2)} - \cdots\cdots - \frac{(r_T-\delta_T)m_{T-1} p_T}{(1+r_1)(1+r_2)\cdots\cdot(1+r_T)}]
\end{aligned}
$$

where λ is a Lagrange multiplier.

The first order conditions for the constrained maximization problem are

$$\frac{\partial \mathcal{L}}{\partial c_t} = [1 - b\alpha(\frac{\omega}{\beta})^{t-1}]\beta^{t-1} c_t^{-1} - \lambda p_t (1+r_1)^{-1}(1+r_2)^{-1}\cdots(1+r_t)^{-1}$$

$$= 0 \text{ if } c_t > 0$$

$$< 0 \text{ if } c_t = 0, \quad t = 1,2,\cdots,T \qquad (1.12)$$

$$\frac{\partial \mathcal{L}}{\partial m_{t-1}} = a\alpha\omega^{t-1} m_{t-1}^{-1} - \lambda p_t (r_t - \delta_t)(1+r_1)^{-1}(1+r_2)^{-1}\cdots(1+r_t)^{-1}$$

$$= 0 \text{ if } m_{t-1} > 0$$

$$< 0 \text{ if } m_{t-1} = 0,$$

$$t = 1,2,\cdots,T \qquad (1.13)$$

$$\frac{\partial \mathcal{L}}{\partial \lambda} = 0$$

$$= W_0 + \frac{p_1 y_1}{(1+r_1)} + \frac{p_2 y_2}{(1+r_1)(1+r_2)} + \cdots + \frac{p_T y_T}{(1+r_1)(1+r_2)\cdots(1+r_T)}$$

$$- \frac{p_1 c_1}{(1+r_1)} - \frac{p_2 c_2}{(1+r_1)(1+r_2)} - \cdots - \frac{p_T c_T}{(1+r_1)(1+r_2)\cdots(1+r_T)}$$

$$- \frac{(r_1 - \delta_1)m_0 p_1}{(1+r_1)} - \frac{(r_2 - \delta_2)m_1 p_2}{(1+r_1)(1+r_2)} - \cdots - \frac{(r_T - \delta_T)m_{T-1} p_T}{(1+r_1)(1+r_2)\cdots(1+r_T)}$$

$$(1.14)$$

Note if $r_t < \delta_t$, that is, bonds are dominated by money as a store of value, money is definitely held whereas no bonds would be held. This is a less interesting case for our analysis. So we assume that $r_t > \delta_t$ $\forall t$ and investigate the case where a positive money demand is possible even when money is dominated as a store of value. In this case, the only reason for the agent to hold money is to carry out transactions, therefore the restriction

$$M_{t-1} \leq p_t c_t, \text{ or } \frac{m_{t-1}}{c_t} \leq 1 \tag{1.15}$$

must hold.

Assuming that $\{c_t\}_{t=1}^{T}$ and $\{m_{t-1}\}_{t=1}^{T}$ are all positive, we can solve for their optimal values in terms of exogenous variables (demand functions) as follows.

To eliminate λ and express $\{c_t\}_{t=1}^{T}$ and $\{m_{t-1}\}_{t=1}^{T}$ in terms of m_0, using $\frac{\partial \mathcal{L}}{\partial m_0} = a\alpha m_0^{-1} - \lambda p_1 (1+r_1)^{-1} = 0$ and other first order conditions from (1.12) and (1.13), we obtain

$$c_1 = \frac{1-b\alpha}{a\alpha} (r_1 - \delta_1) m_0 \tag{1.16}$$

Note that $0 < b, \alpha < 1$ so $(1-b\alpha) > 0$.

$$c_t = \frac{[1-(\frac{\omega}{\beta})^{t-1}]\beta^{t-1}}{a\alpha} (r_t - \delta_t)(1+r_2)(1+r_3)\cdots(1+r_t) \frac{p_1}{p_t} m_0$$

$$t=2,3,\cdots,T \tag{1.17}$$

$$m_{t-1} = \omega^{t-1} \left(\frac{r_1-\delta_1}{r_t-\delta_t}\right)(1+r_2)(1+r_3)\cdots(1+r_t)\frac{p_1}{p_t} m_0$$

$$t=2,3,\cdots,T \qquad (1.18)$$

Substitute (1.16)-(1.18) into (1.14), divide both sides by $(1+r_1)^{-1}$ and then let

$$W =$$

$$[(1+r_1)W_0 + p_1 y_1 + \frac{p_2 y_2}{(1+r_2)} + \frac{p_3 y_3}{(1+r_2)(1+r_3)} + \cdots + \frac{p_T y_T}{(1+r_2)\cdots(1+r_T)}] / p_1$$

$$= (1+r_1)\frac{W_0}{p_1} + y_1 + \sum_t^T \prod_{i=2}^t (1+r_i)^{-1}(1-\pi_i)^{-1} y_t$$

we obtain the demand for real money balances at the beginning of period 1:

$$m_0^{*} = a\alpha \frac{1}{r_1-\delta_1} D^{-1} W \qquad (1.19)$$

where $D = [\sum_{t=0}^{T-1}\beta^t + (a-b)\alpha\sum_{t=0}^{T-1}\omega^t] = [\frac{1-\beta^T}{1-\beta} + (a-b)\alpha\frac{1-\omega^T}{1-\omega}]$.

Note that as $T\rightarrow\infty$, $D = [\frac{1}{1-\beta} + (a-b)\alpha\frac{1}{1-\omega}]$.

Therefore, the demand function for $\{c_t\}_{t=1}^T$ and $\{m_{t-1}\}_{t=1}^T$ sequenes are obtained from (1.16)-(1.18) as

$$c_1^* = (1-b\alpha)\, D^{-1} W \tag{1.20}$$

$$c_t^* = [1-(\frac{\omega}{\beta})^{t-1}]\,\beta^{t-1}\,(1+r_2)\,(1+r_3)\cdots(1+r_t)\,\frac{p_1}{p_t}\,D^{-1} W$$

$$= [1-(\frac{\omega}{\beta})^{t-1}]\,\beta^{t-1}\,(1+r_2)\,(1+r_3)\cdots(1+r_t)\,\frac{p_1}{p_2}\,\frac{p_2}{p_3}\cdots\frac{p_{T-1}}{P_T}\,D^{-1} W$$

$$= [1-(\frac{\omega}{\beta})^{t-1}]\,\beta^{t-1}\,[\prod_{i=2}^{t}(1-\pi_i)(1+r_i)]\,D^{-1} W$$

$$t = 2, 3, \cdots, T \tag{1.21}$$

$$m_{t-1}^* = \omega^{t-1}\, a\,\alpha\frac{1}{(r_t-\delta_t)}\,(1+r_2)\,(1+r_3)\cdots(1+r_t)\,\frac{p_1}{p_t}\,D^{-1} W$$

$$= \omega^{t-1}\, a\,\alpha\frac{1}{(r_t-\delta_t)}\,[\prod_{i=2}^{t}(1-\pi_i)(1+r_i)]\,D^{-1} W$$

$$t = 2, 3, \cdots, T \tag{1.22}$$

where $\pi_t = \dfrac{p_t - p_{t-1}}{p_t}$

3 Properties of the Money Demand Function

From the multiperiod utility maximization framework, we obtain the current desired money holdings written as

$$m_0^* = a\,\alpha\frac{1}{r_1-\delta_1}\,D^{-1} W\,, \tag{1.19}$$

This can be expressed in the log form

$$\ln(m_0^*) = \ln(a\alpha D^{-1}) - \ln(r_1-\delta_1) + \ln(W) \tag{1.23}$$

where $W = (1+r_1)\dfrac{W_0}{p_1} + y_1 + \sum_t \prod_{i=2}^{t}(1+r_i)^{-1}(1-\pi_i)^{-1}y_t$ and

$D = [\dfrac{1-\beta^T}{1-\beta} + (a-b)\alpha\dfrac{1-\omega^T}{1-\omega}]$

Therefore, the desired real money balances depends on four major factors:

(a) the agent's life-time resources (W).

(b) the nominal return on money (δ_1) and on the alternative asset (bond) (r_1).

(c) the tastes and preferences of the agent represented by parameters (α,β,ω).

(d) transactions technology (a,b)

To simplify, we write (1.19) as

$$m_0^* = f(r_1,\delta_1;u)W \tag{1.24}$$

where u contains (α,β,ω,a,b).

From above, a number of properties are in order about this function:

First, two rates of interest rates entering the function are one-period rates and no other future rates appear. In other words, given real wealth W, it is the short-term interest rate that is relevant in this model.

The directions of the response of m_0^* to r_1 and δ_1 are

$$\frac{\partial m_0^*}{\partial r_1} = -a\alpha \frac{1}{(r_1 - \delta_1)^2} D^{-1} W < 0 \qquad (1.25)$$

$$\frac{\partial m_0^*}{\partial \delta_1} = a\alpha \frac{1}{(r_1 - \delta_1)^2} D^{-1} W > 0 \qquad (1.26)$$

which display usual substitution effect.

Second, as in all demand analyses resting on maximization of a utility function defined in terms of "real" magnitudes, this function is independent of nominal units used to measure money variables. If the unit in which prices and money income are expressed is changed, the amount of money demanded should change proportionately. That is, nominal money demand (M_0^*) function is homogeneous of the first degree in W_0 and p_t. As seen in (1.19), the demand for real balances is expressed as a function of "real" wealth independent of nominal monetary values. In particular, m_0^* is proportional to W, and it has a unitary wealth elasticity, other things being the same. As Friedman's (1957) permanent income hypothesis suggests that permanent income is treated as the income flow resulting from a stock of wealth, we define $y_p = rW$, where r is the yield applicable to wealth so (1.19) can also be expressed as

$$m_0^{\bullet} = \frac{f(r_1, \delta_1; u)}{r} y_p \tag{1.27}$$

and (1.27) in nominal terms becomes

$$M_0^{\bullet} v = Y_p \tag{1.28}$$

where $v = [\frac{f(r_1, \delta_1; u)}{r}]^{-1}$. In this form (1.28) is in the usual quantity theory form. v is (permanent) income elasticity. Third, other than the wealth constraint and substitution effect on the demand for money, the tastes and preferences of the agent and the transactions condition also determine the form of the demand function. These effects are summarized by a variable $u = u(\alpha, \beta, \omega, a, b)$. The effect of u may be better examined as that of the issue of the stability of the money demand function.

4 Stability of the Money Demand Function

The empirical issue regarding the stability of money demand function can be investigated by understanding the preference parameters (α, β, ω) and transactions technology parameters (a, b). Note any changes in tastes (U) or the transactions technology (L) induced by social, political and economic changes can fundamentally change in the specification of U and L and then alter the functional form of the money demand function, which certainly causes instability of the money demand function. Nevertheless we assume that these functional specifications are the same under different circumstances. In other words, we do not complicate the issue by considering the problem of the Lucas critique (1976). In the following, we consider the effect of the parameters "a" and "b" on $m_0^{\bullet}$. The partial derivatives are computed as

$$\frac{\partial m_0^{\bullet}}{\partial a} = \alpha \frac{1}{r_1 - \delta_1} \left[\frac{1-\beta^T}{1-\beta} - b\alpha \frac{1-\omega^T}{1-\omega} \right] D^{-2} W \tag{1.29}$$

$$\frac{\partial m_0^{\bullet}}{\partial b} = a\,\alpha^2 \frac{1}{r_1 - \delta_1} \left(\frac{1-\omega^T}{1-\omega} \right) D^{-2} W > 0 \tag{1.30}$$

As mentioned before, changes in the parameter "a" may reflect the shift of the monetary regime (e.g. commodity versus money economy), financial regulatory changes or innovations in payment technology, which in turn affect the costs of using money and then change the desired money holdings, for a given wealth and opportunity cost. The direction of the effect of the parameter "a" on m_0^* depends on sign of $[\frac{1-\beta^T}{1-\beta} - b\alpha\frac{1-\omega^T}{1-\omega}]$; when

$$0< b \le \frac{(1-\omega)(1-\beta^T)}{\alpha(1-\beta)(1-\omega^T)},\quad \frac{\partial m_0^*}{\partial a} \ge 0$$

Furthermore, a change in the parameter "b" alters m_0^* in the same direction. Intuitively, when a direct exchange of goods and services becomes harder, the agent is motivated to use a medium of exchange to conduct transactions to reduce transactions costs and therefore hold more money.

The effect of transactions costs on the structure of transactions (barter versus money exchange) can be understood by examining the relationship between m_0^* and transactions c_1^* from (1.16)

$$m_0^* = \frac{a\alpha}{1-b\alpha}\frac{1}{r_1-\delta_1}c_1^* \qquad (1.16)$$

Define the average amount of money held per dollar of transactions

$$h = \frac{m_0^{*}}{c_1^{*}} = \frac{a\alpha}{1-b\alpha}\frac{1}{r_1-\delta_1}$$

As we argued above, (1.15) must hold, that is $h \leq 1$, therefore,

$$a\alpha \leq (1-b\alpha)(r_1-\delta_1) \tag{1.31}$$

and

$$0 < a \leq \frac{(1-b\alpha)(r_1-\delta_1)}{\alpha}$$

The partial derivative of h with respect to the parameters "a" and "b" are

$$\frac{\partial h}{\partial a} = \frac{\alpha}{1-b\alpha}\frac{1}{r_1-\delta_1} > 0 \tag{1.32}$$

$$\frac{\partial h}{\partial b} = \frac{a\alpha^2}{(1-b\alpha)^2}\frac{1}{r_1-\delta_1} > 0 \tag{1.33}$$

That is, the ratio of transactions with money to total transactions is increasing in transactions costs parameters "a" and "b". If payments become easier or barter transactions become cumbrous, the use of money is encouraged in the exchange process so the proportion of the trades through barter are reduced. Note the amount of transactions with money is m_0^*, therefore,

$$(c_1^* - m_0^*) = (1 - \frac{a\alpha}{1-b\alpha}\frac{1}{r_1-\delta_1})\, c_1^* \geq 0 \tag{1.34}$$

is the quantity of transactions via direct barter.

Note (1.16) is not a demand function in our model. However, this relationship between desired consumption, the demand for real balances, and the nominal interest rates is frequently studied in empirical work so we also investigate the stability of this relationship. As in the analysis of the "true" money demand function (1.19), a stable relationship between the desired real balances, transactions variable and opportunity cost variable is determined by the constancy of tastes (α, β) and transactions technology (a, b). In practice, measured income (y, such as GDP) is frequently used as a proxy of transactions variable (c). If transactions and income (or final output) obey a constant relationship over time, say $c_t = n\, y_t$, where n is a constant, (1.16) can be written in terms of y as

$$m_0^* = \frac{a\alpha}{1-b\alpha}\frac{1}{r_1-\delta_1}\, n\, y_1 \tag{1.35}$$

Thus, we expect (1.35) has same properties as (1.16). Furthermore, the ratio of money to income, the reciprocal of (current) income velocity, can be expressed as

$$\frac{m_0^*}{y_1} = \frac{1}{v_y} = n\frac{a\alpha}{1-b\alpha}\frac{1}{r_1-\delta_1} \qquad (1.36)$$

which is independent of income and negatively related to the spread of interest rates, other things being equal.

5 *Cash-in-Advance Model versus the Shopping Time Model*

Next we compare the transactions money demand model developed by Lucas (1988) and our model. By assuming a constant relative risk aversion utility function of a homogeneous of degree one function of consumption and imposing a cash-in-advance constraint,

$$P\,a\cdot c \leq M \text{ (Eq.4 in Lucas)} \qquad (1.37)$$

where P is the price of goods, a and c are vectors; $a_i \in [0,1]$ is the fraction of purchases of good i that must be covered by money, Lucas (1988) derives a proportional relationship between desired real money balances and consumption

$$\frac{M}{P} = \sum_i a_i g_i(r)c = h(r)c \quad \text{(Eq.15 in Lucas)} \tag{1.38}$$

where $c = k(Q)W$, (Eq.16 in Lucas); Q are securities prices; W is wealth.

Interestingly, we notice that (1.38) has a similar form to be

$$\dot{m_0} = \frac{a\alpha}{1-b\alpha}\frac{1}{r_1-\delta_1}\dot{c_1} \tag{1.16}$$

in our model. From his cash-in-advance constraint (1.37), we reason that Lucas (1988) implicitly recognizes that barter exchange is possible since

$$P\,a\cdot c \leq M \leq P\cdot c$$

However, he does not consider it explicitly in the model. Hence, by assuming a logarithm utility function of consumption and leisure[9] (a monotonic transformation of Cobb-Douglas function) and the Cobb-Douglas leisure function of consumption and real balances, our model gives a possible

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[9] When we considered different specifications of the utility functions, $U(c_t, L_t) = \frac{1}{1-\gamma}[\frac{c_t^{1-\alpha}-1}{1-\alpha} + d\frac{L_t^{1-\beta}-1}{1-\beta}]^{1-\gamma}$, see (Mankiw, Rotemberg, and Summers, 1985), the proportional relationship between real balances and wealth (or consumption) is no longer maintained in the model.

missing explanation for the existence of his cash-in-advance constraint, namely, the time saving aspect of the use of money. Note the utility function is also a constant relative risk aversion function[10]. Our model demonstrates that even without imposing that money must be used, we can still generate a positive money demand for transactions purpose even when money is dominated as a store of value.

6 Empirical Issues: Arguments in the Money Demand Function

Since an agent cannot hold all the money he might want, in practice, there is the question of the constraint that is imposed on money balances - whether the appropriate constraint is a measure of wealth, permanent income, or current income. In our model, the first two legitimately enter the "demand function" for money, which can be seen from (1.19) and (1.27).

Traditionally the use of current income as an argument in the demand function for money often has been associated with the notion that money is used primarily to effect a given transactions volume (Keynes 1936, Baumol 1952, Tobin 1956, Miller and Orr 1956), and that a demand for cash

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[10] The risk aversion coefficients are $\frac{-u''(c_t)c_t}{u'(c_t)}$ = 1 and $\frac{-u''(L_t)L_t}{u'(L_t)}$ = 1.

balances depends on costs and yields. In our model, even the money demand function is subject to a wealth constraint, the model is also able to capture this relationship. This can be seen from

$$\dot{m}_0 = \frac{a\alpha}{1-b\alpha}\frac{1}{r_1-\delta_1}\dot{c}_1 \tag{1.16}$$

This relationship is frequently studied in the empirical work. Interestingly, the desired money holdings are proportional to consumption and negatively related with the opportunity cost variable. Several empirical studies, such as, (Lucas 1988, Rasche 1990, Hoffman and Rasche 1991, Stock and Watson 1993, Hoffman, Rasche, and Tieslau 1995, Rasche and Hoffman 1996) find a unitary income elasticity and a negative short-term interest effect. These studies find this relationship as a steady-state equilibrium property. If, in addition, in the steady-state $c-ny$ is stationary, then the results follow our model. Thus, this model may help understand these empirical findings.

A second issue is concerned with whether a short-run interest rate or a long-term rates is the relevant variable in the money demand function[11]. Unlike Poole (1988), among

[11] The importance of this issue can be understood through the following paragraph quoted from Laidler (1966). 'If investment is more sensitive to long rates of interest and if the first impact of a change in the money supply is on short rates, then in order to have a theory as to how

others, who has advocated the use of a long-term in the money-demand function, our model suggests that a short-term rate is the appropriate argument since, given real wealth, no other future rates appear in the function.

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changes in the money stock affect the economy, one must have a theory as to how interest rates of various terms are interrelated. Thus, ... the theory of term structure of interest rates comes to be a central topic, essential to any description of the mechanism by which changes in the money supply affect the real variables in an economy.' (pp.554-555).

III: THE STABILITY OF THE AGGREGATE MONEY DEMAND FUNCTION AND ITS POLICY IMPLICATION

Although the Keynesian and Monetarists have emphasized the "demand" side of the economy in their policy prescriptions, they held a different view of the stability/instability of the aggregate demand for money relationship.

For example, Keynes (1936) explicitly postulated potential instability in the aggregate demand for money function, which was asserted to shift erratically and unpredictably with rumor and expectations so he doubted the effectiveness of monetary policy[12]. Keynesian analysis usually treated monetary policy as subsidiary to fiscal policy as an income-stabilization device. On the contrary, Friedman (1956) stated that the demand for money was "empirically" stable function of a few arguments. He argued that a stable money demand for money function played a vital role in the analysis of the economy as a whole, such as the

[12] Keynes (1936) viewed that money affected the economy through its direct effect on the interest rate, which in turn affected the investment and employment, and finally through the multiplier process (if marginal propensity to consume is less than 1) to affect the aggregate demand and the output (or employment) However, he doubted that the quantity of money was able to work its way into the economy in the end. For example, the changes in interest rate may be small since the speculative demand for money may shift around due to the expectation of the monetary policy or the investment is inelastic to the interest rate (see Chapter 13, Keynes 1936).

level of money income or of price. Although he believed that money mattered in the short-run and recognized that manipulation of the quantity of money was a powerful (perhaps dangerous) policy tool, Friedman (1968) doubted the ability of the use of money to stabilize the economy due to both inside lag and outside lag in the monetary policy of a central bank. Consequently, he developed a policy idea of the money-supply growth rule.

Although the stability/instability of money demand relationship is not the only difference and would not be the decisive factor which distinguished Monetarism from Keynesian[13], we recognize that the stability issue is the central assertion in Monetarism and it is certainly associated with the importance of the quantity of money in the economy. To fix the idea that money matters and how a stable money demand relationship can help formulate monetary policy, as a first step we need to understand the quantity theory of money.

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[13] In a simple IS-LM analysis framework, how effective monetary policy can depend more than the slope of the LM curve. The responsiveness of aggregate demand to changes in interest rates as well as the direct responsiveness of expenditures to changes in the quantity of money are just as important. A stable money demand function does not guarantee the potence of monetary policy.

1 *Quantity Theory of Money and Monetarism*

The ultimate objectives of monetary policy are to promote price stability and economic growth. The efficacy of using monetary aggregate targets to achieve ultimate policy objective depends importantly on the existence of a stable or predictable demand for money. To put it another way, the velocity of money must be reasonably stable and predictable. From the equation of exchange (Fisher 1911), we know

$$MV = PY \tag{1.39}$$

in terms of growth rates

$$\dot{M} + \dot{V} = \dot{P} + \dot{Y}. \tag{1.40}$$

Substituting $M^s = M^d$,

$$V = \frac{PY}{M^d}$$

where M, Y, and P denote the nominal money stock (however defined), real GNP, and general price level, respectively; V represents income velocity, interpreted as a measure of how much money firms and households desire to hold relative to the level of nominal income. If, in the extreme case where

V is constant, as Fisher (1911) claimed, determined by social and economic factors, or the demand for money is inelastic with respect to the variables in M^d, (1.39) is the model of the determination of money income. In addition, if, in the long-run, the real output (income) is determined by "real" side of economy (factors of production, technology and relative prices) such that output is fixed at full employment level, (1.39) says that the money stock determines the general price level. Keeping the nominal supply on a stable growth will stabilize the behavior of the price level.

Friedman (1956) reformulated the quantity theory by arguing that V did not have to be numerically constant over time, but required the stability in the functional relation between V (or the quantity of money demanded) and the variables that determine it. From (1.39)-(1.40), we learn that as long as V is predictable or if the growth in the demand for money accompanying a given growth of income can be predicted, the monetary authority can influence growth of money income (aggregate demand) by controlling monetary growth. Inaccurate predictions of velocity growth can lead to monetary targets that are incompatible with ultimate objective. For example, underestimation (overestimation) of the growth in velocity can lead the monetary authority to follow monetary policies that are unduly expansionary (restrictive), resulting in unacceptably high inflation (or recession).

The principal of the quantity theory of money is easy to understand through the real-balance effect[14]; changes in the money stock may result in changes in real output or prices or both in the short-run. The monetary authority can attempt to achieve the short-run output-inflation combination most consistent with its policy objective by influencing the level of aggregate demand (nominal income). However, (1.39) does not tell us how much a change in PY is reflected in Y and how much in P. To infer this, we need to specify all the determinants of the variables associated with (1.39) to capture the complicate transmission from money to price and output which is hidden in (1.39).

Economists may have different views of the way an economy works and then build their own view of macroeconomic models, however, we will not go further about these models, (for the macroeconomic models of Monetarism,

[14] Friedman (1974) illustrated the quantity theory of money as follows. 'Suppose that the nominal quantity that people hold at a particular moment of time happens to correspond at current prices to real quantity larger than the quantity that they wish to hold. Individuals will then seek to dispose of what they regard as their excess money balances; they will seek to pay out a larger sum for the purchase of securities, goods and services, for the payment of debts, and as gifts.....', 'If prices and income are free to change, the attempt to spend more will raise the volume of expenditures and receipts, expressed in nominal units, which will lead to a bidding up of prices and perhaps also

see Friedman 1974, Monetarism, edited by Stein, 1976). We have seen that a stable money demand function is the central proposition of monetarist models. We also observe that this relationship is an important element in the new classical economics models Sargent and Wallace (1975), New Keynesian analysis (Mankiw, 1991), and in some empirical real business cycle models (King, Plosser, Stock, and Watson, 1991) as well.

We should note that the monetary authorities may use a dynamic macroeconomic model to compare alternative monetary policies (e.g. pegging interest rates versus pegging monetary aggregates) in order to choose a preferable policy instrument to achieve its policy goal. As Lucas (1976) suggests, in making policy comparisons, it is necessary that the model of private behavior, such as the consumption function, the investment function, and the money demand function, be "structural", that is, invariant to the policy rule in effect. A shift in the empirical money demand function due to a policy regime shift (or due to other reasons) certainly invalidates these policy analyses.

From above, we may realize how empirical evidence against a stable money demand function can invalidate these macroeconomic models and undermine the importance of the monetary policy. In the CHAPTER 3 and CHAPTER 4, we will

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to an increase in output (pp.2-3, "A Theoretical Framework investigate the empirical evidence of the stability of the aggregate money demand functions in Taiwan over the period of 1961:4-1997:3 using various econometric techniques. First we describe the variables chosen for the study of the Taiwanese demand for money in the final section of CHAPTER 2.

IV: THE ARGUMENTS IN THE TAIWANESE DEMAND FOR MONEY FUNCTIONS: DATA DESCRIPTION

Economic theories (Keynes 1936, Baumol 1952, Tobin 1956, 1958, Friedman 1956, Miller and Orr 1966, Saving 1971, Lucas 1988, McCallum 1989) suggest that the demand for money can be explained by functional relationships which include a relatively small number of arguments. They can be a scale variable, such as permanent income, wealth, or current income, and an opportunity cost variable such as a nominal interest rate or some measure of the expected inflation rate. If nominal balances have been the dependent variable, the general price level is also included in the function. Algebraically, the demand for money conventionally takes the form of log linear function

$$m_t^{*} = f\,(y_t, R_t) + p_t, \quad f_y > 0, \quad f_R < 0 \qquad (1.41)$$

for Monetary Analysis", Friedman 1974).

where m_t^* is the log of quantity of money demanded; y is scale variable; R a vector of interest rates on the alternatives to money; p the log of the general price level. Since economic theory predicts that the demand for money is a demand for "real" balances (money holdings measured in constant purchasing power terms), a price level elasticity of demand for nominal balances is frequently constrained to one. In addition, f is increasing in y and decreasing in the elements of R.

Many empirical studies estimate the specification

$$(m_t - p_t) = (m_t^* - p_t) = f(y_t, R_t) \tag{1.42}$$

(1.42) is empirically referred to the long-run (equilibrium) money demand function which assumes that the observed quantity of real money balances $(m_t - p_t)$ corresponds to the quantity of real balances that are desired at the current value of the arguments of the money demand function. It is distinguished from the short-run demand function[15] (Goldfeld

[15] The widespread practice in the short-run money demand analysis is to add a lagged dependent variable to the equilibrium specification (1.42). Such stock adjustment specifications are motivated by cost minimizing behavior in the presence of quadratic costs of adjustment toward the agent's long-run demand for money function. In the short-run, some dynamics may appear in the equilibrium specification so that the current money balances can be explained not only by current values of the arguments but also by the lagged real balances and lagged values of the

1973, 1976) developed from the partial adjustment model (Chow 1966), see Rasche and Hoffman 1996 for the clarification of the long-run specification, and the criticisms on the partial adjustment models.

The choice and the measurement of the variables for the study of Taiwanese demand for money which reflect the corresponding economic notions and the data available are described in the following.

1 *Measure of Money*

A stable aggregate demand for money function is an important element in macroeconomics, however we do not have theories suggesting the appropriate definition of money for monetary analysis and particular empirical measures as clearly superior to others which yield more stable money demand relationships. To compare the performance of the money demand functions in Taiwan for alternative definitions of money for the same time period (1961:4-1997:3), we will study four measures of money, namely M1A, M1B, M1BP, and M2.

Currently, the Central Bank in Taiwan has three definitions of monetary aggregates, M1A, M1B, and M2, publishing these three series on a monthly basis. M1A and

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arguments. Hence more general specification of the demand for real balances can be written as $m_t = f(y_t, R_t;$ lagged $y_t, R_t, m_t)$.

M1B are narrowly defined monetary aggregates consisting of currency in circulation and the demand deposits of the individuals and enterprises in the monetary institutions[16]; they largely serve as transactions balances. M2 is a broader concept of money; it consists of M1B and a large proportion of less liquid assets including time and time saving deposits, foreign currency deposits of individuals and enterprises in the monetary institutions, all the deposits in the Postal Savings System (P.S.S), and others. Data on these three aggregates are available from July 1961 in Financial Statistics Monthly. In February 1997, the Central Bank in Taiwan changed the definition of money, especially in the definition of M2 by excluding less liquid components[17], and including new components[18] to take account of financial innovations. The data used in our study are redefined monetary aggregates and are retrieved from the FSM data bank of the Taiwan Economic Data Center[19]. We report the components in the Appendix 1.

[16] Monetary institutions are defined as financial institutions which are able to create money (Financial Statistics Monthly, November 1997, pp.190).

[17] These include bank debentures issued, savings bonds and treasury bills-B issued by the central bank held by enterprises or individuals and foreign exchange trust funds.

[18] They are repurchase agreements and non-resident NT (New Taiwan dollar) deposits.

[19] I wish to thank Professor Lee, C.N. for his assistance with the data in my study. The web site address

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We notice that the passbook savings deposits in the Postal Saving System (P.S.S.)[20] are included in M2 but are excluded in M1B[21]. These postal deposits are also transactions related deposits and are similar to the deposits included in M1B. Furthermore, the amount of this component is nontrivial, around 30% of M1B (on average 1980:12 - 1992:12, see Table 4.2 in C.N Lee Ph.D dissertation, Department of Economics, Michigan State University, 1994). A priori we will reason that the aggregation of the same type of deposits may reveal similar money demand behavior of Taiwanese people. Thus, for experimentation, we construct another monetary aggregate (M1BP) by adding these transactions related deposits of the

of Taiwan Economic Data Center is http://140.111.1.22/moecc/rs/pkg/tedc/tedc2.htm.

[20] The Postal Savings System is a major financial competitor in the deposit market, competing with the domestic banks for demand and time deposits (Semkow 1992, pp.43-44). According to the Central Bank's compilation of monetary aggregates, the Bank groups monetary assets by the type of institution (monetary versus financial institutions), not by the similar type of deposits. Because the Postal Savings System is not defined as a monetary institution, its deposits are not supposed to be counted in the monetary aggregates. However, in the new definition, all deposits are included in M2 as a non-M1B component.

[21] In the study of the equilibrium M1 demand for Japan, Rasche (1990) makes the observation that deposits in the Postal Savings System are excluded in M1 measure which is the most comprehensive measure of transactions money available for Japan.

P.S.S into M1B[22] to see if this seemingly homogeneous monetary aggregate yields quantitatively and qualitatively different conclusions from M1B.

The quarterly average series for each monetary aggregate is measured as the geometric average of the three end-of-month observations for 1961:7 - 1997:9.

2 Scale Variable

Whether the appropriate constraint imposed on money balances is a measure of wealth, or income remains an open question.

The use of income as a constraint in the money demand function is often associated with the notion that money is used principally to carry out a given transactions volume, for example, (Baumol 1952, Tobin 1956, Miller and Orr 1966). The wealth constraint emphasizes the role of money as a productive asset and focuses attention on the equilibrium of the balance sheet, and the allocation of assets and the services that money provides, (e.g. Friedman 1956, Tobin 1958, Saving 1971, Lucas 1988, McCallum 1989). The model we derived in Section II suggests that both a measure of wealth and of income are relevant variables entering the demand for money relationship. As in many empirical money demand

[22] The aggregate (M1BP) is constructed by the sum of the components with retrieval codes listed in the FSM databank M1BP = M1B + LPSS@SD@TA + LPSS@SD@PSD.

studies, we use real GNP as a measure of the scale variable for the four specifications of the money-demand function[23]. The data are from the NIAQ databank of Taiwan Economic Data Center.

The four measures of money and real GNP are deflated by the GNP deflator (1991 = 100). In the following study, they are denoted as RM1A, RM1B, RM1BP, RM2, and y.

3 *Measure of the Opportunity Cost of Holding Money Balances*

The rate of interest entering the demand function is an important variable in a macroeconomic model since it carries the transmission mechanism from a change in the money policy to other sectors in the economy. Hoffman and Tahiri (1994) point out that there is no conclusive evidence as to whether money demand in the developing countries is sensitive to an interest rate measure. Hence, to find a theoretically and empirically relevant variable would help us understand the working of the economy in Taiwan and help the monetary authority select an appropriate policy.

Conventionally, bank deposits rates (e.g. one-, three-, six- month time deposit rate) are employed as a

[23] In the Taiwanese money demand literature, Lyo (1970) used both wealth and real GNP (and unorganized money market rate as an opportunity cost variable) to study the long-run demand for real M1 (currency + demand deposits) and real M2 (M1 + time and savings deposits). The data are annual data (1947-1968).

representative of "the" interest rate variable in the study of the money demand functions in Taiwan. Perhaps it is because of the availability of the data. In Financial Statistics Monthly, the banks rates have longest records (starting 1961:7) and the data on the money market rates, which are frequently used in the money demand studies in the developed countries, are not available until November 1980[24]. It appears that the choice of "the" rate is underdiscussed in the Taiwanese money demand literature since we observe that some studies mistakenly treated a proxy of the own rate as the opportunity cost of holding M2. Moreover, a negative coefficient on the own rate is frequently obtained in the demand for M2 studies[25]. Shen (1995) first makes this observation but does not provide suggestion of an appropriate measure of opportunity cost of holding M2. This fundamental issue should not be overlooked before we go any further to examine the money demand relationships otherwise interpretations of results can be very misleading.

[24] The money market was established in 1976. Note the treasury-bill (TB) market was established in April, 1973 and that the TB rates are available from October 1973 on in Financial Statistics Monthly.

[25] See Lin (1997) for demand for M2 function literature review.

Since bank rates were subject to the ceiling rates set by the central bank before 1989[26], the structure of deposits rates did not vary freely with the money market condition during much of the sample period considered in this study. They may not adequately capture changes in the opportunity cost of holding money balances. Hence, the market determined rate may be a better measure. We argue that another interest rate should also be included in Taiwanese researchers' choice list of opportunity cost variables in the money demand study. In the following, we discuss why we regard the own rate of money and the unorganized money market (U.M.M.) rate as relevant variables a priori[27].

(1) The own rate on money

Since deposits (money) in the commercial banks pay interest and a higher (lower) deposit rate attracts (discourages) the public to hold these assets (due to the same direction of both substitution and income effect), the own rate is a relevant variable in the money demand

[26] Interest rate regulation in Taiwan dates back to 1947. Since then, the maximum rates for various kinds of deposits and loans were determined by the central bank (Wu,1995). The process of interest rates deregulation began in 1980. In 1989, the interest rate regulation was totally removed (Semkow Ch.6 1992).

[27] A couple of Taiwanese money demand studies employed the U.M.M. rate as a cost variable of holding money (e.g. Lyo 1970, Shui 1983, Lin 1997) but little was said about its relevance and its importance.

relation. Unless the own rate is close to zero, we should include it in the study to see if the effect of the own rate is of statistical significance. Since all the components except currency and demand deposits in M1A earn non-zero interest and we observe that the interest rates paid are not negligible, which can be confirmed by the constructed own rates in Figure 1, we construct the own rate for each measure of money.

The proxy of the own rate for each monetary aggregate is constructed on a monthly basis from the own rates on the individual components[28] and is characterized as a short-term rate. The weights are the share of the components in monetary aggregates during the previous month. The own rate on the individual component is from the Financial Statistics Monthly (FSM), listed as "Rates of Banks". The quarterly rates are obtained from the arithmetic average of monthly observations and we denote them i^{a}, i^{b}, i^{bp}, and i^{m2}.

(2) Unorganized money market rate

The importance of the role played by the unorganized money market (a curb market) in the financial structure in Taiwan has been emphasized by several studies in the Taiwanese economic literature (e.g. Ho 1981, Shea 1979a,

[28] The own rate for time deposits is the average of one- three and six- month "Rates of Banks" listed in FSM.

1979b, 1983, 1986, 1991, Liang and Chen 1985, Yang 1984). As pointed out by Wijnbergen in a series of studies for South Korea (1982, 1983a, 1983b, and 1985), the typical financial structure for developing countries is characterized by undeveloped primary securities markets and the existence of the unorganized money market (U.M.M.). This is also observed in Taiwan. Appendix 2 exhibits the source of funds borrowed by private enterprises during 1976-1988 in Taiwan.

Due to severely rationed bank credit, the public lends directly to firms via the unorganized money markets, bypassing the banking system. In Taiwan, the private enterprises actively borrow from the U.M.M.[29] despite the existence of the organized money market (established in 1976) and substantial reforms launched to modernize the Taiwanese financial system in 1980s (see Semkow 1992, 1994). These loans made to the firms are supplied by friends, relatives, and professional money lenders. When looking at the balance sheet of Taiwanese people (private enterprises and individuals), we observe two major monetary assets money (currency and bank deposits) and the loans outstanding at

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[29] Postdated checks are one example of the U.M.M. instruments. They are very popular means of raising funds in Taiwan, especially among small firms. Invoices for goods are stated on the basis of cash payment, but are negotiated into 60 to 90 days postdated check payments, often including rate of interest for the supplier of the purchaser's loan. The rate is about three times that of money market rate and bank rates, (rate on postdated check = 22.44%, 1-month commercial paper rate = 7.86%, October 1997, FSM November 1997).

UMM appearing on the sheet[30] so that money and loans extended at U.M.M. are gross substitutes, obeying the adding up constraint. On a priori we have reason to regard the U.M.M. rate as the relevant opportunity cost of holding money in Taiwan before testing its significance.

Finding a statistically significant negative effect of the U.M.M. rate has an important policy message. Wijnbergen (1982, 1983a,b, 1985), by incorporating facts about the financial sector of LDCs (absence of security markets, existence of a curb market), analyzes credit policy in a macro-model in which the money demand function and the aggregate supply function are negatively related to the U.M.M. rate. He uses the existence of the curb market to explain the empirical observation that monetary policy in the developing countries has effects that differ from what we would expect given the predictions of the standard macro-model. For example, tight credit policies may lead to higher inflation and less output (stagflation) in the short run

[30] It is not argued that there are not other paper assets (e.g. bonds and equities) appearing on the balance sheet of Taiwanese households and firms. The assumption is made on observations, for example, that capital markets in Taiwan are still in the infancy (see Semkow 1994) and most treasury bills are bought and held by commercial banks as liquidity assets (Ch.8, Semkow 1992). So those paper assets may comprise only small portion in their portfolio. Of course, we may observe more varieties of assets held in their portfolio because liberation of financial markets and financial renovation since the 1980s provide Taiwanese people with more financial opportunities.

rather than an lower inflation and less output as traditionally assumed because of additional transmission channel of monetary policy into the supply side of the economy via the real costs of working capital ("Cavallo-effect") other than demand side transmission channel of monetary policy. If this is the case, policy makers should be aware of the adverse effects of the monetary policy documented in these papers and select appropriate measures to offset these short-term macro-economic consequences.

The U.M.M. rates are survey data[31]. The survey was initially conducted by the Bank of Taiwan until 1970:3. The data are available monthly from September 1961 in the Financial Statistics Monthly. In our study we use the "Unsecured Rates in Taipei City" recorded in FSM. However, the series appears inconsistent before and after 1970:3 due to changes in the way of surveying. In order to remove the inconsistency and obtain the longest series for this variable[32], the two series obtained in FSM: series 1 (1961:9-1971:2) and series 2 (1970:3-1997:9) are chained at March 1970 by first multiplying 1961:9-1970:2 data by the ratio of the average of series 2 to the average of series 1

[31] The earliest records on the U.M.M. rate dates back to May 1947. See Yuan (1986) for detailed research on the unorganized money market rate in Taiwan.

[32] Professor Rasche suggested that it was problematic to obtain credible statistical results if the sample period

during 1970:3-1971:2 and then this adjusted series is linked with the series 2 to obtain the entire series. The quarterly data are arithmetic average of the monthly data, denoted umm.

All the data series are seasonally unadjusted. The four real monetary aggregates and real GNP are in logarithms. We plot these series and the spread between the umm rate and the own rate of money (umm-i) in Figure 1. The annual growth rates for y, RM1A, RM1B, RM1BP, RM2 and inflation rates (computed from GNP deflator) are presented in Figure 2.

Figure 1 suggests that the own rate and the interest spread for each aggregate display similar trend. They peak in 1973 and 1979 when two oil crises occurred and inflation rate climbed up to double-digit. Taiwan does not have a particular high inflation rate over the whole sample period except the episode of two oil price shocks. On average, inflation rate is 5.30% annually. As Figure 2 suggests that real GNP grows at an average annual rate of 8.75% over the sample period and displays the mid-70's and early 80's recessions common to developed countries. The annul rates of growth for four real monetary aggregates have been very similar to one another, both on a trend and on a cyclical basis; with the two oil price shocks, four real monetary

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is too short, especially in conducting cointegration aggregates grow at a negative rate. The average annual growth rates for four measures of money are RM1A 11.59%, RM1B 14.05%, RM1BP 14.62%, and 15.48% for RM2.

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analysis.

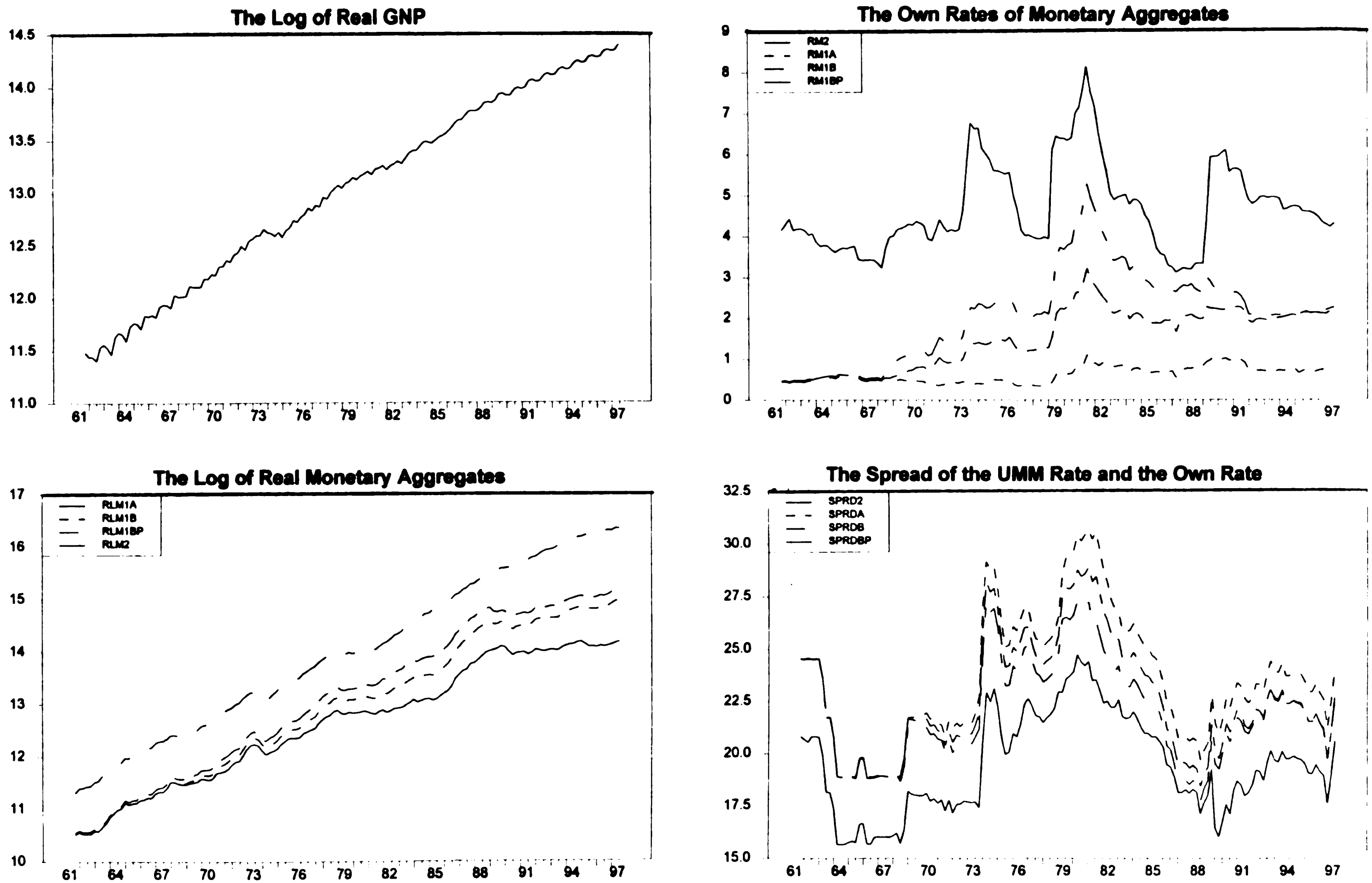

Figure 1: Taiwanese money demand data

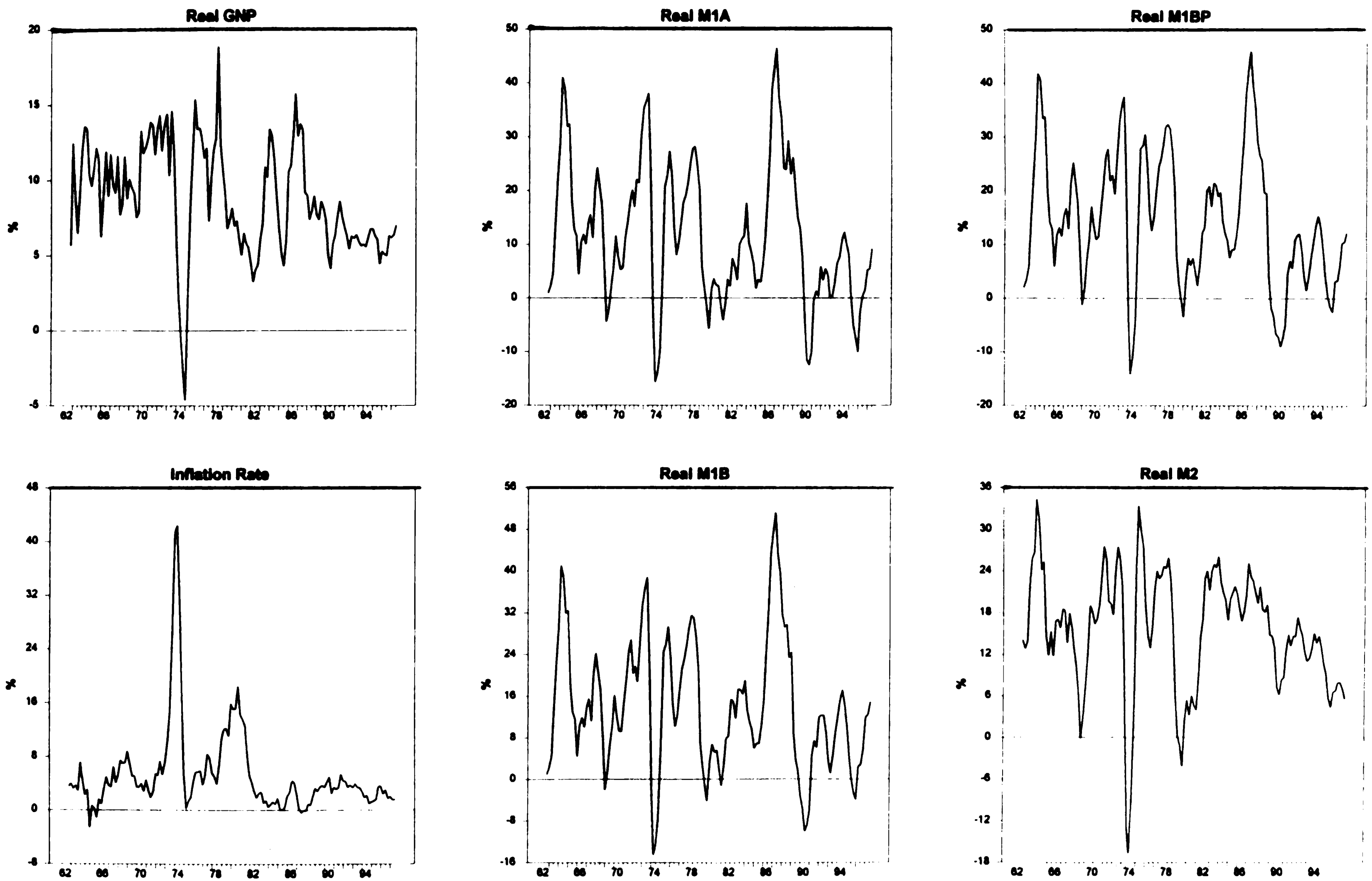

Figure 2: Growth rates of Taiwanese money demand data: 1961:4-1997:3

CHAPTER 3

STATISTICAL PROPERTIES OF UNIVATIATE TIME SERIES IN THE DEMAND FUNCTIONS FOR REAL BALNACES IN TAIWAN: 1961:4-1997:3

I: INTRODUCTION

The rapid development of time series analysis of models with unit roots has had a major impact on our understanding of the response of economic systems to shocks.

It is common practice in macroeconomics to decompose real variables such as output into a growth component, which is associated with real factors such as capital accumulation, population growth, and technological change, and a cyclical component. According to the conventional view of the business cycle, for example, Monetarist (e.g. Friedman 1974) and neo-Keynesian theories, fluctuations in output are assumed to be driven primarily by shocks to aggregate demand, such as monetary policy, fiscal policy, or animal spirits, and shocks to aggregate demand are assumed to have only a temporary effect on output; in the long-run the economy returns to the natural rate. The time series of real output has a tendency to return to a deterministic path, or to revert toward a trend, following a shock. Since transitory fluctuations are assumed to dissipate over time, a decline in real GNP below trend today has no effect on forecasts of the level of real GNP in the far future. This class of stationary processes is called trend-stationary processes. Many empirical examinations on real GNP using the United States data, for example, Nelson and Plosser (1982) and Campbell and Mankiw (1987), find a great deal persistence in real GNP and are skeptical about this implication, suggesting that shocks to GNP are largely permanent. Instead of modeling the growth component as a deterministic trend and attributing all variations in output changes to the cyclical component, a series whose fluctuations are partly temporary and partly permanent can be modeled as a combination of a stationary series and a random walk (stochastic trend, Beveridge and Nelson 1981). The random walk carries the permanent part of a change and the stationary component carries transitory part of a change[1]. Thus, real business cycle models (Kydland and

[1] Any first-difference stationary process can be decomposed as the sum of stationary(x_t^s) and random walk component(x_t^p): Let $\Delta x_t = u + c(L)\varepsilon_t$, where the roots of c(L) lie outside the unit circle; $x_t = x_t^s + x_t^p$. There are two kinds of decompositions commonly used in the literature:
(1) The Beveridge and Nelson (1981) decompositions: $x_t = x_t^s + x_t^p$, where $x_t^p = \mu + x_{t-1}^p + c(1)\varepsilon_t$ and $x_t^s = d(L)\varepsilon_t$, $d_t = -\sum_{r=t+1}^{\infty} c_j$. Note the innovations in x_t^s and x_t^p are identical so they are perfectly correlated.
(2) An unobserved components model: $x_t = x_t^s + x_t^p$, where $x_t^p = \mu + x_{t-1}^p + \eta_t$ and $x_t^s = b(L)\delta_t$, $E(\eta_t\delta_t)$ arbitrary. See Nelson

Prescott 1982), which attribute a major role to supply shocks in fluctuations of real output, and the traditional view of demand shocks, can coexist to explain the business cycle. In addition, if the output is characterized as an integrated stochastic process, we may question the natural rate hypothesis; aggregate demand shocks may have permanent effects of the level of output as described by models of multiple equilibria (e.g. Diamond 1984).

From a forecasting perspective, the implication of integrated series is that if fluctuations in output are dominated by the permanent components, the series has a tendency to move farther away from any given initial state as time goes on. Thus a decline in output today lowers forecasts of output into the infinite future.

From a purely statistical viewpoint, many statistical procedures rely critically on whether the series is integrated or stationary (including trend stationarity). The presence of a unit root in its autoregressive representation produces serious problems of statistical inference. Granger and Newbold (1974) find that two independent integrated series can display spurious correlation, characterized by high coefficient of multiple correlation R^2 and an extremely low value for the Durbin-Watson statistic in the

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and Plosser (1982), Cochrane (1987), and Stock and Watson (1988). Note the variance of the random walk is the same for regression. Phillips (1986) provides asymptotic theory for spurious regressions that relates general integrated random processes. He demonstrates that the usual t-ratio and F significance test do not possess a limiting distribution, but diverge as the sample size becomes large. The use of conventional asymptotics in setting the critical values of these tests leads to the rejection of no relationship at a rate increases with the sample size.

In order to avoid the spurious regressions and non-standard limiting distribution of coefficients in the regressions which involve integrated variables and to choose appropriate tests and estimation procedures to study the relationships among the variables, we first need to carefully investigate the nature of the individual time series.

II: UNIVARIATE UNIT-ROOT TESTS

A standard procedure of detecting unit roots in a data series is to apply Dickey-Fuller type univariate unit root tests (Fuller 1976, Dickey and Fuller 1979, 1981), or Phillips (1987), Phillips and Perron (1988) unit-root tests. Since all of our data series in the study are seasonally

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any decomposition of a unit root process into stationary unadjusted, they exhibit substantial seasonality in the levels, especially in real GNP, which can be seen in Figure 1. We should account for the effects of seasonality in the unit root tests.

In the literature, three classes of time-series models are commonly used to model seasonality: (1) purely deterministic seasonal processes, (2) stationary seasonal processes, and (3) integrated seasonality process (see Hylleberg, Engle and Granger, Yoo, 1990). The use of seasonal dummies variables, as in Barsky and Miron (1989), is not appropriate if the observed seasonality is generated by an integrated process. When applying seasonal difference filter, for example, for quarterly data, $(1-L^4)$, as advocated by Box and Jenkins (1970), analysts implicitly assume that there are seasonal unit root roots. Hence without careful detection of the presence of seasonal unit roots, the mechanical application of the seasonal difference filter is likely to produce series misspecification. Furthermore, Dickey-Fuller type tests assume that the root of interest not only has a modulus of one but is precisely one. Such a root corresponds to a zero-frequency peak in the spectrum. In addition, Dickey-Fuller tests assume that there are no other unit roots in the system. The properties of Dickey-fuller tests in the presence of unit roots at

components and random walk components (Cochrane 1987).

frequency other than zero have been investigated by Ghysels, Lee and Noh (1994). They show Dickey-fuller t-statistics remain valid even when seasonal unit-roots are present if at least three lagged dependent variables are included in the usual regression model. However, the normalized-bias statistic should be divided by four to have the same limiting distribution as that in Dickey-Fuller (1979). To see this, let x_t be a univariate stochastic process:

$$x_t = \alpha x_{t-4} + u_t, \quad t = 1, 2, \cdots, T \tag{2.1}$$

where u_t is either a stationary process with zero mean and constant variance or, alternatively, depending on the context and test u_t is a martingale difference sequence following the regularity conditions appearing in Phillips (1987) and Chan and Wei (1988). We can rewrite the time series in (2.1) as follows:

$$\Delta x_t = \phi_1 x_{t-1} + \phi_2 \Delta x_{t-1} + \phi_3 \Delta x_{t-2} + \phi_4 \Delta x_{t-3} + u_t$$

where $\phi_1 = \alpha - 1$ and $\phi_2 = \phi_3 = \phi_4 = -\alpha$.

Under the hypothesis that $\alpha = 1$ (or $\phi_1 = 0$) Ghysels, Lee, and Noh (1994) show the following relations holds:

$$(T/4)\hat{\phi}_1 \rightarrow \{W(r)^2 - 1\} / 2\int_0^1 W(r)\,dr$$

$$t_{\phi_1} \rightarrow \{W(r)^2 - 1\} / 2[\int_0^1 W(r)^2\,dr]^{1/2}$$

where $W(r)$ denotes a standard Brownian motion on [0,1].

The investigation of seasonal unit roots logically precedes the examination of other kinds of seasonality since tests can produce spurious results if seasonal unit roots are present but not accounted for. Our investigation of seasonal unit roots is conducted using the approach developed by Hylleberg, Engle, Granger and Yoo (HEGY)(1990). This is a general procedure that allows tests for unit roots at some seasonal frequencies in processes that may also exhibit deterministic or stationary stochastic seasonality without maintaining that unit roots are present at all seasonal frequencies.

1 The HEGY Seasonal Unit-Root Tests[2]

Suppose x_t is the series of interest and generated by a general autoregression of the form:

$$\varphi(L)x_t = \varepsilon_t$$

where $\varphi(L)$ is a polynomial in the lag operator and ε_t is a white noise process. Let γ_k be the roots of the characteristic polynomial associated with $\varphi(L)$. Assume for the moment that deterministic terms, such as seasonal dummies or time trend, are known to be absent from the process for x_t. In general, some or all of the γ_k may be complex.

The frequency associated with a particular root is the value of ω in $e^{i\omega}$, the polar representation of the root. A root is seasonal if $\omega = 2\pi j/S$, $j = 1,\ldots\ldots,S-1$, where S is the number of observations per year. For quarterly data, the seasonal unit roots are -1, i, $-i$, which correspond to 2 cycles per year and one cycle per year. The last root, $-i$,

[2] I wish to thank Professor Schmidt for providing me some related references regarding the effect of seasonal filter on unit-root tests which lead me to find the appropriate procedures to implement unit-root tests when using seasonally unadjusted data series. I am totally responsible for all the mistakes which I may possibly make in my study.

is indistinguishable from the one at i with quarterly data (the aliasing phenomenon) and is therefore also interpreted as the annual cycle.

To test the hypothesis that the roots of $\varphi(L)$ lie on the unit circle against the alternative that they lie outside the unit circle, the testing procedure developed by HEGY (1990) consists of linearizing the polynomial $\varphi(L)$ around the zero frequency unit root (1) plus three seasonal unit roots $(-1, i, -i)$. Thus, write $\varphi(L)$ as

$$\varphi(L) = \sum_{k=1}^{4} \lambda_k \Delta\varphi(\mathrm{L}) \frac{1-\delta_k(L)}{\delta_k(L)} + \Delta(L)\varphi^{*}(L) \qquad (2.2)$$

where $\delta_k(L) = 1 - \frac{1}{\theta_k} L$, $\lambda_k = \frac{\varphi(\theta_k)}{\prod_{j\neq k}\delta_j(\theta_k)}$, $\Delta\varphi(L) = \prod_{k=1}^{4}\delta_k(L)$.

$\varphi^{*}(L)$ is a remainder with roots outside the unit circle with $\varphi^{*}(0) = 1$; the θ_k are the unit roots $(1, -1, i, -i)$. Thus, (2.2) becomes,

$$\begin{aligned}\varphi(L) = {} & \lambda_1 L(1+L)(1+L^2) + \lambda_2(-L)(1-L)(1+L^2) \\ & + \lambda_3(-iL)(1-L)(1+L)(1-iL) \\ & + \lambda_4(iL)(1-L)(1+L)(1+iL) \\ & + \varphi^{*}(L)(1-L^4) \qquad (2.3)\end{aligned}$$

It is clear that the polynomial $\varphi(L)$ will have a root at θ_k if and only if $\lambda_k = 0$. Thus testing for unit roots can be carried out equivalently by testing for parameters $\lambda_k = 0$. Since $\varphi(L)$ is real, λ_3 and λ_4 must be complex conjugates. Let $\pi_1 = -\lambda_1$, $\pi_2 = -\lambda_2$, $2\lambda_3 = -\pi_3 + i\pi_4$, and $2\lambda_4 = -\pi_3 - i\pi_4$ After some algebraic manipulation, (2.3) can be re-written as

$$\varphi^{*}(L)\, y_{4t} = \pi_1 y_{1t-1} + \pi_2 y_{2t-1} + \pi_3 y_{3t-2} + \pi_4 y_{3t-1} + \varepsilon_t \qquad (2.4)$$

where $y_{1t} = (1+L+L^2+L^3)\,x_t$ [3]

$$y_{2t} = -(1-L+L^2-L^3)\,x_t$$

$$y_{3t} = -(1-L^2)\,x_t$$

$$y_{4t} = (1-L^4)\,x_t = \Delta_4\, x_t$$

which renders a testable regression.

(2.4) can be estimated by ordinary least squares, possibly with additional lags of y_{4t} to whiten the errors. For the root 1 this is simply a test for $\pi_1 = 0$ by use of the 't' value on π_1 and for -1 by using 't' value to test $\pi_2 = 0$. These two t values are distributed as the usual

[3] In y_{1t}, we have removed the seasonal unit roots at frequency $\pi/2$, and π and preserved the zero frequency unit root, while y_{2t} and y_{3t} contain the π and $\pi/2$ frequencies unit root but not the unit roots at other frequencies.

Dickey-Fuller t and the tables supplied by Fuller (1976) can be used. For the complex roots, λ_3 will have absolute value of zero only if both π_3 and π_4 equal zero, under $\pi_4 = 0$, the t value on π_3 has a distribution like the one developed by Dickey, Hasza, and Fuller (1984) for a process

$$x_t = \alpha x_{t-2} + \varepsilon_t .$$

The proposed HEGY t-statistics for π_1, π_2, and π_3 if (π_4 = 0) will be denoted t_1 through t_4.

In order to obtain power against relevant alternatives, the hypothesis tests can be extended to the case where the alternative includes a constant, and seasonal dummies, and/or a time trend[4]. The equation is still estimated by OLS, but the asymptotic and finite sample distributions change[5]. Thus, the HEGY tests involve running a different regression,

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[4] From his simulation results, Ghysels et al (1994) suggests when the data-generating processes have deterministic seasonal components, the regression without seasonal dummies leads to a large size distortion and too low power. In empirical applications, inclusion of deterministic terms (i.e. a constant, seasonal dummies, and/or a trend (possible irrelevant) in the model, which tends to reduce the power of the tests, is the safe strategy (pp.436).

[5] As HEGY (1990) noted, the intercept and trend portions of the deterministic mean influence only the

$$\Delta_4 x_t = \pi_1 y_{1t-1} + \pi_2 y_{2t-1} + \pi_3 y_{3t-2} + \pi_4 y_{3t-1}$$

$$+ \sum_{j=1}^{k} c_j \Delta_4 x_{t-j} + \text{[set of fixed regressors]} + \varepsilon_t$$

where the set of fixed regressors include a constant, seasonal dummies, and a linear trend. We consider four alternatives and denote them as I (with an intercept), ISD (with an intercept and seasonal dummies), IT (with an intercept and a linear trend), and ISDT (with an intercept, seasonal dummies and time trend). Since autocorrelation destroys the properties of the test, the correct augmentation in the regression is crucial to the interpretation of test results[6]. The lag length k in the augmented regression is selected on the basis of a t-test following a procedure suggested by Perron and Vogelsang (1992). k is chosen such that the coefficient on the last included lag of the fourth-differences of the data is

distribution of π_1 because they have all their spectral mass at zero frequency. Once the intercept is included, the remaining three seasonal dummies do not affect the limiting distribution. Beaulieu and Miron (1993) explicitly derive the distribution of the HEGY t-statistics for monthly data when the regression includes deterministic terms. They show t_1 is invariant to the inclusion of seasonal dummies as long as a constant is included; the distributions of $t_2 \ldots\ldots t_{12}$ are independent of constant and trend terms.

[6] If the number of lags is above/below the number necessary to render the errors white noise, the test power and size are subject to bias; that is, too many lags reduce

significant at 0.05 level (that is, t-statistic of the OLS estmator $\hat{c}_k$ is greater than 1.6 in absolute value) and the coefficient on the last included lag in higher order autoregressions is insignificant up to some a priori specified maximum order (k_{max}). Initially, we set $k_{max} = 12$; if the last lag (12) is still significant, $k_{max} = 16$.

In Table 1, we present the results of applying the HEGY test procedure to the variables in the Taiwanese demand function for real balances. The quarterly series are (y), four monetary aggregates deflated by GNP deflator (RM1A, RM1B, RM1BP, and RM2), the interest rates (i^a, i^b, i^{bp}, i^{m2}, umm) and the spreads between the own rate and umm rate (umm-i). All variables are in log level except the interest rates. The sample period is 1961:4 to 1997:3.

From Table 1, the p-values for Q statistic suggest the residuals of the regression are white noise. The results strongly indicate that there is a unit root at the zero frequency in all of the series except RM2 since we cannot reject the null hypothesis in any of the auxiliary regressions at the 5% level. The assumption of $\pi_1 = 0$ is rejected marginally for RM2 in the regression (IT) at the 0.05 level but not at the 0.10 level.

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the power and too few lags distorts the test size. See Hyllberg (1995).

Next we consider a unit root at $\pi/2$. In order to establish a unit at $\pi/2$, we need both $\pi_3 = 0$ and $\pi_4 = 0$. Since we fail to reject both hypothesis simultaneously for each series at either 5% or 10% level, we reject the existence of a root at this frequency for all series. Moreover, the tests reject that there are any seasonal unit roots in all of the interest rates except i^a because t_2, t_3, and t_4 are all below the 5% critical values in four models.

At π, we cannot reject $\pi_2 = 0$ for y and RM2 in ISD and ISDT at either 10% or 5% level. For i^a, $\pi_2 = 0$ cannot be rejected at the 5% level in all models.

To summarize, we fail to reject the hypothesis of a unit root at frequency zero for all series. The data reject the presence of any seasonal unit roots in all of the interest rates except i^a. Series y, RM2, and i^a may contain a unit root at the biannual frequency (π).

Table 1
HEGY tests for seasonal unit roots in quarterly aggregate series for the Taiwanese demand functions for real balances: 1961:4 - 1997:3

Model I: Regression with an Intercept

		0	π	$\pi/2$		Q^a
Series	Lags(k)	t_1	T_2	t_3	t_4	p-value
y	5	-2.46	-1.79*	-2.54**	-1.96**	0.22
RM1A	7	-1.36	-2.77**	-0.60	-2.79**	0.34
RM1B	7	-1.15	-3.11**	-0.32	-3.00**	0.57
RM1BP	7	-1.61	-3.00**	-0.33	-3.24**	0.76
RM2	6	-0.98	-1.86*	-1.84*	-3.20**	0.75
umm-i^a	0	-2.13	-8.32**	-4.46**	-7.40**	0.66
umm-i^b	0	-2.32	-8.41**	-4.70**	-7.14**	0.66
umm-i^{br}	0	-2.56	-8.51**	-4.89**	-6.87**	0.61
umm-i^{m2}	0	-1.92	-1.97**	-5.47**	-6.35**	0.98
i^a	9	-1.49	-1.60*	-0.58	-2.64**	0.99
i^b	9	-1.31	-2.66**	-1.11**	-3.05**	0.98
i^{br}	6	-1.80	-3.42**	-1.80*	-2.79**	0.52
i^{m2}	2	-2.88	-5.69**	-3.31**	-0.06	0.93
umm	1	-2.06	-6.66**	-3.75**	-6.90**	0.93
Critical Values[b]						
5%	**	-2.89	-1.91	-1.88	-1.68	
10%	*	-2.58	-1.58	-1.53	-1.31	

An * and ** indicates significance at the 10% and 5% levels, respectively.
(a) This is the marginal significance level of the Ljung-Box Q statistic (1979) for the test of no serial correlation in the residuals of the regression.
(b) Critical values are from Hylleberg, Engle, Granger, and Yoo (1990).

Table 1 (cont'd)
Model ISD: Regression with an intercept and seasonal dummies

		0	π	$\pi/2$		Q
Series	Lags(k)	t_1	T_2	t_3	t_4	p-value
y	5	-2.45	-1.25	-2.15	-2.77**	0.16
RM1A	5	-1.55	-2.72*	-1.45	-4.73**	0.21
RM1B	7	-1.08	-3.04**	-0.53	-3.86**	0.65
RM1BP	7	-1.54	-3.10**	-0.47	-3.91**	0.78
RM2	6	-0.94	-2.17	-1.79	-3.33**	0.58
umm-i^{n}	0	-2.12	-7.95**	-4.37**	-7.44**	0.77
umm-i^{h}	0	-2.33	-7.98**	-4.63**	-7.21**	0.78
umm-i^{hr}	0	-2.57	-7.97**	-4.88**	-6.98**	0.76
umm-i^{m2}	0	-1.95	-7.59**	-5.49**	-6.41**	0.99
i^{n}	9	-1.38	-2.84*	-1.35	-4.01**	0.99
i^{h}	1	-1.41	-6.54**	-4.15**	-7.68**	0.20
i^{hr}	6	-1.75	-3.58**	-1.99	-3.20**	0.42
i^{m2}	1	-2.84*	-6.54**	-3.64**	-7.92**	0.84
umm	1	-1.99	-6.66**	-3.87**	-7.06**	0.81
Critical values						
5%	**	-2.94	-2.90	-3.44	-1.96	
10%	*	-2.62	-2.59	-3.11	-1.52	

Table 1 (cont'd)
Model IT: Regression with an intercept and a linear trend

		0	π	$\pi/2$		Q
Series	Lags(k)	t_1	T_2	t_3	t_4	p-value
y	5	-1.15	-1.78*	-2.53**	-1.92**	0.34
RM1A	7	-1.73	-2.72**	-0.65	-2.82**	0.37
RM1B	7	-1.66	-3.03**	-0.39	-3.03**	0.59
RM1BP	7	-0.61	-2.97**	-0.34	-3.24**	0.77
RM2	5	-3.48**	-3.47**	-1.96**	-1.56*	0.52
umm-i^a	0	-2.26	-8.30**	-4.49**	-7.38**	0.64
umm-i^b	0	-2.35	-8.37**	-4.70**	-7.12**	0.65
umm-i^{br}	0	-2.59	-8.46**	-4.89**	-6.85**	0.61
umm-i^{m2}	0	-2.11	-7.94**	-5.49**	-6.33**	0.90
i^a	9	-2.63	-1.57	-0.57	-2.57**	0.99
i^b	6	-2.49	-3.56**	-2.23**	-3.15**	0.92
i^{br}	6	-1.44	-3.41**	-1.78*	-2.78**	0.53
i^{m2}	1	-2.98	-6.44**	-3.53**	-7.69**	0.93
umm	1	-2.24	-6.60**	-3.76**	-6.86**	0.72
Critical values						
1%		-4.09				
5%	**	-3.46	-1.96	-1.90	-1.64	
10%	*	-3.16	-1.63	-1.52	-1.23	

Table 1 (cont'd)
Model ISDT: Regression with an intercept, seasonal dummies and a linear trend

		0	$\pi/2$	$\pi/2$		Q
Series	Lags(k)	t_1	T_2	t_3	t_4	p-value
y	5	-1.08	-1.24	-2.15	-2.73**	0.26
RM1A	5	-2.47	-2.64*	-1.49	-4.45**	0.26
RM1B	5	-2.57	-2.72*	-1.56	-4.52**	0.16
RM1BP	5	-1.58	-2.83*	-1.45	-4.88**	0.15
RM2	3	-2.72	-2.52	-3.73**	-6.00**	0.20
umm-i^{n}	0	-2.27	-7.92**	-4.41**	-7.43**	0.75
umm-i^{b}	0	-2.37	-7.95**	-4.64**	-7.20**	0.77
umm-i^{br}	0	-2.61	-7.92**	-4.88**	-6.97**	0.76
umm-i^{m2}	0	-2.16	-7.56**	-5.52**	-6.41**	0.99
i^{n}	9	-2.36	-2.78*	-1.33	-3.86**	0.99
i^{b}	1	-1.98	-6.40**	-4.16**	-7.34**	0.26
i^{br}	6	-1.39	-3.56**	-1.98**	-3.19**	0.43
i^{m2}	1	-2.88	-6.51**	-3.62**	-7.83**	0.84
umm	1	-2.18	-6.61**	-3.89**	-7.01**	0.80
Critical values						
5%	**	-3.52	-2.93	-3.44	-1.94	
10%	*	-3.21	-2.61	-3.12	-1.51	

2 The Dickey-Fuller Unit-Root Tests

Even the presence of the seasonal unit roots, we can still apply the Dickey-Fuller tests to test for a unit root at zero frequency. Table 2 presents the values of test statistics of the DF-t test and normalized bias test. We compute DF statistics based on

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

where set of fixed regressors are the same as before; k is chosen according the same procedure as described in the HEGY test. A t-statistic is based on $\hat{\alpha}$, the OLS estimator; Tc($\hat{\alpha}$ - 1) is the normalized bias test, where T is the number of the observations, c = $(1-\sum_{j=1}^{k}\hat{\alpha}_j)^{-1}$. As we mentioned earlier, in the case of the presence of the seasonal unit root, k should be no less than three and the normalized bias statistic should be divided by four. From the HEGY tests, we observe no seasonal unit roots in any of the interest rates except i^{u}, so k is not restricted for the interest rates.

From Table 2, the results unambiguously indicate that y, RM1B, RMBP, and all interest rates except i^{u} and i^{m2} contain a unit root, since none of the tests can reject the null hypothesis of a unit root. The null is rejected at the 5% level by normalized bias tests $Tc(\hat{\alpha} - 1)$, for RM1A in ISDT, for RM2 in IT, ISDT, for i^{a} in I, IT, and for i^{m2} in ISD but not by the statistic for the presence of seasonal unit root, $\frac{Tc(\hat{\alpha}-1)}{4}$. Thus, the outcome depends on whether these two series have a unit root at frequency π (since we have ruled out $\pi/2$, see above).

From the HEGY test, the t_2 test cannot reject a unit root at π at the 5% level for RM1A in ISDT, RM2 in ISDT, and i^{a} in IT so we may use $\frac{Tc(\hat{\alpha}-1)}{4}$ for inference. Hence, the null hypothesis of a unit root is rejected for these series in the corresponding models.

Table 2

DF tests for unit roots at zero frequency in quarterly aggregate series for the Taiwanese demand functions for real balances: 1961:4 - 1997:3

Model I: Regression with an intercept

Series	Lags (k)	$\hat{\alpha}$ [a]	t-test	Tc($\hat{\alpha}$-1)	$\frac{Tc(\hat{\alpha}-1)}{4}$	Q[b] p-value
y	9	0.996	-2.20	- 0.47	-0.12	0.25
RM1A	10	0.995	-1.29	- 0.85	-0.21	0.41
RM1B	11	0.997	-1.12	- 0.49	-0.12	0.56
RM1BP	11	0.996	-1.56	- 0.63	-0.16	0.75
RM2	11	0.999	-0.74	- 0.21	-0.05	0.98
umm-i^{a}	2	0.961	-1.95	- 8.09	-2.02	0.65
umm-i^{b}	2	0.954	-2.11	- 9.19	-2.30	0.62
umm-i^{br}	2	0.944	-2.32	-10.86	-2.71	0.54
umm-i^{m2}	0	0.958	-1.76	- 6.07	-1.52	0.80
i^{a}	9	0.936	-2.36	-23.3**	-5.84	0.37
i^{b}	12	0.982	-1.31	-2.67	-0.67	0.98
i^{br}	10	0.978	-1.77	-5.89	-1.47	0.53
i^{m2}	1	0.928	-2.92	-10.20	-2.55	0.92
umm	2	0.961	-1.96	-8.52	-2.13	0.77
Critical value[c]						
5%	**		-2.89	-13.70	-13.70	

An ** indicates significance at 5% level.

(a) $\hat{\alpha}$ is the OLS estimate of the parameter (α) in the autoregressive representation of the variable, $x_t = \alpha x_{t-1} + \sum_{i=1}^{k} \alpha_i \Delta x_{t-i}$ + [set of fixed regressors] + ε_t.

(b) This is the marginal significance level of the Ljung-Box Q statistic (1978) for the test of no serial correlation in the residuals of the regression.

(c) Critical values are taken from Ghysels, Lee, and Noh (1994).

Table 2 (cont'd)
Model ISD: Regression with an intercept and seasonal dummies

Series	Lags (k)	$\hat{\alpha}$	t-test	$Tc(\hat{\alpha}-1)$	$\frac{Tc(\hat{\alpha}-1)}{4}$	Q p-value
y	12	0.995	-2.11	- 0.46	-0.12	0.21
RM1A	9	0.996	-1.40	- 1.04	-0.26	0.21
RM1B	11	0.999	-1.09	- 0.47	-0.12	0.61
RM1BP	11	0.996	-1.56	- 0.61	-0.15	0.76
RM2	11	0.999	-0.73	- 0.21	-0.05	0.94
umm-i^{a}	2	0.961	-1.98	- 8.59	-2.15	0.79
umm-i^{b}	2	0.954	-2.15	- 9.86	-2.47	0.78
umm-i^{br}	2	0.943	-2.36	-11.77	-2.94	0.71
umm-i^{m2}	0	0.960	-1.69	- 5.78	-1.45	0.93
i^{a}	12	0.965	-1.38	-5.16	-1.29	0.99
i^{b}	6	0.982	-1.50	-4.51	-1.13	0.70
i^{br}	10	0.979	-1.76	-5.81	-1.45	0.42
i^{m2}	2	0.927	-2.86	-17.6**	-4.35	0.83
umm	2	0.961	-1.98	-8.64	-2.16	0.83
Critical value						
5%	**		-2.89	-13.70	-13.70	

Table 2 (cont'd)

Model IT: Regression with an intercept and a linear trend

Series	Lags (k)	$\hat{\alpha}$	t-test	$Tc(\hat{\alpha}-1)$	$\frac{Tc(\hat{\alpha}-1)}{4}$	Q p-value
y	9	0.985	-0.52	- 1.61	- 0.40	0.29
RM1A	11	0.947	-1.90	-20.06	- 5.01	0.47
RM1B	11	0.948	-1.66	-14.89	- 3.72	0.62
RM1BP	11	0.987	-0.54	- 2.40	- 0.60	0.77
RM2	11	0.917	-2.22	-36.08**	- 9.02	0.97
umm-i^{a}	2	0.958	-2.08	- 8.92	- 2.23	0.64
umm-i^{h}	2	0.954	-2.14	- 9.40	- 2.35	0.61
umm-i^{hr}	2	0.943	-2.34	-11.06	- 2.76	0.54
umm-i^{m2}	0	0.953	-1.91	- 6.77	- 1.69	0.80
i^{a}	13	0.890	-2.42	-36.9**	-9.13	0.99
i^{h}	11	0.950	-1.69	-12.91	-3.23	0.99
i^{hr}	12	0.979	-1.24	-4.34	-1.09	0.84
i^{m2}	2	0.923	-2.92	-18.39	-4.60	0.92
umm	2	0.957	-2.12	-9.57	-2.39	0.76
Critical value						
5%	**		-3.45	-21.40	-21.40	

Table 2 (cont'd)
Model ISDT: Regression model contains an intercept, seasonal dummies, and a linear trend

Series	Lags (k)	$\hat{\alpha}$	t-test	$Tc(\hat{\alpha}-1)$	$\frac{Tc(\hat{\alpha}-1)}{4}$	Q p-value
y	9	0.986	-0.49	- 1.54	-0.39	0.25
RM1A	9	0.941	-2.31	-28.40**	-7.10	0.28
RM1B	11	0.950	-1.61	-14.05	-3.51	0.65
RM1BP	11	0.987	-0.52	- 2.31	-0.58	0.78
RM2	11	0.918	-2.18	-34.95**	-8.74	0.94
umm-i^{a}	2	0.958	-2.11	- 9.50	-2.38	0.79
umm-i^{b}	2	0.953	-2.18	-10.10	-2.52	0.78
umm-i^{br}	2	0.942	-2.39	-11.95	-2.99	0.72
umm-i^{m2}	0	0.955	-1.86	- 6.51	-1.63	0.93
i^{a}	12	0.909	-2.36	-21.24	-5.31	0.99
i^{b}	6	0.944	-2.32	-16.91	-4.23	0.76
i^{br}	10	0.978	-1.34	-6.02	-1.51	0.42
i^{m2}	2	0.923	-2.91	-18.43	-4.61	0.83
i^{umm}	3	0.952	-2.31	-10.27	-2.57	0.82
umm	2	0.957	-2.15	- 6.19	-1.55	0.85
Critical value						
5%	**		-3.45	-21.40	-21.40	

3 The KPSS Stationarity Tests

Several studies have argued that the Dickey-Fuller test, has low power against reasonable alternatives, for example, DeJong et al. (1989) provides evidence that the Dickey-Fuller tests have low power against a unit root near unity. Ghysels, Lee, and Noh (1994) evaluated the small sample properties of the HEGY test, finding the HEGY also has low power against the relevant alternatives. Unlike the Dickey-Fuller tests and the HEGY tests whose the null hypothesis is a unit root process, Kwiatkowski, Phillips, Schmidt, and Shin (KPSS) (1992), propose a test of the null hypothesis of stationarity (short-memory) against the alternative of a unit root. The tests complement unit root tests, such as the Dickey-fuller tests. By testing both the unit root hypothesis and the stationary hypothesis, we can distinguish series that appear to be stationary, series that appear to have a unit root, and series for which the data (or the tests) are not sufficiently informative to be sure whether they are stationary or integrated. To ensure the existence of a unit root at zero frequency, we also apply the KPSS test to the data series.

In the study we consider two null hypotheses[7]; one is the level stationarity hypothesis, the other is a trend

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[7] Canova and Hansen (1993) extend the KPSS test to the seasonal case and propose a test, the CH test, based on the residuals from a regression extracting the deterministic

statonarity. The resulting statistics are denoted η_{μ}, and η_{τ}, respectively. For each test, we will consider values of the lag truncation parameter (q) from 2 to 8[8].

The test results are presented in Table 3 and 4. First we consider the null hypothesis of stationarity around a level. From Table 3, we reject the hypothesis of level stationarity for all series except the interest rates (all the spreads, i^{m2}, and umm) no matter what value of q we choose. The outcome for the spreads and i^{m2} and umm depends on the lag truncation parameter (q). Since obvious deterministic trends are present in real GNP and four real monetary aggregates (see Figure 1), the rejection of the null is not surprising. We then proceed to test the trend stationary hypothesis. From Table 4, for all of the series

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seasonal components and other deterministic components, see Hylleberg (1995). We also consider the regression with seasonal dummies for both hypotheses, however, we do not report the results in the study because the values of the statistics are almost identical to those in the corresponding regression without seasonal dummies.

[8] Since the data series are highly dependent over time and the residuals from the regressions are serially correlated, it is not realistic to assume iid errors under the null and use q = 0, no correction for autocorrelation, in estimation of the long-run variance. The choice of eight as the maximal value of q, is based on the observation that the value of the test statistic has settled down by the time we reach q = 8 and the simulation results in KPSS (1992), which suggest q = 8 is a compromise between the large size distortion under the null for l = 4 and the very low power under the alternative for l = 12.

except RM2 the null is rejected at either 5% or 10% level for all values of q.

For the interest rates, KPSS tests of level stationarity , the ADF-t, ADF-F and HEGY-t_1 tests of a unit root against level stationarity unambiguously suggest that i^{a}, i^{b}, and i^{bp} contain a unit root at zero frequency since we can reject the trend stationary hypothesis and we cannot reject a unit root hypothesis. Similarly, combining the KPSS test of the trend stationarity and the ADF-t, ADF-F, and HEGY t_1 test of a unit root against the trend stationary alternative, we find that all the series except RM2 appear to have a unit root at zero frequency. For RM2 and interest rates (the spreads and i^{m_2}, umm), since both the null of a unit root and the null of stationarity are rejected, they may be characterized as long memory processes. In the following analysis, we will treat them as integrated processes.

In order to determine the order of integration maintained in each variables, we apply the ADF-t test to the first-difference of the series. The alternative models considered are again I, ISD, IT, and ISDT as defined above. We report the results in Table 5. From Table 5, we reject the null hypothesis of a unit root against all of the alternatives for all series at 5% level. Therefore, we model each of the variables under study in the Taiwanese demand for real balances as an I(1) process.

Table 3
KPSS stationarity tests for quarterly aggregate series for the Taiwanese demand functions for real balances (1961:4 - 1997:3)

Model I: Regression with an intercept

	Lag truncation parameter (q)						
Series	2	3	4	5	6	7	8
	η_μ: 5% critical value is 0.463 10% critical value is 0.347						
y	3.39**	2.62**	2.20**	1.96**	1.82**	1.73**	1.70**
RM1A	3.34**	2.58**	2.17**	1.91**	1.79**	1.71**	1.68**
RM1B	3.38**	2.61**	2.20**	1.94**	1.81**	1.73**	1.69**
RMBP	3.38**	2.61**	2.20**	1.96**	1.81**	1.73**	1.69**
RM2	3.39**	2.62**	2.20**	1.97**	1.82**	1.74**	1.70**
umm-i^a	0.59**	0.46**	0.39*	0.35*	0.33	0.32	0.31
umm-i^b	0.44*	0.34*	0.29	0.27	0.25	0.24	0.24
umm-i^{bp}	0.34*	0.27	0.23	0.21	0.20	0.19	0.19
umm-i^{m2}	0.63**	0.49**	0.42*	0.38*	0.35*	0.34*	0.33
i^a	1.76**	1.37**	1.17**	1.05**	0.98**	0.94**	0.92**
i^b	2.70**	2.09**	1.77**	1.58**	1.46**	1.40**	1.37**
i^{bp}	1.78**	1.38**	1.17**	1.05**	0.97**	0.93**	0.90**
i^{m2}	0.42**	0.33	0.29	0.27	0.26	0.25	0.25
umm	0.64**	0.50**	0.42*	0.38*	0.36*	0.34	0.34

Test statistics are computed using a Newey-West procedure and the Bartlett lag window as suggested in KPSS (1992).
An * and ** indicates significance at 10% and 5% level, respectively.
Critical values are taken from Kwiatkowski, Phillips, Schmidt, and Shin (1992)

Table 4
KPSS stationarity tests for quarterly aggregate series for the Taiwanese demand functions for real balances (1961:4 - 1997:3)

Model IT: Regression with an intercept and a trend

	LAG truncation parameter (q)						
Series	2	3	4	5	6	7	8
	η_τ: 5% critical value is 0.146 10% critical value is 0.119						
Y	0.62**	0.50**	0.42**	0.38**	0.36**	0.35**	0.34**
RM1A	0.31**	0.25**	0.22**	0.20**	0.19**	0.18**	0.18**
RM1B	0.26**	0.21**	0.18**	0.17**	0.16**	0.16**	0.15**
RM1BP	0.43**	0.34**	0.30**	0.27**	0.26**	0.25**	0.25**
RM2	0.11*	0.09	0.08	0.08	0.07	0.07	0.07
umm-i^{a}	0.45**	0.35**	0.30**	0.27**	0.26**	0.25**	0.24**
umm-i^{b}	0.43**	0.33**	0.29**	0.26**	0.24**	0.23**	0.23**
umm-i^{br}	0.34**	0.27**	0.23**	0.21**	0.20**	0.19**	0.19**
umm-i^{m2}	0.46**	0.36**	0.31**	0.28**	0.26**	0.25**	0.25**
i^{a}	0.24**	0.19**	0.16**	0.15**	0.14*	0.13*	0.13*
i^{b}	0.42**	0.33**	0.28**	0.26**	0.24**	0.23**	0.23**
i^{br}	0.64**	0.49**	0.42**	0.38**	0.35**	0.34**	0.33**
i^{m2}	0.23**	0.19**	0.17**	0.15**	0.15**	0.14*	0.14*
umm	0.44**	0.35**	0.30**	0.27**	0.25**	0.24**	0.24**

See notes to Table 3.

Table 5
DF-t tests for unit roots in the first difference of the aggregate series for the Taiwanese demand functions for real balances: 1961:4 - 1997:3

Regression	I(with an intercept)			ISD(with an intercept and seasonal dummies)		
	Lags (k)	t-test	Q[a] p-value	Lags (k)	t-test	Q p-value
y	7	- 3.88**	0.38	7	-3.86**	0.29
RM1A	8	- 3.33**	0.07	6	-4.37**	0.07
RM1B	10	- 4.04**	0.61	10	-4.09**	0.60
RM1BP	10	- 4.04**	0.80	10	-4.03**	0.81
RM2	10	- 3.88**	0.98	10	-3.84**	0.94
umm-i^{a}	0	- 8.20**	0.55	0	-7.90**	0.76
umm-i^{b}	0	- 8.43**	0.52	0	-8.12**	0.75
umm-i^{br}	0	- 8.67**	0.42	0	-8.36**	0.69
umm-i^{m2}	0	-10.08**	0.89	0	-9.75**	0.98
i^{a}	9	- 3.75**	0.16	10	-4.67**	0.99
i^{b}	11	- 3.82**	0.98	9	-3.53**	0.81
i^{bp}	9	- 3.63**	0.51	10	-4.33**	0.97
i^{m2}	10	- 4.00**	0.82	10	-4.04**	0.70
umm	0	- 8.21**	0.70	0	-7.98**	0.82
Critical value[b]						
5%	**	-2.89			-2.89	

See notes to Table 2.

Table 5 (cont'd)

Regression	IT(with an intercept and a trend)			ISDT(with an intercept, trend and seasonal dummies)		
	Lags (k)	t-test	Q p-value	Lags (k)	t-test	Q p-value
y	7	- 4.61**	0.20	7	-4.60**	0.15
RM1A	10	- 3.58**	0.40	9	-3.54**	0.19
RM1B	10	- 4.15**	0.56	7	-4.19**	0.60
RM1BP	10	- 4.33**	0.74	10	-4.32**	0.75
RM2	9	- 4.53**	0.98	9	-4.43**	0.93
umm-i^{a}	0	- 8.17**	0.56	0	-7.87**	0.76
umm-i^{b}	0	- 8.41**	0.52	0	-8.09**	0.75
umm-i^{bp}	0	- 8.67**	0.43	0	-8.35**	0.69
umm-i^{m2}	0	-10.06**	0.90	0	-9.73**	0.69
i^{a}	10	- 4.94**	0.99	10	-4.65**	0.99
i^{b}	11	- 3.84**	0.98	10	-4.36**	0.96
i^{bp}	9	- 3.80**	0.45	11	-3.64**	0.77
i^{m2}	10	- 4.12**	0.82	10	-4.05**	0.68
umm	0	- 8.19**	0.71	0	-7.95**	0.82
Critical values						
5%	**	-3.45		**	-3.45	

See notes to Table 2.

CHAPTER 4

THE LONG-RUN DEMAND FOR MONEY FUNCTIONS IN TAIWAN (1961:4-1997:3): COINTEGRAITON EVIDENCE

I: EQUILIBRIUM RELATIONSHIP AND COINTEGRATION

Economic theories suggest that the long-run money demand relationship can be explained by functional relationships:

$$m_t - p_t = f(y_t, R_t) \tag{3.1}$$

m, p, y, and R are defined as in (1.42). We can write a stochastic version for the long-run equilibrium relationship (3.1) as

$$m_t - p_t = f(y_t, R_t) + v_t \tag{3.2}$$

v_t are disturbances. Let X_t be a vector consisting of $(m_t - p_t, y_t, R_t)$, if $f(\cdot)$ is linear in y and R (or some transformations thereof, the equilibrium relationship in (3.1) can be stated as $\alpha' X_t = 0$. In most time periods, X_t will not be in equilibrium and the univariate quantity

$$v_t = \alpha' X_t$$

can be called the equilibrium error.

In the unit root tests, we have established that all series chosen for the demand for money functions in Taiwan are individually I(1). It means that an individual economic variable can wander extensively whenever it is hit by shocks. As far as the demand for money function is concerned, economic theories suggest that the variables in the function are expected to move so that they do not drift too far apart at least in the long-run, if not in the short run. An economic theory involving equilibrium concepts can be interpreted in terms of the concept of cointegration in econometric literature (Engle and Granger 1987). If each element of a vector series X_t first achieves stationarity after differencing, but a linear combination $\alpha' X_t$ is already stationary, the time series X_t are said to be cointegrated with cointegrating vector α. Interpreting

$$\alpha' X_t$$

as a long-run equilibrium, cointegration implies that deviations from equilibrium are stationary, with finite variance, even though the series are non-stationary and have infinite variance.

However the theories suggest, there is no prior belief that these I(1) variables in the study necessarily obey the functional form as theories predict. Whether cointegration occurs or not is an empirical question. In this chapter, we examine the empirical evidence of cointegration in the Taiwanese long-run demand for real balances for four measures of monetary aggregates (RM1A, RM1B, RM1BP, and RM2). The choice of scale variable is real GNP (y). The cost of holding money for each measure of money is the difference between the own rate of money (i^a, i^b, i^{bp}, i^{m2}) and the unorganized money market rate (***umm***). We first introduce how to test cointegration between I(1) variables.

II: RESIDUAL BASED COINTEGRATION TESTS

1 <u>Introduction</u>

Engle and Granger (1987) suggest that the residuals from OLS estimation of the cointegrating regression can be examined for the presence of a unit root in the autoregressive representation. If the I(1) series are not cointegrated then there must be a unit root in these residuals; this is therefore the null of no cointegration. If the series are cointegrated, then the residuals will be stationary.

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

$$x_{1t} = \hat{\alpha}' X_{2t} + \hat{v}_t, \tag{3.3}$$

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

$$\hat{v}_t = \hat{\rho}\hat{v}_{t-1} + \hat{\varepsilon}_t, \quad \hat{\rho} = 1, \tag{3.4}$$

that is to test $\hat{\rho} = 1$.

[1] Engle and Granger (1991) point out "in almost all cases, there should be an intercept in the cointegrating regression. Only strong priors should set it to zero. With seasonal process which have deterministic seasonals, seasonal dummy variables can be added either in the

2 *Phillips and Ouliaris (1990) Z_α and Z_t Tests*

The testing procedure of no cointegration is first to run the cointegrating regression (3.3) by OLS as described above and then apply an standard unit root test, such as the Dickey-fuller unit root test (1976 and 1979), or Phillips'(1987) Z_α and Z_t tests on the residuals of the cointegrating regression (3.4). The test statistics are calculated as in the context of univariate unit root testing, but the asymptotic distributions of the test statistics differ. Phillips and Ouliaris (1990) examine both the ADF test and Phillips' Z_α and Z_t tests under the assumption that no deterministic trends are present in the regressors or the regression equation. The tabulated distributions are also provided. Hansen (1992a) further discovers that such cointegration tests depend on the actual trends in the data if the data is not detrended. Since the distributions under the null and alternative hypotheses are sensitive to the true trend process and detrending procedures, careful consideration of the nature of the trends in the regressors must be given if tests of the cointegration hypothesis are to have the correct size and optimal power. From unit root tests, we have established all individual series contain a unit root. However many

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cointegrating regression or in the error correction form." (PP.11) variables, for example, the series (y, RM1A, RM1B, RM1BP, and RM2) appear to have deterministic trends. This can be seen from visual inspection of the graph (Figure 1). These series may be actually best thought of as "I(1) with drift", which is the sum of an I(1) process with zero-mean increments and a linear trend. In the following analysis, we will conduct tests incorporating this data property[2].

Although a simple Dickey-Fuller test would be appropriate in the model, we consider Z_a and Z_t tests, which have advantages to correct potential serial correlation and heteroskedasticity in the cointegrating errors.

While the null hypothesis of no cointegration in (3.4) holds when $\hat{\rho} = 1$, the alternative holds for $\hat{\rho} < 1$. Following Hansen (1992a), we consider two procedures. The first is to run the unrestricted OLS regression in which a time trend is included to test no cointegration of series

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[2] We employed Dickey-Fuller F tests (Dickey and Fuller 1981) of the hypothesis $(\alpha,\beta,\rho) = (0,0,1)$ and of the hypotheses $(\alpha,\beta,\rho) = (\alpha,0,1)$ in $Y_t = \alpha + \beta t + \rho Y_{t-1} + \sum_{i}^{l} \alpha_i \Delta Y_{t-i} + e_t$, where Y_t is univariate time series. The likelihood ratio test of the hypothesis that the true model is a unit root process with zero drift $(\alpha,\beta,\rho) = (0,0,1)$ is rejected at the 5% level and that $(\alpha,\beta,\rho) = (\alpha,0,1)$ is not rejected for these series (y, RM1A, RM1B, RM1BP, RM2)(the estimated α ranges from 0.17 to 0.96). Hence the formal tests suggest that the series (y, RM1A, RM1B, RM1BP, and RM2) contain a unit root with possible drift.

which have a drift[3]; this is equivalent to detrending the series first[4]:

$$x_{1t} = \hat{\mu} + \hat{\alpha}' x_{2t} + \hat{\beta} t + \hat{v}_t \qquad (3.5)$$

The second is the restricted OLS regression which excludes a time trend by constraining $\hat{\beta} = 0$.

$$x_{1t} = \tilde{\mu} + \tilde{\alpha}' x_{2t} + \tilde{v}_t ,$$

Hansen (1992a) notes that cointegration tests based on the restricted regression are more powerful than cointegration tests based on the unrestricted regression. Note when some or all of the I(1) variables are involved deterministic time trends, cointegration implies that cointegrating vectors should be sufficient to eliminate both the trend components and stochastic components of the series for the linear combination of these variables being stationary (see pp.574-575, Hamilton 1994).

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[3] Engle and Granger (1991, pp.14) note that this is a test that the series are not cointegrated even after extracting a linear time trend.

[4] The inclusion of time trends in the regression has the advantage of rendering estimates of the cointegrating vector invariant to the presence of trends in the regressors and simplifies the asymptotic theory, see Phillips and Hansen (1990), Hansen (1992a).

The Phillip's (1987) test statistics Z_a and Z_t are formed using a "bias-corrected" serial correlation coefficient estimates

$$\hat{\rho}^{*} = \frac{\sum_t (\hat{v}_t \hat{v}_{t-1} - \hat{\lambda})}{\sum_t \hat{v}_t^2}$$

and are computed as

$$Z_a = T(\hat{\rho}^{*} - 1), \text{ T: sample size} \tag{3.6}$$

$$Z_t = \frac{\hat{\rho}^{*} - 1}{\hat{s}}, \quad \hat{s}^2 = \frac{\hat{\sigma}^2}{\sum_t \hat{v}_t^2} \tag{3.7}$$

where $\hat{\lambda} = \sum_{j=1}^{T} \omega(j/M) \frac{1}{T} \sum_t \hat{\varepsilon}_{t-j} \hat{\varepsilon}_t$, $\hat{\varepsilon}_t = \hat{v}_t - \hat{\rho}\hat{v}_{t-1}$; $w(\cdot)$ is kernel weight; $\hat{\sigma}^2 = \frac{1}{T} \sum_t \hat{\varepsilon}_t^2 + 2\hat{\lambda}$.

The statistics for the restricted regression are constructed from $\tilde{v}_t$ in the same way. The long-run variance ($\hat{\sigma}^2$) is estimated using a prewhitened quadratic spectral kernel (QS)[5] with a first-order autoregression for the

[5] Any kernel that yields positive semi-definite estimates can be used. These include the Bartlett, Parzen,

prewhitening[6] and an automatic bandwidth estimator[7] (see Andrews 1991, Andrews and Monahan, 1992).

3 OLS Estimates: Asymptotic Results

If the hypothesis of no cointegration is rejected, the OLS estimates are consistent for the true cointegrating coefficient (Stock 1987) even though there exist simultaneous equation bias and serial correlation. This estimator is superconsistent converging to its true value at a rate of T^{-1} rather than the usual $T^{-1/2}$. In the case with $r>1$ cointegrating vectors, Wooldridge (1991) shows that OLS estimation selects the relation whose residuals are uncorrelated with any other I(0) linear combinations of regressors, also see Chapter 19, Hamilton 1994.

Since the OLS estimator has a non-normal limit distribution, which involves three parts: a mixture of normals, the unit root term, and the term caused by

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and quadratic spectral (QS) kernel. As Andrews (1991) shows that QS kernel has the best performance with respect to asymptotic truncated mean square error (MSE).

[6] Monte Carlo results in Andrews and Monahan (1992) show that prewhitening is very effective in reducing bias, and reducing over-rejection of t statistics constructed using standard kernel heteroskedasticity and autocorrelation consistent (HAC) covariance matrix estimator.

[7] Since the test outcomes sometimes depend on the choice of bandwidth (M), the use of a data dependent bandwidth parameter removes the arbitrariness associated with the choice of bandwidth.

simultaneous equation bias arising from the endogeneity of the regressors, (Phillips and Loretan 1991). Note that the last two terms produce a finite sample bias (asymptotically the bias will vanish) in median and mean and invalidate the use of standard distributions for testing hypothesis about the cointegrating vector. The inference about the cointegrating vectors based on the standard least squares output can be misleading. The usual test statistics (t-ratios and Chi-squared criteria) need to be modified to allow for the serial dependence in disturbances from cointegrating errors and endogeneity of the regressors.

4 Inference of Cointegrating Vectors

Although the estimates based on least squares estimation are consistent, there exist alternative estimates that are superior. Asymptotically efficient estimation of long-run equilibrium relationship can be achieved by a variety of methods, for example, full systems MLE (restricted by the imposition of unit roots)(Johansen 1988, 1991), fully modified OLS (with semiparametric serial correlation and endogeniety corrections) (Phillips and Hansen 1990, Hansen 1992a), and leads and lags estimator (dynamic OLS) (Stock and Watson[8] 1993), (for other

[8] Unlike Johansen and Phillips and Hansen's estimation in which each series is individually I(1), DOLS is developed

estimators see an informative review of Phillips and Loretan 1991, Gonzalo 1994). Each of these methods achieves full efficiency in the limit by working to estimate and eliminate the effects of long-run feedback between the errors on the long-run equilibrium relationship and the errors that drive the regressors. These methods are asymptotically equivalent and lead to conventional chi-squared criteria for inferential purposes with respect to cointegrating coefficients.

5 *Empirical Study: The Existence of the Long-Run Money Demand Functions in Taiwan (1961:4 - 1997:3)*

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

$$(m_t - p_t) = u + \alpha_y y_t + \alpha_i i_t + v_t$$

Therefore, the natural normalization is to take real balance $(m_t - p_t)$ as the dependent variable. Without prior knowledge

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for cointegrating regressions among general I(d≥1) variables with general deterministic components.

about whether the two interest rates (the own rate i, and unorganized money market rate umm) have equal coefficients with opposite signs, for each test we consider two basic models (with no structural break).

Model I: the interest rates enter the regression unrestricted

$$(m_t - p_t) = u + \alpha_y y_t + \alpha_i i_t + \alpha_{umm} umm_t + [\beta trend] + v_t \quad (3.8)$$

Model II: the spread is used in the regression

$$(m_t - p_t) = u + \alpha_y y_t + \alpha_{umm-i} (umm_t - i_t) + [\beta trend] + v_t \quad (3.9)$$

Relaxing price homogeneity restriction, we find that the tests results from the nominal specifications are similar to those in the real specifications. In Appendix 3, we show that the price homogeneity is not rejected by the data using the t-ratios constructed from the fully modified estimator of Philips and Hansen (1990). Therefore, in the following study, we only consider the real specification (3.8)-(3.9).

In Table 6-9[9], for each aggregate, we report the Z_α and Z_t tests, and the data dependent band width parameter $\hat{M}$.

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[9] Although the data are not seasonally adjusted, we do not add deterministic seasonal dummies in the regressions. However, the test results generated from the regressions with these dummies are very close to those in Table 6-9. Besides, the seasonal dummies are not significantly

The OLS estimates and their standard errors are also presented. Note as we mentioned above, the OLS estimates is a poor candidate for inference, so we should be cautious when using these standard errors to construct t-ratios to test significance of parameters of interest. In the bottom of each table, three outputs from OLS (DW, $\bar{R}^2$ and SEE) are presented. Because of its simplicity, the Durbin-Watson statistic (DW) was used by Engle and Granger (1987) to test the hypothesis of no cointegration[10]. Although a low DW might be expected as all dynamics are omitted, it should not be too low if the variables are cointegrated. Since the critical values for the test DW=0 are only available for bi-variable system in Engle and Granger (1987)[11], no inference is made from this statistic for our three- and four-variable system. We report it just for a quick check if we have a symptom of spurious regression: high $\bar{R}^2$ and a too low DW (Granger and Newbold 1974, Phillips 1986).

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different from zero, individually, and jointly using the t-ratios and Wald statistics, respectively, which are constructed from FM estimation. Therefore, the results are not reported here but are available upon request.

[10] Other than the DW statistic, six alternative tests of no cointegration (e.g. DF, ADF, restricted VAR(RVAR), augmented RVAR, unrestricted VAR (UVAR), augmented UVAR) are also considered in Engle and Granger (1987).

(1) The demand for real M1A equation

From Table 6, the null hypothesis of no cointegration is rejected at the 5% level by the Z_a and Z_t tests only when the regression includes a time trend. It suggests that there is some sort of long-run cointegrating relationship between the variables in demand for real M1A equation. It appears that the time trend is important to the conclusion of cointegration. Comparing test results from Model I and II, we note that these test results do not depend on how the opportunity cost variable enters the regression. That is, two interest rates enter restricted, or the spread is used.

From the OLS estimates, we notice that the coefficient on income in Regression (B) and (D) (where time trend is included) is huge. It is not only greater than unity, which suggests that M1A is a luxury for Taiwanese people, but also greater than two, which is rarely seen for the narrow aggregate in the literature. Without the time trend, the coefficients from Regression (A) and (Cs) are 1.28, which look more reasonable for this aggregate. Moreover, the coefficients on interest rates are correct. $\hat{\alpha}_{i^a} > 0$ as the theories predict and $\hat{\alpha}_{umm}<0$ and $\hat{\alpha}_{umm-i^a}<0$ as we conjecture

[11] Note the empirical distribution of the DW statistic in Engle and Granger (1987) is generated without considering a time trend in the cointegrating regression.

that the umm rate is a measure of the opportunity cost of holding real M1A in Taiwan.

(2) The demand for real M1B equation

From Table 7, we find that all of the tests reject the null of no cointegration at the 5% level. The results strongly suggest that there is a long-run relationship between the variables in this equation. The coefficients on interest rates have correct signs and the magnitudes are similar in the regressions with/without a trend.

(3) The demand for real M1BP equation

Like the demand for real M1B equation, from Table 8 we find that the null hypothesis of no cointegration is rejected at the 5% level by all of our tests. It strongly suggests that there exists a long-run relationship between the variables in the demand for real M1BP data. The signs are correct and the magnitudes of the estimates are similar to those in the real M1B equation. This seems no surprise because M1BP was constructed by aggregating similar deposits in the commercial banks and the Postal Savings System.

(4) The demand for real M2 equation

From Table 9, we reject the no cointegration hypothesis at the 5% level for all of the tests. It appears that there is also a long-run relationship in this specification for the full sample period (1961:4-1997:3) in Taiwan.

We should note that the coefficients on income are very different in the regressions with/without a trend. From Regression (A) and (C), the income elasticity is about 1.71, which suggests that M2 is a luxury. It becomes less than 1 (0.72) in Regression (B) and (D), indicating that there is economy of large scale in holding M2. Although $\hat{\alpha}_{umm}<0$ as we conjectured, $\hat{\alpha}_{i^{m2}}>0$, which contrasts with the prediction of theories.

Note that adopting different normalization, that is using y as the dependent variable, we find that the test results are similar to those in Table 6-9 for each aggregate. After re-normalizing the obtained coefficients such that the coefficient equals -1.00, we find that the re-normalized coefficients from the regressions without a trend are close to those in Table 6-9. However, with a trend in the regression, different normalization produces substantially different coefficients.

We also test if the ***umm*** rate and the own rate on each money are cointegrated. Regardless of normalization, Z_t and Z_a tests cannot reject the hypothesis of no cointegration occurred between these two interest rates. We report the results in Appendix 4.

Table 6
Testing for no cointegration in Taiwanese demand for real M1A

Model I: $RM1A_t = u + \alpha_y y_t + \alpha_{i^a} i_t^a + \alpha_{umm} umm_t + [\beta trend] + v_t$

		OLS estimates				Cointegration tests					
(A)	u	y	i^a	umm	trend	Z_t	Z_α	$\hat{M}$	$\bar{R}^2$	DW	SEE
	-3.80 (0.17)	1.28 (0.01)	0.037 (0.067)	-0.008 (0.003)		-3.27	-20.57	0.68	0.989	0.30	0.117
(B)	-13.25 (1.52)	2.11 (0.13)	0.107 (0.060)	-0.013 (0.002)	-0.018 (0.002)	-5.98**	-56.2**	0.57	0.991	0.78	0.104

Model II: $RM1A_t = u + \alpha_y y_t + \alpha_{umm-i^a} (umm_t - i_t^a) + [\beta trend] + v_t$

		OLS estimates			Cointegration tests					
(C)	u	y	umm-i^a	trend	Z_t	Z_α	$\hat{M}$	$\bar{R}^2$	DW	SEE
	-3.83 (0.15)	1.28 (0.01)	-0.008 (0.003)		-3.27	-20.49	0.74	0.989	0.29	0.116
(D)	-13.95 (1.51)	2.17 (0.13)	-0.014 (0.002)	-0.018 (0.002)	-5.18**	-44.5**	0.66	0.991	0.62	0.101
								Critical values for DW=0		
								1%	5%	10%
								0.511	0.386	0.322

Note: Z_t and Z_α tests are calculated using a prewhitened quadratic spectral kernel and an automatic bandwidth estimator.
The critical values of DW = 0 are taken from Table II in Engle and Granger (1987).
An * and ** indicates significance at the 5% and 10% levels, respectively.
SEE represents standard errors of regression estimate.
For critical values for Z_t and Z_α, see Table 10.

Table 7
Testing for no cointegration in Taiwanese demand for real M1B

Model I: $RM1B_t = u + \alpha_y y_t + \alpha_{i^b} i_t^b + \alpha_{umm} umm_t + [\beta trend] + v_t$

	OLS estimates					Cointegration tests					
(A)	μ	y	i^b	umm	trend	Z_t	Z_α	$\hat{M}$	$\bar{R}^2$	DW	SEE
	-6.50 (0.29)	1.52 (0.02)	0.052 (0.030)	-0.020 (0.003)		-4.83**	-39.8**	0.62	0.995	0.56	0.093
(B)	-9.01 (1.39)	1.75 (0.12)	0.048 (0.030)	-0.021 (0.003)	-0.004 (0.002)	-5.60**	-50.5**	0.50	0.995	0.70	0.093

Model II: $RM1B_t = u + \alpha_y y_t + \alpha_{umm-i^b} (umm_t - i_t^b) + [\beta trend] + v_t$

	OLS estimates				Cointegration tests					
(C)	u	y	umm-i^b	trend	Z_t	Z_α	$\hat{M}$	$\bar{R}^2$	DW	SEE
	-6.94 (0.12)	1.55 (0.01)	-0.015 (0.002)		-4.83**	-40.0**	0.74	0.995	0.57	0.094
(D)	-9.67 (1.34))	1.80 (0.12)	-0.017 (0.002)	-0.005 (0.002)	-5.65**	-51.4**	0.61	0.995	0.73	0.093
								Critical values for DW=0		
								1%	5%	10%
								0.511	0.386	0.322

See notes to Table 6.

Table 8
Testing for no cointegration in Taiwanese demand for real M1BP

Model I: $RM1BP_t = u + \alpha_y y_t + \alpha_{i^{bp}} i_t^{bp} + \alpha_{umm} umm_t + [\beta trend] + v_t$

	OLS estimates					Cointegration tests					
	u	y	i^{bp}	umm	trend	Z_t	Z_α	$\hat{M}$	$\bar{R}^2$	DW	SEE
(A)	-6.84 (0.20)	1.58 (0.01)	0.096 (0.030)	-0.030 (0.003)		-5.14**	-44.2**	0.30	0.995	0.62	0.095
(B)	-8.90 (1.64)	1.75 (0.14)	0.086 (0.016)	-0.029 (0.004)	-0.003 (0.002)	-5.70**	-52.1**	0.07	0.995	0.73	0.094

Model II: $RM1BP_t = u + \alpha_y y_t + \alpha_{umm-i^{bp}} (umm_t - i_t^{bp}) + [\beta trend] + v_t$

	OLS estimates				Cointegration tests					
	u	y	umm-i^{bp}	trend	Z_t	Z_α	$\hat{M}$	$\bar{R}^2$	DW	SEE
(C)	-7.70 (0.15)	1.63 (0.01)	-0.017 (0.003)		-4.39**	-34.5**	0.76	0.994	0.50	0.104
(D)	-12.94 (1.40)	2.09 (0.12)	-0.019 (0.003)	-0.009 (0.003)	-6.07**	-58.0**	0.61	0.995	0.82	0.100

Critical values for DW=0		
1%	5%	10%
0.511	0.386	0.322

See notes to Table 6.

Table 9
Testing for no cointegration in Taiwanese demand for real M2

Model I: $RM2_t = u + \alpha_y y_t + \alpha_{i^{m2}} i_t^{m2} + \alpha_{umm} umm_t + [\beta trend] + v_t$

		OLS estimates				Cointegration tests					
(A)	u	y	i^{m2}	umm	trend	Z_t	Z_α	$\hat{M}$	$\bar{R}^2$	DW	SEE
	-7.68 (0.10)	1.72 (0.01)	-0.015 (0.010)	-0.024 (0.004)		-6.17**	-56.9**	0.54	0.996	0.79	0.082
(B)	-3.19 (0.75)	0.76 (0.07)	-0.015 (0.006)	-0.018 (0.001)	0.020 (0.001)	-4.67**	-39.7**	1.18	0.998	0.59	0.051

Model II: $RM2_t = u + \alpha_y y_t + \alpha_{umm-i^{m2}} (umm_t - i_t^{m2}) + [\beta trend] + v_t$

		OLS estimates			Cointegration tests					
(C)	u	y	$(umm-i^{m2})$	trend	Z_t	Z_α	$\hat{M}$	$\bar{R}^2$	DW	SEE
	-7.59 (0.12)	1.71 (0.01)	-0.035 (0.003)		-5.90**	-53.0**	0.69	0.996	0.61	0.086
(D)	3.66 (0.85)	0.72 (0.07)	-0.027 (0.002)	0.021 (0.001)	-4.50**	-36.8**	0.92	0.998	0.51	0.059
								Critical values for DW=0		
								1%	5%	10%
								0.511	0.386	0.322

See notes to Table 6.

(5) Summary

From the residual based cointegration test, there is strong evidence suggesting that cointegration occurs over 1961:4 - 1997:3 in the variables chosen for the Taiwanese demand for real M1B, real M1BP, and real M2 equations. These results are invariant to the detrending procedures. However, the time trend is important to the conclusion of cointegration for real M1A equation.

Table 10
Critical values for residual-based cointegration tests

	Model I		Model II	
Test	5%	10%	5%	10%
Z_t				
C	-4.20	-3.89	-3.85	-3.55
C/T	-4.53	-4.23	-4.20	-3.89
Z_a				
C	-31.02	-26.94	-26.00	-22.33
C/T	-35.30	-31.02	-31.02	-26.94

Note: Critical values for Z_t and Z_a are from Mackinnon (1990) and Haug (1992), respectively.

III: THE STABILITY OF THE LONG-RUN DEMAND FOR MONEY FUNCTIONS IN TAIWAN: 1961:4-1997:3

In CHAPTER 2, we discussed how important role a stable long-run money demand function played in the macroeconomic analysis. In this section, we examine the empirical evidence of the stability in the long-run money demand relationships in Taiwan.

From the analysis above, while the conventional residual based cointegration tests in the previous section provide us evidence of a long-run relationship in money demand for real M1A, M1B, M1BP, and M2 in Taiwan for entire sample period, they do not provide us with much information concerning the question of whether or not there is a regime shift in these cointegrating relationships. Moreover, another unsettled issue in the money demand analysis is whether the narrowly defined or broadly defined monetary aggregates yield a more reliable long-run relationship: "under differing institutional arrangements, changes in the social and political environment and changes in economic conditions" (Meltzer 1963). Lucas (1988) also argues that the demand for money function will be stable over time provided that preferences and trading technology are stable. The following events in Taiwan may play a part in explaining potential changes in the long-run parameters.

1963: the New Taiwan dollar was pegged to the US dollar

1973:3-4: the first oil crisis

1976: the creation of the organized money market

1978: the adoption of a floating exchange rate system

1979: the second oil crisis

1980: interest rate deregulation (ending in 1989)

1987: liberalization of the foreign exchange control system

1989: bank entry deregulation

With these events in mind, we investigate the stability in these four demand equations for real balances (RM1A, RM1B, RM1BP, and RM2) in Taiwan (1961:4-1997:3).

In order to examine the stability in the cointegrating relation and to search for a correct model, we apply the extended cointegration tests with regime shifts developed by Gregory and Hansen (1996) and the parameter stability tests of Hansen (1992b).

1 Residual-Based Tests for Cointegration in Models with Regime Shifts: Gregory and Hansen (1996) Z_t^* and Z_α^* Tests

In the standard tests for no cointegration introduced in the previous section, it is presumed that the cointegrating vector is time-invariant under the alternative hypothesis:

$$H_1: \; x_{1t} = u + \alpha' x_{2t} + v_t, \quad t = 1, 2, \ldots, n$$

where $X_t = (x_{1t}, x_{2t}')'$ are I(1) and v_t is I(0). In this model the parameters μ and α are cointegrating vector and are considered as time-invariant. In our empirical study (especially for the demand for real M1A data), we wish to entertain the possibility that the series are cointegrated, in the sense that a linear combination (the cointegrating vector) has shifted at one unknown point in the sample; hence, the standard cointegration tests are not appropriate. Gregory and Hansen (1996) extend the ADF, Z_t, and Z_t tests designed to test the null of no cointegration against the alternative of cointegration in the presence of a possible regime shift. Such tests can detect cointegrating relations when there is a change in the intercept and/or slope coefficient occurring at unknown date. Rejection of the null hypothesis, therefore, provides evidence in favor of this specification.

The principle of testing for cointegration in model with regime shifts is the same as the standard cointegration procedure as introduced in Section II, which is a test of unit root in the cointegrating residuals. While the standard testing procedure is to set up the null of no cointegration against the alternative of cointegration, the extended tests set the alternative hypothesis to be cointegration while allowing for the structural change which is reflected in changes in the intercept and/or changes to the slope.

Specifically, Gregory and Hansen (1996) consider three alternative hypotheses. In order to model structural change, we first define a dummy variable. As in Perron and Vogelsang (1992) and Zivot and Andrews (1992), we assume that there is at most one change and we denote the date of break, should it occur, by t_b, with $1 < t_b < n$, where n is the sample size. Hence, the dummy variable is defined as

$$D(t_b)_t = 1 \text{ if } t > t_b$$
$$= 0, \text{ otherwise.}$$

We also define the location of the break fraction $\tau = t_b/n$, $\tau \in (0,1)$.

Three alternative hypotheses of cointegration with structural changes ($H_2 - H_4$) are

H_2: Level Shift (C)

$$x_{1t} = u_1 + u_2 D(t_b)_t + \alpha' x_{2t} + v_t, \quad t = 1,2\ldots,n \qquad (3.10)$$

This model specifies that there is a level shift in the cointegrating relationship, which is modeled as a change in the intercept μ, while the slope coefficients α are held constant. This implies that the equilibrium equation has shifted in a parallel fashion. μ_1 represents the intercept before the shift, and μ_2 represents the change in the intercept at the time of the shift. As in Gregory and Hansen (1996), we also denote this model C.

H_3: Level shift with trend (C/T)

$$x_{1t} = u_1 + u_2 D(t_b)_t + \alpha' x_{2t} + \beta t + v_t \qquad t = 1,2,\ldots,n \qquad (3.11)$$

In this model, we allow for a time trend into the level shift (C) model.

H_4: Regime shift (C/S)

$$x_{1t} = u_1 + u_2 D(t_b)_t + \alpha_1' x_{2t} + \alpha_2' x_{2t} D(t_b)_t + v_t \qquad t = 1,2\ldots,n \qquad (3.12)$$

This model allows both the intercept and the slope vector to shift. This permits the equilibrium relation to rotate as well as shift parallel. In this case μ_1 and μ_2 are as in the level shift model (C), α_1 denotes the cointegrating slope coefficients before the regime shift, and α_2 denotes the change in the slope coefficients.

We use the same approach described in the standard residual based cointegration tests, but take account the unknown break date to compute the statistics Z_t^* and Z_α^*. For each possible τ (or t_b), we estimate one of the models C, C/T, C/S by OLS, yielding the residuals $\hat{v}_{t\tau}$. The subscript τ on the residuals denotes the fact that the residual sequence depends on the choice of change point τ. The resulting Phillips' test statistics for each for each τ (or t_b) is denoted $Z_t(\tau)$ and $Z_\alpha(\tau)$ for each τ (or t_b). The testing scheme is to choose the breakpoint that gives the least favorable result for the null hypothesis of no cointegration; the statistic of interest is the then smallest value of each of the above statistics across all values of τ, denoted by Z_t^* and Z_α^*. In other words, τ (or t_b) is chosen to minimize $Z_\alpha(\tau)$ and $Z_t(\tau)$ for testing that $\hat{v}_{t\tau}$ is I(1).

As suggested in the earlier literature, such as, Banerjee, Lansdaine, Stock (1992), and Andrews (1990), we apply some trimming such that $.15 < \tau < .85$. For our study, the sample size $n = 144$, so the possible break dates $t_b \in ([22, 122])$.

2 Hansen (1992b) LM Tests for Parameter Instability in Cointegrating System

The test statistics proposed by Hansen (1992b) are designed to test the hypothesis of no regime shift against the alternative of a regime shift. Unlike the tests in the previous section whose null hypothesis is no cointegration, the null hypothesis of Hansen's tests is Engle-Granger cointegration. Hansen's (1992b) tests are joint tests on all regression parameters in a cointegrating regression. A statistically significant test statistic is taken as evidence against the standard cointegration in favor of the regime shift model.

Under the null hypothesis of cointegration, we assume that the generating mechanism for $X_t = (x_{1t}, x_{2t}')'$ is the cointegrated system

$$x_{1t} = \alpha' x_{2t} + u_{1t} \tag{3.13}$$

$$x_{2t} = \Pi t + S_{2t} \tag{3.14}$$

$$\Delta S_{2t} = u_{2t} \tag{3.15}$$

As before $u_t = (u_{1t}, u_{2t}')'$ is p-vector stationary series; the assumption characterizing the innovation vector u_t see Hansen (1992a). (3.13) can be thought of as a stochastic version of the linear long-run equilibrium relationship x_{1t}

$= \alpha' x_{2t}$ with u_{1t} representing stationary deviations from equilibrium. (3.14)-(3.15) is a reduced form which specifies x_{2t} as a general integrated process, the outcome of superimposed shock u_{2s} (s≤t) that influence the process period after period.

Since the asymptotic distributions of the test statistics depend on the nature of the trends in the regressors x_{2t} (Hansen 1992b), as in Section II we estimate both unrestricted model:

$$x_{1t} = \mu + \beta t + \alpha' x_{2t} + u_{1t} \tag{3.16}$$

and the restricted model:

$$x_{1t} = \mu + \alpha' x_{2t} + u_{1t} \tag{3.17}$$

For simplicity, we rewrite (3.13) and (3.14) as

$$x_{1t} = A X_{2t} + u_{1t} \tag{3.18}$$

where X_{2t} = (constant, t, x_{2t}).

The constancy tests require an estimation of A in (3.18) that has a mixed normal asymptotic distribution so estimators can be the fully modified (FM) estimator of

Phillips and Hansen (1990), the maximum likelihood estimator (MLE) of Johansen (1998b,1991), or the "leads and lags" estimator of Saikkonen (1991) and Stock and Watson (1993)[12]. The test statistics using these asymptotically equivalent estimators would have the same asymptotic distributions. Following Hansen (1992b), we also consider the FM estimator of Phillips and Hansen (1990), which uses semiparametric methods for serial correlation and for endogeneity.

The LM tests for parameter stability are L_c, MeanF, and SupF tests[13]. The three proposed tests are all tests of the same null hypothesis of cointegration with stable parameters but differ in their choice of alternative hypothesis. The SupF test is appropriate to discover whether there is a swift shift in regime, like Z_t^{*} and Z_α^{*} which assume a single structural break occurring at unknown date $\hat{t}_b$. The other two statistics capture the notion of an unstable model that gradually shifts over time. If the parameter variation is relatively constant throughout the sample, the L_c is the appropriate choice. In practice, all of the tests will tend to have power in similar directions.

[12] Also see Phillips and Loretan (1991) for a review

[13] For constructing these statistics, see Hansen 1992b.

3 *Empirical Study: The Stability of the Long-Run Money Demand Functions in Taiwan (1961:4 - 1997:3)*

In Table 11 - 14, we report the extended tests (Z_α^*, Z_t^*), parameter stability tests - (L_c, MeanF, and SupF), the FM estimates, and their standard errors. In each specification, the covariance parameters for constructing statistics are estimated using a QS kernel on residuals prewhitened with a VAR(1) for Hansen (1992b) procedure, and AR(1) for Gregory and Hansen (1996) procedure, respectively. In both procedures, the bandwidth parameter is selected according to the recommendations of Andrews (1991), using univariate AR(1) approximating models[14]. All of the statistics are computed using the trimming region (.15.85) as in Section II.

In the Hansen's (1992b) tests and estimation procedure, we will proceed in a "general to simple" specification search. We first examine the unrestricted regression in which a simple time trend is included and the two interest rates enter the regression unrestricted

$$(m_t - p_t) = u + \alpha_y y_t + \alpha_i i_t + \alpha_{umm} umm_t + \beta t + v_t$$

[14] For details, see Hansen 1992b, Gregory and Hansen 1996.

Subsequently, we re-estimate the regression by dropping insignificant variables and constraining the two interest rates to have equal coefficients with opposite sign if the restriction is compatible with the data. We use the usual t ratio constructed from FM to test the significance of the individual coefficients; and the Wald statistics to test the linear restriction $\alpha_i + \alpha_{umm} = 0$. Under the null the Wald statistic is distributed as a $\chi^2(1)$.

For comparison we present all the results from the unrestricted and restricted regressions. However, the final conclusion is drawn from the restricted regression in which all the coefficient are significant. We label these regressions by (A), (B), etc.

For each aggregate, the results from the restricted regression, where the conclusion is drawn, are also presented graphically in Figure 3-10, which contains time plot of $Z_t(\tau)$ and $Z_\alpha(\tau)$ over the truncated sample along with 5% and 10% critical values for the minimum Z_t^* and Z_α^* statistics[15]. Figure 11-14 plot the sequence of F statistics along with the 5% critical value for the SupF statistic for the truncated sample.

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[15] For comparison, we graph all extended tests Z_t^* and Z_α^* even when time trend is not significant.

$\hat{t}_b$ denotes the estimated break point which is the point in the sample where the smallest value of the test statistic (Z_t^* and Z_α^*) is obtained and where the largest value of F statistic (SupF) is obtained.

(1) The demand for real M1A equation

From unrestricted regression (B) in Table 11, we found that i^a is insignificant and has a negative sign, which contrasts with the prediction of the theories. Using DOLS estimation (Stock and Watson 1993), the coefficient has the correct sign but still insignificant, with/without a time trend in the regression[16]:

$$\underset{}{RM1A_t} = \underset{(6.79)}{-16.24} + \underset{(0.61)}{2.38 y_t} + \underset{(0.23)}{0.13 i^a_t} - \underset{(0.01)}{0.02 umm_t} - \underset{(0.01)}{0.02 t} + v_t$$

$$RM1A_t = \underset{(0.63)}{-3.88} + \underset{(0.05)}{1.27 y_t} + \underset{(0.26)}{0.14 i^a_t} - \underset{(0.01)}{0.01 umm_t} + v_t$$

Dropping i^a, we re-estimate regression (D). In regression (D), all coefficients are significant at either 5% or 10% level (the critical values of the t-ratio is 1.65 for 5%, 1.28 for 10%) and have correct signs (the time trend is

[16] The leads and lags in the DOLS estimation are included in the regressions but their coefficients are not reported here.

negative from data information). Therefore, the conclusion from test results are based on this regression.

In regression (D), all three stability tests L_c, MeanF and SupF are not rejected at the 5% level, implying this long-run relationship remains stable over the whole sample period. Hansen (1992b) discovered that the L_c can be regarded as a test of the null of cointegtration against the alternative of no cointegration since the lack of cointegration is a special case of the alternative hypothesis (an unstable intercept) where the intercept follows a random walk. Since all of the parameters appear stable, including the constant, this implies that the null of cointegration hypothesis is not rejected.

From Panel B of Table 11, we can see that the null hypothesis of no cointegration is rejected at the 5% level by Z_α^* and Z_t^* tests only when the regression includes a time trend (the C/T formulation). We should note that although the null is rejected by Z_α^* and Z_t^* tests, providing evidence in favor of C/T specification (Eq.(3.10)), we are unable to conclude from these two tests that there is a structural break. The reason is that the alternative hypotheses $H_2 - H_4$ (Eq.(3.10)-(3.12)) in the extended tests contain as a special case the standard model of cointegration with no regime shift, a conventional cointegrated system with constant parameters could produce results of cointegration from these extended regressions. Using the umm rate alone, we also conduct the standard cointegration tests Z_t (= -5.80**) and Z_α (= -53.80**), finding that the null of no cointegration is also rejected.

Turning to the FM estimates in Regression (D), we notice that the income elasticity (1.86) is huge as the OLS estimates in Table 6 and significantly different from unity. As we mentioned earlier that this is rarely seen in the literature for this narrow aggregate. The estimates obtained in the regressions without a time trend Regression (A), (C), and (E) are more reasonable.

The conclusion of cointegration depends on the detrending procedure and it appears that the linear combination of the I(1) variables under study reverts to a trend. However the significant trend gives us difficulty in interpreting the results. Firstly, economic theories only suggest that $(m_t - p_t) - f(y_t, R_t)$ should have an equilibrium, that is, the discrepancy should not contain a trend. Moreover, to accept this long-run relationship, we need a reasonable explanation for a huge income elasticity. Balancing the evidence we have obtained, we doubt this equilibrium relationship.

We notice that the test results and coefficients from the regression (F) in which the spread is used are very close to those from regression (D) where the umm rate is used.

Finally we present the test results Z_t^*, and Z_α^* for C, C/T, and C/S (Panel B in Table 11) graphically in Figure 3-4, and the sequence of F statistic from Regression (D) in Figure 11. We can see that the sequences of $Z_t(\tau)$ and $Z_\alpha(\tau)$ for C/T regression cross the 5% Z_t^* and Z_α^* critical values several times, suggesting rejection of the no cointegration. The sequence of F statistics for structural change is well below the 5% critical value, indicating the parameters are stable.

(2) <u>The demand for real M1B equation</u>

From the unrestricted regression (B) in Table 12, the estimated coefficient of i^b and the time trend are not significant at the 5% level, which is also robust to the DOLS estimation, so we re-estimated regression by dropping them. In Regression (C), all of the coefficients are significant at the 5% level and possess correct signs. The following analysis is based on Regression (C).

The L_c, MeanF and SupF tests cannot reject the null hypothesis of a stable cointegrating relationship at the 5% level. As in the analysis of demand for real M1A, the L_c

test implies that the hypothesis of cointegration is not rejected.

Turning to the Z_{α}^{*} and Z_{t}^{*} tests (Panel B in Table 12), the null of no cointegration is rejected at the 5% level for all of the tests (10% for the C formulation). It is consistent with the L_{c} test, suggesting strong evidence of cointegration among real M1B data. Again, the Z_{α}^{*} and Z_{t}^{*} tests do not provide us with much information about any shift in this relationship for the reason we mentioned in the analysis of the demand for real M1A. Using one single interest rate, umm, we compute the standard cointegration tests as Z_{t}(= -5.80**) and Z_{α}(= -53.80**). Hence the null of no cointegration is also rejected. We note that the stability tests results are robust to the regressions we used; the only test which rejects the null at both 5% and 1% level is the L_{c} test in Regression (A). This suggests that this long-run relationship among the variables we chose is quite stable over the sample period. This result contrasts with that in Lee (1994) who finds a downward shift in the demand for real M1B around 1982:4.

In Regression (C), the estimated income elasticity (1.56) is significantly different from unity at the 5% level so M1B is a luxury to Taiwanese people. The interest semi-elasticity is around - 0.015. Using the sample mean value of umm rate (24.12%) to convert the interest semi-elasticity, we obtain the interest elasticity is of the order of -0.36.

Constraining $\alpha_{r} = -\alpha_{umm}$, which is not rejected by the Wald statistic, and then using the interest rate spread, we find that the test results and estimates in regression (E) are similar to Regression (C).

Finally we present the test results Z_t^*, and Z_α^* for C, C/T, and C/S graphically (from Panel B in Table 12) in Figure 5-6, and the sequence of F statistic from Regression (C) in Figure 12, along with the critical values.

(3) <u>The demand for real M1BP equation</u>

From the unrestricted regression (B) in Table 13, we find that the time trend is not significant, which is also observed in DOLS, so we drop it and re-estimate Regression (A). The coefficients in Regression (A) are all significant at the 5% level and the signs are consistent with we predicted. Thus we interpret the results based on Regression (A) in the following.

All of three parameter stability tests reject the null at the 5% level, casting some doubt about the existence of a stable long-run relationship in this specification. All Z_α^* and Z_t^* tests (from Panel A in Table 13) reject the null of no cointegration at the 5% level in favor of the structural break specification.

As we mentioned, this monetary aggregate is obtained by grouping seemingly homogenous deposits in the commercial banks (M1B) and in the Postal Savings System (passbook savings deposits). Since we have shown that the demand for real M1B displays considerably constant over the same sample period, the instability in this equation may arise from the behavior of postal saving deposits. After examining the graphs of data series (RM1B, RM1BP, and interest rates) in Figure 1 for these two equations, we found that the trend and cycle of two data set exhibit similar pattern over the whole sample period. Casual inspection of the time series provides us with little information about the possible source of nonconstancy. However, parameters instability in the demand for real M1BP may be explained by the difference in liquidity characteristics in M1B and passbook savings deposits in the Postal Savings System. J.C. Lee (1994) finds that the annul turnover rate of the demand deposits in the Postal Savings System is significantly lower than those in the commercial banks[17]. Deposit holders seem to "regard" the deposits in the commercial banks to be more liquid than

[17] The turnover rates per year for various deposits are: 2.9, and 0.7 for passbook savings deposits and time deposits in the Postal Savings System, respectively. 200 for checking accounts, 50 for passbook deposits, and 20 for

those in the Postal Savings System even though there is no substantial difference in the easiness to convert the deposits into a medium of exchange. This indicates that these postal deposits on the average are not being actively used for transaction purpose, and they may be more like investment balances or savings vehicles to deposit holders, therefore, their behavior may be influenced by wealth, for example, instead of a measure of transactions (real GNP) or other factors. Thus, the specification for this monetary aggregates might not adequately capture a stable money demand relationship.

To summarize, the existence of cointegration in this specification is problematic. This specification appears subject to a structural change and may not be correctly specified.

The sequences of $Z_t(\tau)$ and $Z_\alpha(\tau)$ (from Panel A in Table 13) in Figure 7 and 8 indicate that there is a well-defined minimum from these six tests. The smallest statistic is obtained roughly at one third of the sample (τ=0.34, $\hat{t}_b$=73:3); the break date coincides the first oil crisis. In Figure 13, we plot the sequence of F statistic for the restricted Regression (A).

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passbook savings deposits in commercial banks, J.C. Lee (note 9, Chapter 9, 1994).

(4) The demand for real M2 equation

From the unrestricted regression (B) in Table 14, each of the coefficients is significant at either 5% or 10% level. However we notice that $\hat{\alpha}_{r^{m2}} < 0$, which is frequently obtained in the studies of the Taiwanese literature and contrasts with the prediction of the theories. $\hat{\alpha}_{r^{m2}} < 0$ is also observed in the estimation of OLS (Table 9), DOLS, and Johansen's procedure with/without a trend.

Based on Regression (B), the null of cointegraion with constant parameters is rejected by the L_c test, MeanF and SupF at the 5% level. A stable long-run relationship in the demand for real M2 is doubtful. The negative coefficient on the own rate seems to signal this instability. Because of the instability occurred in real M1BP equation instability for this equation seems no surprise. Since this specification may be subject to a structural break, it may not be correctly specified. Before it is re-examined, we cannot provide any explanation for the negative own rate coefficient in this study. It needs further research which is beyond the scope of this study. Note the results of instability are robust to the detrending procedure, which can be seen from Regression (A).

The sequence of $Z_t(\tau)$ and $Z_\alpha(\tau)$ for C, C/T, and C/S (from Panel A in Table 14) are presented in Figure 9 and 10, and, the sequence of F statistic for regression in Figure 14.

(5) Summary

From the analysis above, it appears that the narrowly defined monetary aggregate (M1B) yield more reliable long-run money-demand relationship in Taiwan in spite of several potential regime shifts over the whole sample period (1961:4 - 1997:3). The broadly defined monetary aggregates (M1BP, and M2) do not provide evidence of a stable long-run relationship; the specifications in our study may not adequately describe the "true" money demand behavior for these two aggregates. Moreover, the most stable long-run money demand relationship occurs in the demand for real M1B. The time trend is important to the conclusion of a stable long-run demand for real M1A function; however, to accept this relationship, one need to provide a reasonable explanation for a huge income elasticity.

One the one hand, that M1B yields most stable money demand relationship can be justified by the transactions-demand theory, with money held as a medium of exchange. The real quantity of money demanded is an increasing function of some measure of the volume of real transactions (real GNP) and a decreasing function of the opportunity cost of holding money (umm or the spread).

On the other hand, the demand for broadly defined money (RM1BP and RM2) might be associated with the portfolio and speculative demand, with money as one of several possible assets in which wealth may be held. Thus, instability detected in these two aggregates is explained by misspecification. That is, the resulting money demand depends on wealth rather than income (the proxy of the volume of transactions) and some measure of the volatility of alternative assets' returns in addition to their expected returns. However, we leave this conjecture unanswered in our study.

Table 11
Cointegration tests with regime shifts and parameter constancy tests: real M1A data

Panel A: $RM1A_t = u + \alpha_y y_t + \alpha_{i^a} i_t^a + \alpha_{umm} umm_t + [\beta trend] + v_t$

	FM estimates					Hansen stability tests					
(A)	u	y	i^a	umm	trend	L_c	MeanF	SupF	Wald $(\alpha_{i^a} + \alpha_{umm} = 0)$ [a]		
	-3.62	1.29	-0.060	-0.022		4.53**	7.96	19.6**	0.15		
	(0.53)	(0.05)	(0.215)	(0.010)				[74:1]			
(B)	-10.52	1.89	-0.046	-0.019	-0.012	3.33**	9.21	17.28	0.23		
	(3.40)	(0.30)	(0.136)	(0.006)	(0.006)			[89:2]			
							Gregory and Hansen cointegration tests				
						Z_t^*			Z_α^*		
						C	C/T	C/S	C	C/T	C/S
						-4.03	-7.32**	-5.24	-31.98	-77.7**	-45.6
						[80:3]	[87:2]	[87:2]	[80:3]	[87:2]	[87:2]

Panel B: $RM1A_t = u + \alpha_y y_t + \alpha_{umm} umm_t + [\beta trend] + v_t$

	FM estimates				Hansen stability tests					
(C)	u	y	umm	trend	L_c	MeanF	SupF			
	- 3.67	1.29	-0.014		2.65**	6.14	17.83**			
	(0.61)	(0.05)	(0.013)				[67:1]			
(D)	-10.37	1.86	-0.014	-0.012	0.70	6.67	9.66			
	(4.39)	(0.39)	(0.008)	(0.008)			[82:1]			
						Gregory and Hansen cointegration tests				
					Z_t^*			Z_α^*		
					C	C/T	C/S	C	C/T	C/S
					-3.34	-7.20**	-4.88	-21.38	-75.5**	-40.27
					[73:3]	[87:2]	[87:2]	[73:3]	[87:2]	[87:4]

Table 11 (cont'd)

Panel C: $RM1A_t = u + \alpha_y y_t + \alpha_{umm-i^a}(umm_t - i_t^a) + [\beta trend] + v_t$

	FM estimates				Hansen stability tests					
(E)	u	y	umm-i^a	trend	L_c	MeanF	SupF			
	- 3.66	1.28	-0.014		2.73**	6.15	19.07**			
	(0.63)	(0.05)	(0.013)				[67:1]			
(F)	-10.82	1.90	-0.015	-0.013	0.66	6.21	9.05			
	(4.39)	(0.39)	(0.009)	(0.008)			[82:1]			
					Gregory and Hansen cointegration tests					
					Z_t^*			Z_α^*		
					C	C/T	C/S	C	C/T	C/S
					-3.95	-7.29**	- 4.82	-28.89	-78.4**	-39.99
					[91:4]	[87:2]	[87:4]	[91:4]	[87:2]	[87:4]

Note (a): Critical values for $\chi^2(1)$ are 3.84 and 2.71 for 5% and 10% level, respectively.
** indicates significant at the 5% level.

Figures in [ ] are the estimated breakpoint $\hat{t}_b$.
See Table 15-16 for critical values.

Table 12
Cointegration tests with regime shifts and parameter constancy tests: real M1B data

Panel A: $RM1B_t = u + \alpha_y y_t + \alpha_{i^b} i_t^b + \alpha_{umm} umm_t + [\beta trend] + v_t$											
	FM estimates					Hansen stability tests					
(A)	u	y	i^b	umm	trend	L_c	MeanF	SupF	Wald($\alpha_{i^b} + \alpha_{umm} = 0$)[a]		
	-6.42	1.53	0.062	-0.024		1.79**	5.16	10.08	0.19		
	(0.89)	(0.07)	(0.094)	(0.011)				[74:1]			
(B)	-7.50	1.62	0.033	-0.022	-0.002	1.26**[1]	7.39	11.40	0.02		
	(3.75)	(0.33)	(0.082)	(0.009)	(0.006)			[87:4]			
						Gregory and Hansen cointegration tests					
						Z_t^*			Z_α^*		
						C	C/T	C/S	C	C/T	C/S
						-5.31**	-6.50**	-6.38**	-47.19	-65.8**	-62.59
						[80:3]	[86:4]	[87:1]	[73:3]	[87:1]	[87:1]
Panel B: $RM1B_t = u + \alpha_y y_t + \alpha_{umm} umm_t + [\beta trend] + v_t$											
	FM estimates					Hansen stability tests					
(C)	u	y		umm	trend	L_c	MeanF	SupF			
	- 6.96	1.56		-0.015		0.76	3.19	5.60			
	(0.44)	(0.03)		(0.009)				[74:2]			
(D)	-10.01	1.83		-0.017	-0.006	0.42	4.23	5.97			
	(4.13)	(0.36)		(0.008)	(0.007)			[87:3]			
						Gregory and Hansen cointegration tests					
						Z_t^*			Z_α^*		
						C	C/T	C/S	C	C/T	C/S
						-5.25**	-6.50**	-6.42**	-45.45*	-65.1**	-62.9**
						[73:3]	[86:4]	[87:2]	[73:3]	[86:4]	[87:1]

Note 1: L_c is not significant at the 1% level.

Table 12 (cont'd)

Panel C: $RM1B_t = u + \alpha_y y_t + \alpha_{umm-i^b}(umm_t - i_t^b) + [\beta trend] + v_t$

	FM estimates				Hansen stability tests					
(E)	u	y	$(umm-i^b)$	trend	L_c	MeanF	SupF			
	- 6.81	1.55	-0.017		0.76	3.19	5.60			
	(0.48)	(0.03)	(0.010)				[74:2]			
(F)	- 9.86	1.82	-0.02	-0.006	0.43	3.99	5.72			
	(4.11)	(0.36)	(0.01)	(0.008)			[87:3]			
					Gregory and Hansen cointegration tests					
					Z_t^*			Z_α^*		
					C	C/T	C/S	C	C/T	C/S
					-5.21**	-6.35**	-6.19**	-44.66*	-63.0**	-60.7**
					[73:3]	[86:4]	[87:2]	[73:3]	[86:4]	[87:2]

See notes to Table 11.

Table 13

Cointegration tests with regime shifts and parameter constancy tests: real M1BP data

Panel A: $RM1BP_t = u + \alpha_y y_t + \alpha_{i^{bp}} i_t^{bp} + \alpha_{umm} umm_t + [\beta trend] + v_t$

	FM estimates					Hansen stability tests				
(A)	u	y	i^{bp}	umm	trend	L_c	MeanF	SupF	$Wald(\alpha_{i^{bp}} + \alpha_{umm} = 0)$ [a]	
	-7.11 (0.55)	1.59 (0.04)	0.083 (0.039)	-0.027 (0.011)		2.46**	10.20**	39.73** [67:1]	3.14*	
(B)	-2.80 (4.13)	1.22 (0.36)	0.106 (0.041)	-0.029 (0.010)	0.008 (0.007)	2.10**	20.02**	24.13** [82:1]	5.01**	
						Gregory and Hansen cointegration tests				
						Z_t^*			Z_α^*		
						C	C/T	C/S	C	C/T	C/S
						-6.35** [73:3]	-7.37** [74:1]	-6.90** [77:1]	-65.4** [73:3]	-80.6** [74:1]	-71.6** [77:1]

Panel B: $RM1BP_t = u + \alpha_y y_t + \alpha_{umm-i^{bp}} (umm_t - i_t^{bp}) + [\beta trend] + v_t$

	FM estimates					Hansen stability tests					
(C)	u	y	$umm-i^{bp}$		trend	L_c	MeanF	SupF			
	- 8.07 (0.48)	1.64 (0.03)	-0.004 (0.010)			3.01**	5.19	13.51 [67:1]			
(D)	-11.88 (3.74)	1.99 (0.33)	-0.014 (0.009)		-0.008 (0.007)	1.18**	9.98	15.30 [87:3]			
						Gregory and Hansen cointegration tests					
						Z_t^*			Z_α^*		
						C	C/T	C/S	C	C/T	C/S
						-6.74** [73:3]	-8.22** [74:1]	-7.29** [87:2]	-67.6** [73:3]	-88.4** [74:1]	-77.2** [87:2]

See notes to Table 11.

Table 14
Cointegration tests with regime shifts and parameter constancy tests: real M2 data

Panel A: $RM2_t = u + \alpha_y y_t + \alpha_{i^{m2}} i_t^{m2} + \alpha_{umm} umm_t + [\beta trend] + v_t$

	FM estimates					Hansen stability tests					
(A)	u	y	i^{m2}	umm	trend	L_c	MeanF	SupF	Wald$(\alpha_{i^{m2}} + \alpha_{umm} = 0)$ [a]		
	-7.60	1.72	-0.011	-0.030		2.67**	16.43**	106.1**	6.96**		
	(0.21)	(0.02)	(0.020)	(0.007)				[67:1]			
(B)	5.51	0.54	-0.025	-0.010	0.020	3.36**	7.76**	20.21**	7.24**		
	(1.78)	(0.17)	(0.018)	(0.006)	(0.004)			[89:3]			
						Gregory and Hansen cointegration tests					
						Z_t^*			Z_α^*		
						C	C/T	C/S	C	C/T	C/S
						-7.83**	-5.84**	-9.07**	-89.0**	-57.0**	-106**
						[88:4]	[88:4]	[82:2]	[88:4]	[88:4]	[82:3]

Panel B: $RM2_t = u + \alpha_y y_t + \alpha_{umm} umm_t + [\beta trend] + v_t$

	FM estimates					Hansen stability tests					
(C)	u	y		umm	trend	L_c	MeanF	SupF			
	- 7.64	1.72		-0.030		1.48**	10.43**	16.49			
	(0.20)	(0.02)		(0.004)				[89:1]			
(D)	4.31	0.66		-0.020	0.022	1.58**	6.51	11.58			
	(2.17)	(0.19)		(0.004)	(0.004)			[88:3]			
						Gregory and Hansen cointegration tests					
						Z_t^*			Z_α^*		
						C	C/T	C/S	C	C/T	C/S
						-7.61**	-5.32**	-7.89**	-80.8**	-47.0	-87.5**
						[67:1]	[67:1]	[82:3]	[67:1]	[67:1]	[82:3]

Table 14 (cont'd)

Panel C: $RM2_t = u + \alpha_y y_t + \alpha_{umm-i^{m2}}(umm_t - i_t^{m2}) + [\beta trend] + v_t$										
	FM estimates				Hansen stability tests					
	u	y	$(umm-i^{m2})$	trend	L_c	MeanF	SupF			
(E)	- 7.77	1.71	-0.027		2.12**	18.31**	27.05**			
	(0.24)	(0.02)	(0.007)				[87:3]			
(F)	7.21	0.39	-0.019	0.028	2.86**	8.53**	19.01**			
	(2.20)	(0.19)	(0.006)	(0.004)			[88:3]			
					Gregory and Hansen cointegration tests					
					Z_t^*			Z_α^*		
					C	C/T	C/S	C	C/T	C/S
					-7.53**	-5.28	-7.65**	-79.7**	-48.84	-85.2**
					[67:4]	[67:4]	[83:1]	[67:4]	[67:4]	[82:3]

See notes to Table 11.

Table 15
Critical values for Z_t^* and Z_α^* tests

	m = 3		m = 2	
Test	5%	10%	5%	10%
Z_t^*				
C	-5.28	-5.02	-4.92	-4.69
C/T	-5.57	-5.33	-5.29	-5.03
C/S	-6.00	-5.75	-5.50	-5.23
Z_α^*				
C	-53.58	-48.65	-46.98	-42.49
C/T	-59.76	-54.94	-53.92	-48.94
C/S	-68.94	-63.42	-58.33	-52.85

Note: m is the number of I(1) regressors. Critical values for Z_t^* and Z_α^* are taken from Gregory and Hansen (1996).

Table 16:
Critical values for L_c, MeanF, and SupF, tests

	m = 3, p = 1		m = 2, p = 1	
	5%	1%	5%	1%
L_c	0.90	1.29	0.778	1.13
MeanF	9.21	12.00	7.69	10.3
SupF	19.3	23.9	17.3	21.4

Note: m is the number of I(1) regressors; p: the power of trend in the data. Critical values are from Hansen (1992b).

Model C/S: Regime Shift

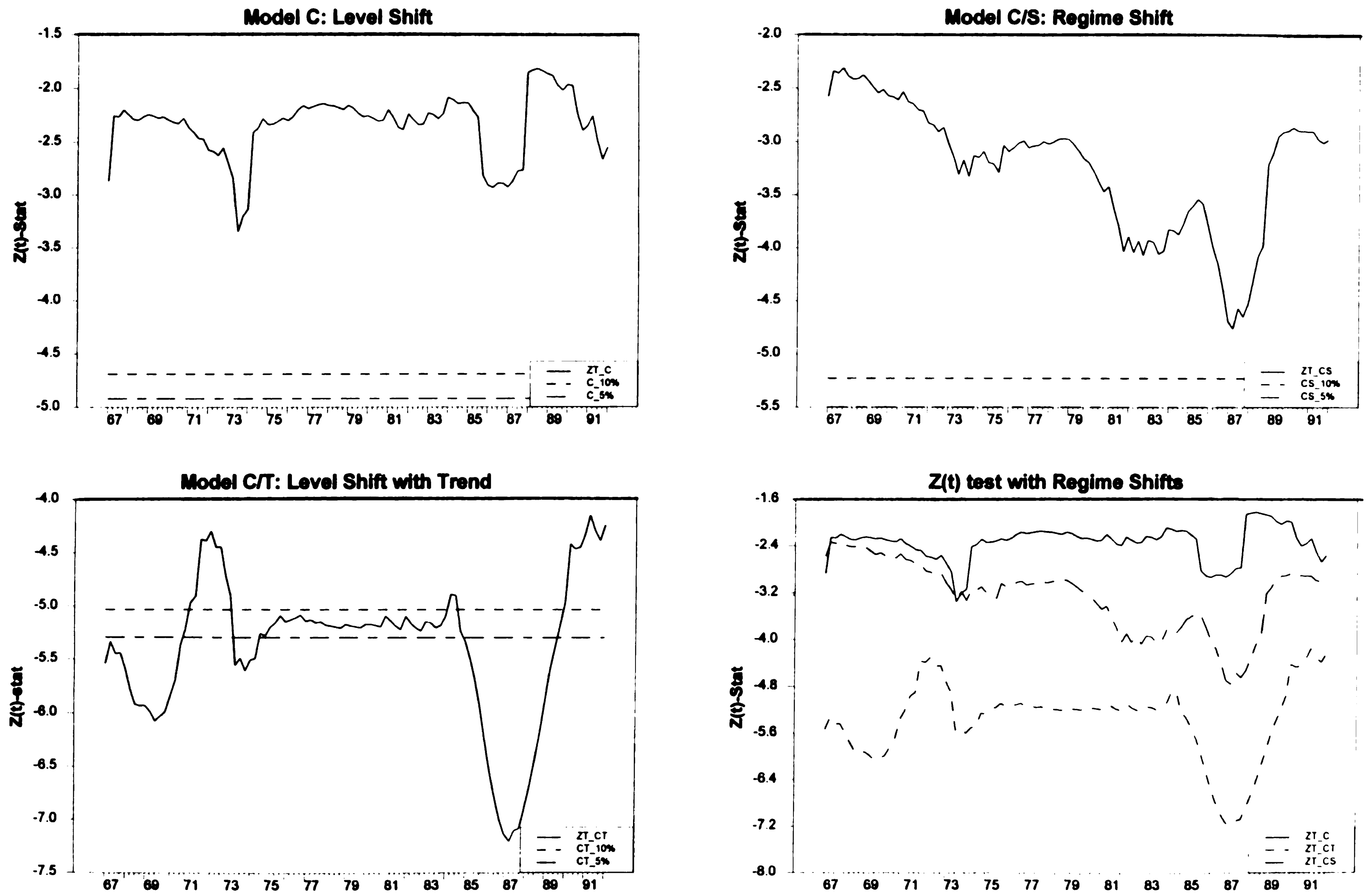

Figure 3: Time plot of Z(t) statistics for the demand for real M1A

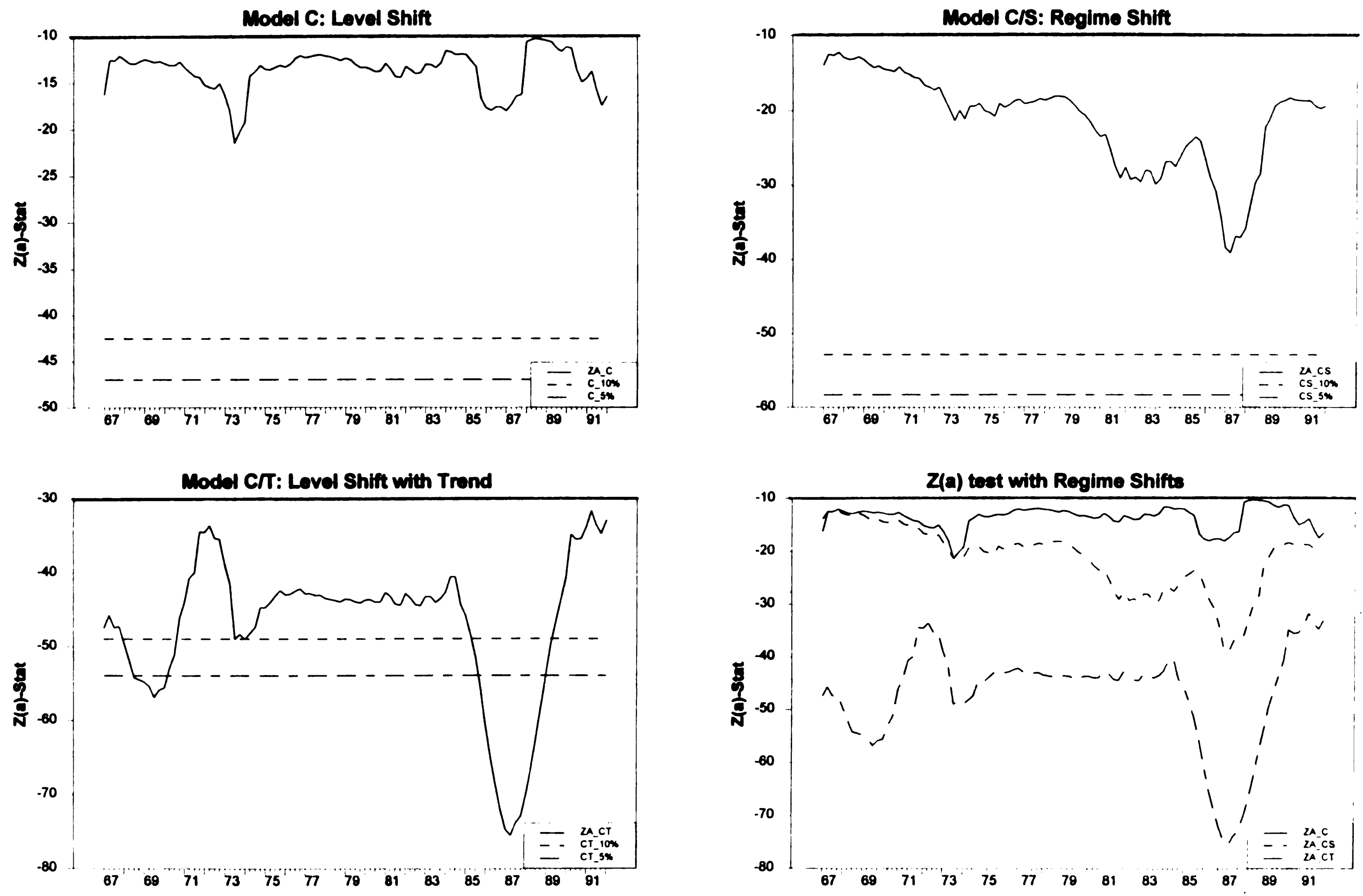

Figure 4: Time plot of Z(a) statistics for the demand for real M1A

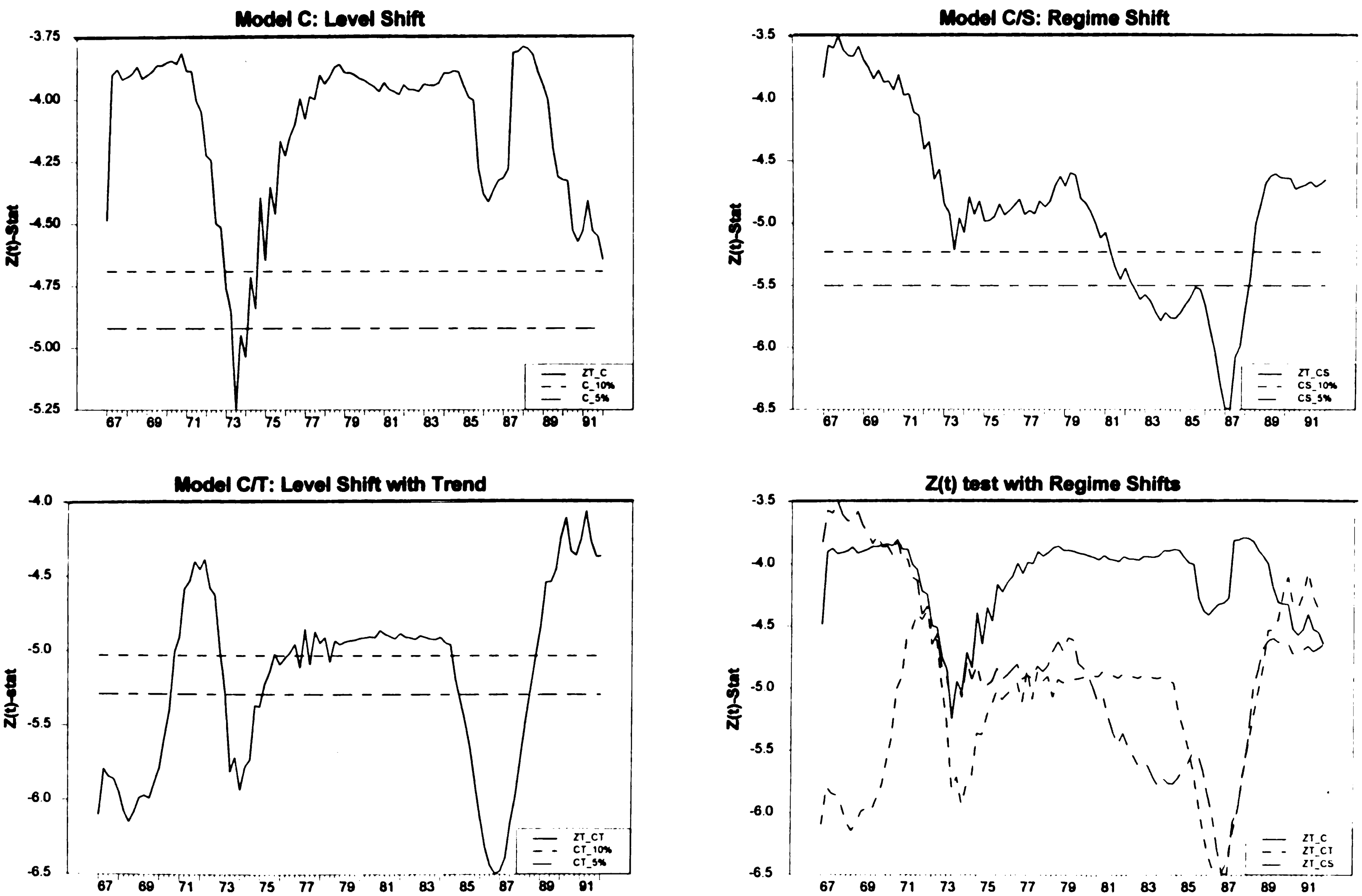

Figure 5: Time plot of Z(t) statistics for the demand for real M1B

27.5

Model C: Level Shift

Model C/S: Regime Shift

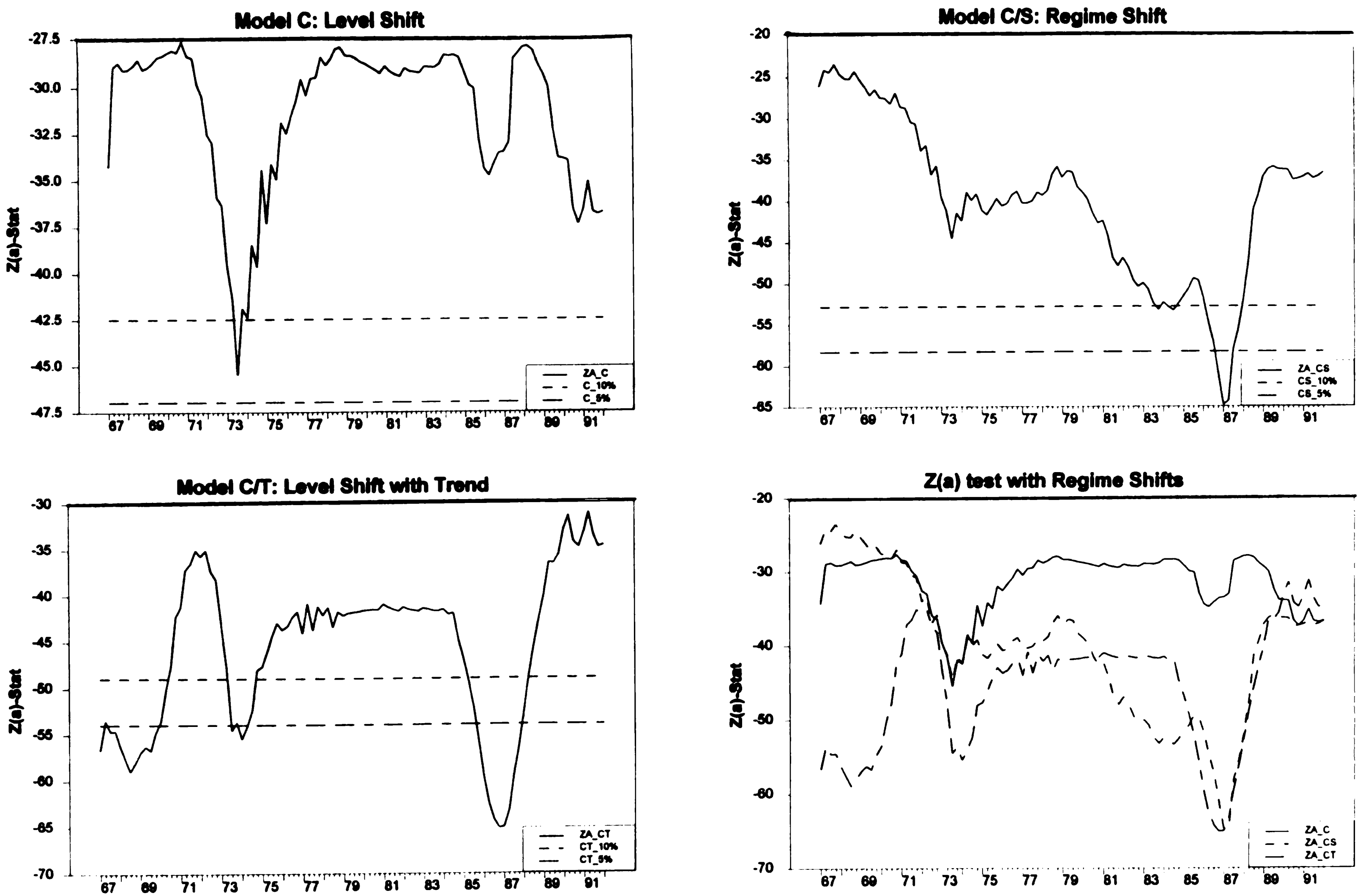

Figure 6: Time plot of Z(a) statistics for the demand for real M1B

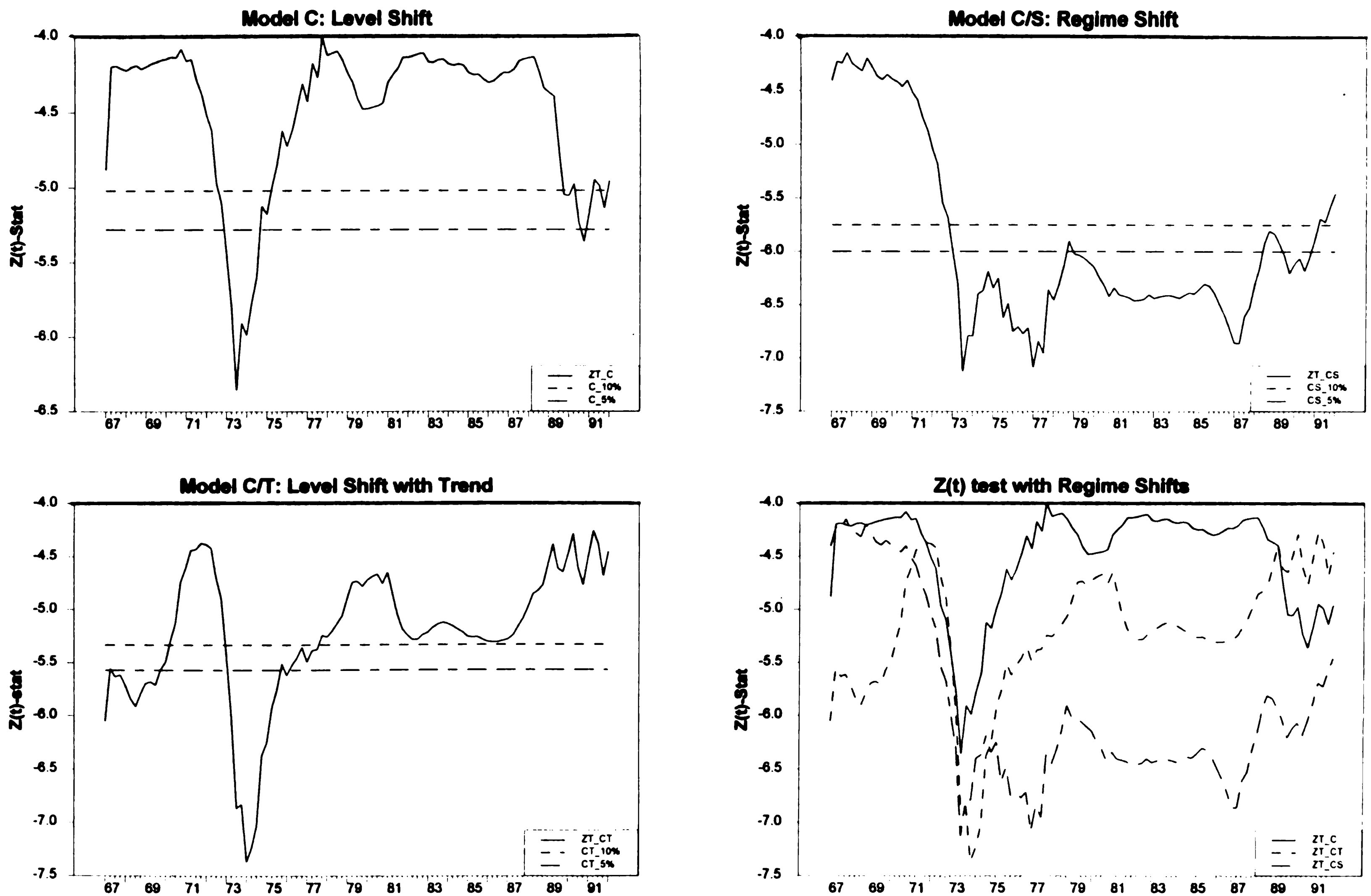

Figure 7: Time plot of Z(t) statistics for the demand for real M1BP

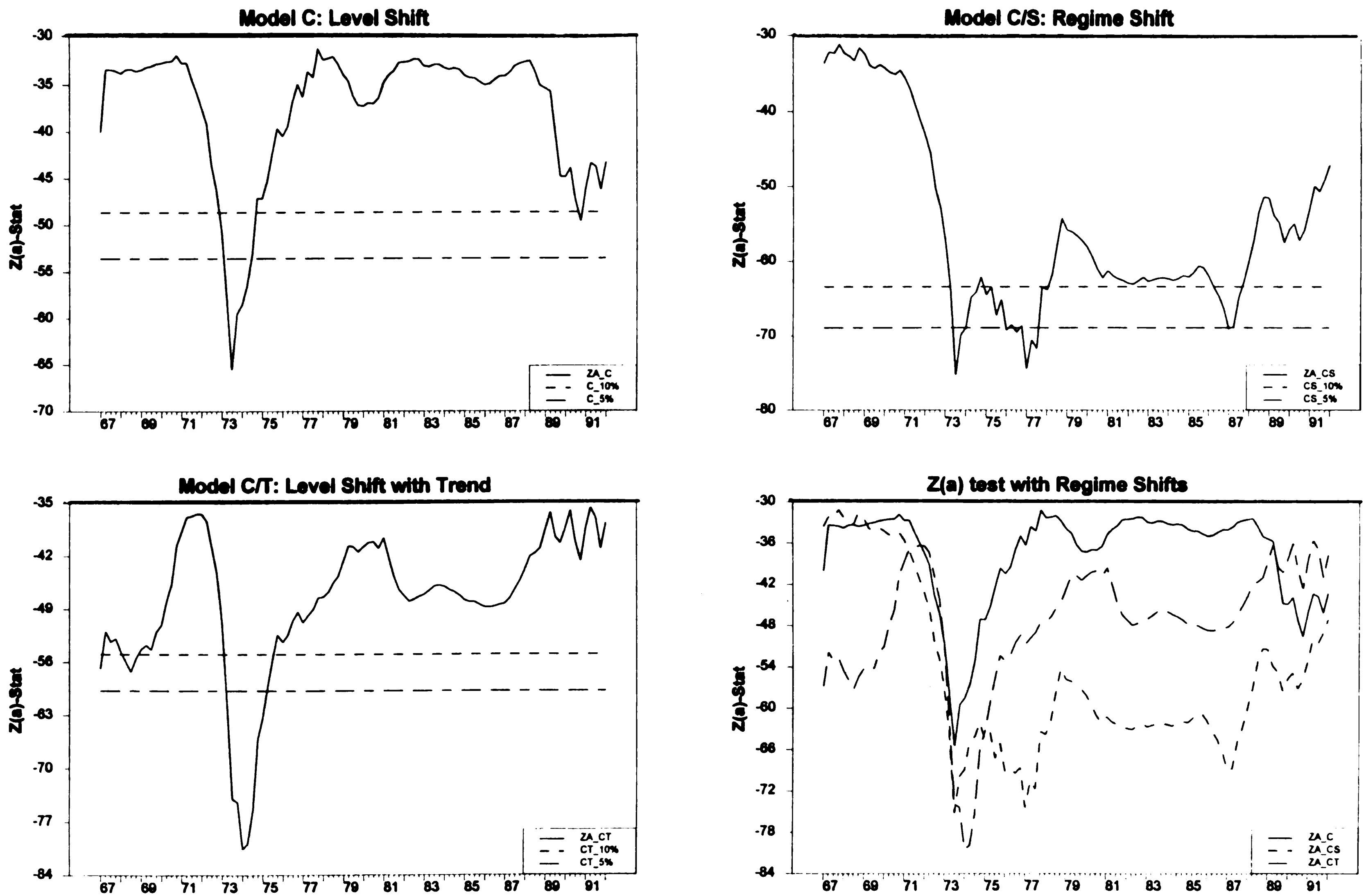

Figure 8: Time plot of Z(a) statistics for the demand for real M1BP

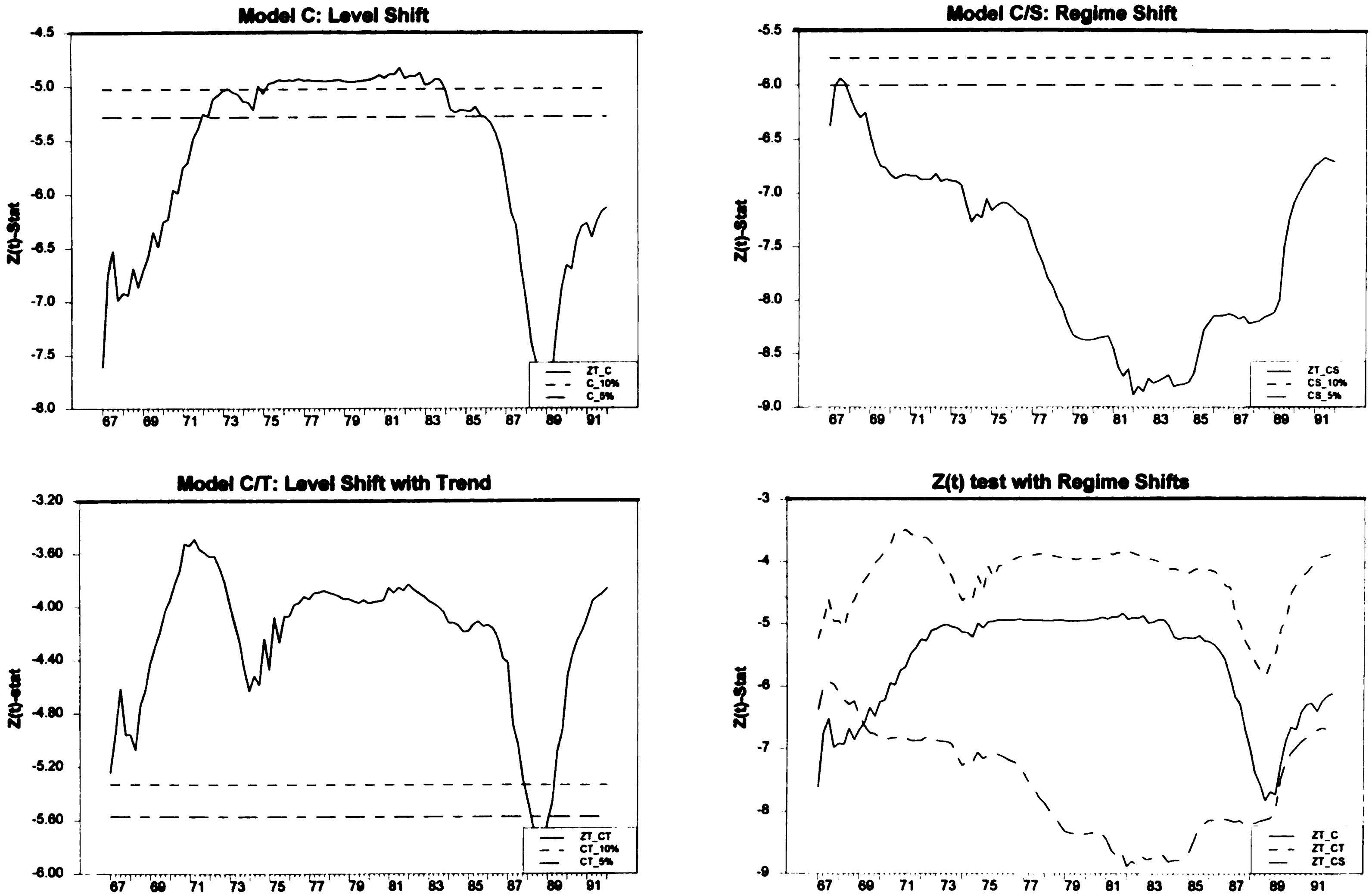

Figure 9: Time plot of Z(t) statistics for the demand for real M2

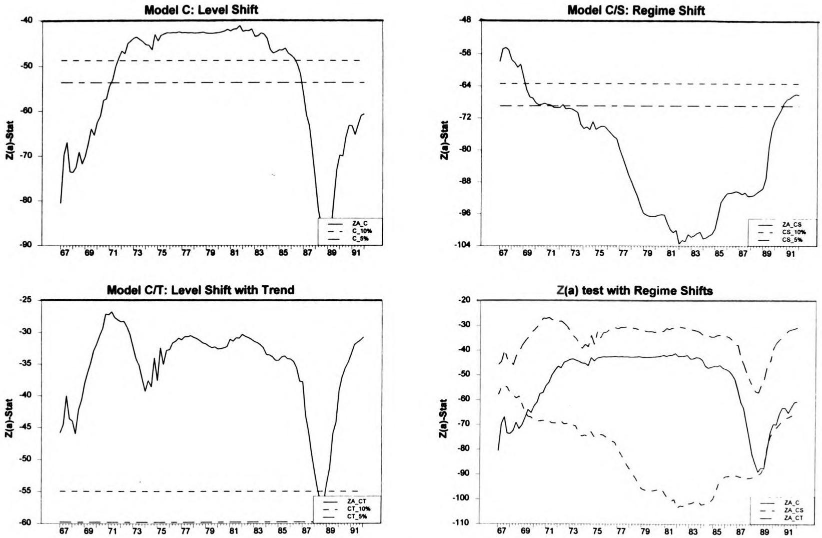

Figure 10: Time plot of Z(a) statistics for the demand for real M2

SupF test for Demand for real M1A

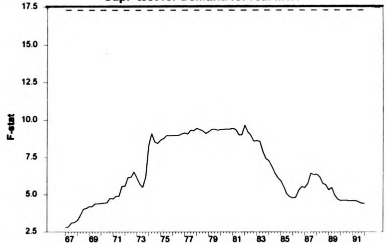

Figure 11

SupF test for Demand for real M1BP

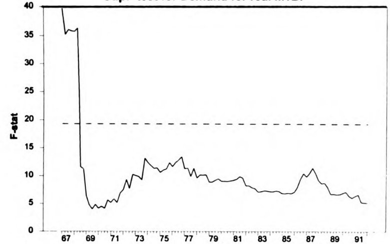

Figure 13

SupF test for Demand for real M1B

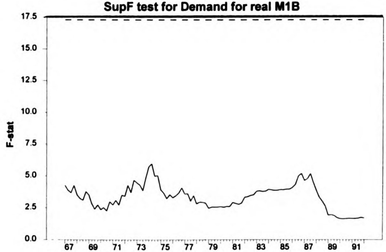

Figure 12

SupF test for Demand for real M2

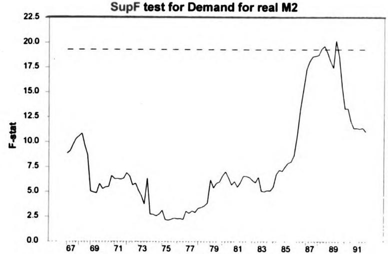

Figure 14

IV: JOHANSEN'S FULL INFORMATION MAXIMUM LIKELIHOOD ESTIMATION

1 *Introduction*

From the residual based cointegration tests in the Section II and III, we have found evidence of cointegration with stable coefficients between the variables in the demand for real M1A and real M1B equations. However, we cannot determine from those tests whether the cointegrating relation is unique or whether there are other linearly independent cointegrating vectors in the system. Approaches other than residual-based tests for cointegration are available, for example, the likelihood ratio tests of cointegration rank of Johansen (1988b, 1991), Johansen and Juselius (1990), and a common stochastic trends test proposed by Stock and Watson (1988). Such tests developed from systems methods enable researchers to avoid invalid restriction from arbitrary normalization in the cointegrating regression of the residual-based tests[19]. For example, the variable on the left hand side of the regression may not appear in the cointegrating relation at all, but a unitary coefficient is wrongly imposed on this variable. These tests also have the advantage of testing for the number of cointegrating relations. Since the Johansen

[19] Conflicts may arise in empirical work where the test outcome depends on the normalization selected.

(1988b, 1991) maximum likelihood methods for the analysis of cointegration can simultaneously detect the number of the cointegration rank in the system, estimate and test for linear hypothesis about the cointegrating vectors and their adjustment coefficients, we will apply them to continue our study of long-run money demand equations in Taiwan.

Testing for the number of cointegrating relations and estimating the cointegrating vectors starts with a VAR(p) representation expressed in first order difference and lagged levels

H_1 :

$$\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots\cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-1} + u_0 + u_1 t + D_t + \varepsilon_t \tag{3.19}$$

where x_t is a p-dimensional vector of I(1) variables; D_t are centered seasonal dummies which sum to zero over a full year and where $\varepsilon_1, \cdots\cdots, \varepsilon_T$ are $IIN_p(0,\Lambda)$ and $x_{-k+1} \cdots\cdots x_0$ are fixed. The Π matrix conveys the long-run information in the data. The hypothesis of r cointegrating vectors is formulated as a reduced rank of the Π-matrix

$$H_2(r): \Pi = \alpha\beta' \tag{3.20}$$

where α and β are pxr matrices of full rank.

Under H_2 (r): $\Pi = \alpha\beta'$, (3.19) can be interpreted as an error correction model (see Engle and Granger 1987, and Johansen 1988a).

2 Empirical Study: The Number of the Cointegrating Vectors in Taiwanese Money Demand Data (1961:4 - 1997:3)

We apply Johansen's maximum likelihood method to the study of the long-run money demand relations (M1A and M1B)[20] in Taiwan. For a comparison with the results in the previous sections (II and III), we consider both 4-dimensional VAR (where x = (m,y,i,umm)) and 3-dimentional VAR (where x = (m,y,umm-i) for the basic model for each money-demand relation, where m denotes the log of real balances.

Since the intercept vector (μ_0) can be decomposed into linear trends in the data and intercepts in the cointegrating relationship (Johansen and Juselius 1990, Johansen 1991b) and the trend coefficient μ_1 can be decomposed into quadratic trend coefficients in the data and linear trend coefficients in the cointegrating relations (see Osterwald-Lenum 1992), the assumptions about the properties of the trend and linear restrictions about μ_0

[20] We also consider the order of the cointegration rank for the other two aggregates (M1BP and M2). We cannot reject the hypothesis of no cointegraion (r=0) at either 5% or 10% even 20% level from the trace and max-λ tests. These add more evidence against a stable long-run relationship for

and μ_1 are important in distinguishing the correct asymptotic distributions of the test statistics.

Since all of the variables except for the interest rates are characterized as I(1) processes with drift, we estimate all of the models under the assumption of a linear trend in the I(1) variables. Specifically, (3.19) without a time trend and centered seasonal dummies[21] is fitted to the Taiwanese demand for real M1B data

Model I:

$$H_2: \Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots\cdots + \Gamma_{k-1}\Delta x_{t-k+1} + \Pi x_{t-k} + u_0 + \varepsilon_t \tag{3.21}$$

Because the time trend is crucial to our conclusion of cointegration with stable parameters for the demand for real M1A data from the single equation methods, we also consider model (3.19) under the assumption of the existence of linear trend in the cointegrating relations (but the absence of quadratic trend) to take into account of a significant

---

these two aggregates. We do not report the results here but the results are available upon request.

[21] Although our data are not seasonally adjusted, we do not include the centered seasonal dummies in the model, because the results obtained in the regression with seasonal dummies are similar to those we report in the text. In addition, from the fully modified estimation we found three seasonal dummies are not significant either individually or jointly for each demand for money equations.

linear trend in FM estimation[22]. That is, under the null we estimate

Model II : H_2^*:

$$\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \alpha(\beta', \beta_1)(x_{t-1}', t)' + u_0 + \varepsilon_t \tag{3.22}$$

(see Osterwald-Lenum 1992).

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

(1) <u>The Taiwanese demand for real M1A data</u>

(A) <u>Misspecification tests</u>

The misspecification tests for the model are reported in Table 17. For the two data sets (three- and four-variable), the choice of k renders the residuals white noise. However, there are indications of excess skewness and kurtosis in the residuals of the interest rate equations, causing the Jarque-Bera test statistic to become

[22] Explicit inclusion of a linear trend in the cointegrating relations implies that there is some linear growth which our model cannot account for given the chosen data set.

significant. We conclude that the residuals from RM1B and y can be assumed to follow a Gaussian process, those from interest rate equations follow an innovation process[23].

(B) Testing for reduced rank

For VAR(4), Model I (3.21) is fitted to the data[24]; for VAR(3) we consider both Model I and II (3.22).

From Table 18, the trace and λ_{max} tests fail to reject $r = 0$ at either 5%` and 10% level even 20% level (significant at 20% for VAR(4)) for all of the model we considered. Since the time trend is important to the conclusion of cointegration with stable coefficients from the residual based tests (Table 6 and 11), the inability to reject H_2^* in VAR(3) adds more evidence against the hypothesis of cointegration with constant coefficients among the variables in the demand for real M1A. Using a single interest rate, umm, we reach the same conclusion of no

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[23] The deviation from normality is not a serious problem. As long as the cumulative sums of errors converge to a Brownian motion, the asymptotic analysis gives the same results as those under the assumption of normality (see Johansen 1991b, Johansen and Juselius 1992, Gonzalo 1994).

[24] We also consider model II under the assumption of a linear trend in the cointegration relations. Although the trace and λ_{max} tests reject $r = 0$ at the 5% level, the estimated income elasticity and interest semi-elasticity are perverse, having wrong signs and being economically meaningless. $(\hat{\beta}_m, \hat{\beta}_y, \hat{\beta}_{i^b}, \hat{\beta}_{umm}, \hat{\beta}_1)$ = (1.00, 12.28, -5.08, -0.28, -0.28).

cointegration which also differs from the conclusion we reached in previous sections (II and III).

Table 17
Residual misspecification tests

VAR(4) with k = 8, x = (RM1A, y, i^a, umm)					
Eq.	S.E.E[a]	Sk[b]	EK[c]	τ_1[d]	τ_2[e] ~ $\chi^2(34)$
ΔRM1A	0.03	-0.33	0.14	1.97	34.18
Δy	0.01	0.21	-0.04	0.77	37.44
Δi^a	0.05	0.37	5.37**	132.37**	45.76
Δumm	0.58	1.06**	3.51**	75.62**	23.22

VAR(3) with k = 5, x = (RM1A, y, umm-i^a)					
Eq.	S.E.E	Sk	EK	τ_1	τ_2 ~ $\chi^2(34)$
ΔRM1A	0.03	-0.14	0.26	0.70	42.00
Δy	0.02	0.13	-0.03	0.36	42.56
Δumm-i^a	0.66	0.89**	5.19**	150.39**	26.78

Note: a. S.E.E denotes the standard error of regression estimate. b. SK and EK are the skewness and kurtosis statistics. A ** indicates the test of the null hypothesis that each is zero is rejected at the 5% level. d. The Jarque and Bera test for normality (Jarque and Bera, 1980), $\tau_1 = \frac{T-m}{6}(SK^2 + \frac{EK^2}{4}) \sim \chi^2(2)$, where m is the number of regressors. C. The Ljung-Box Q-statistic $T(T+2)\left(\sum_{j=1}^{M}\frac{r_j^2}{T-j}\right)$, where r_j is the jth lag autocorrelation of the residuals. M is the number of autocorrelations used, and is selected according to the formula M = min(T/4, 3T^2), with a maximum value of 36.

Table 18
Tests of the cointegration rank

Panel A
Model I: VAR(4)[1] with k = 8, x = (RM1A, y, i^a, umm)

H_2	eigenvalue	trace	λ_{max}
$r = 0$	0.131	41.90	19.07
$r \leq 1$	0.095	22.83	13.51
$r \leq 2$	0.041	9.31	5.74
$r \leq 3$	0.026	3.57	3.57

Panel B
Model I: VAR(3)[2] with k = 5, x = (RM1A, y, umm-i^a)

H_2	eigenvalue	trace	λ_{max}
$r = 0$	0.092	23.92	13.41
$r \leq 1$	0.055	10.52	7.94
$r \leq 2$	0.018	2.58	2.58

Panel C
Model II: VAR(3)[2] with k = 5, x = (RM1A, y, umm-i^a)

H_2^*	eigenvalue	trace	λ_{max}
$r = 0$	0.107	30.12	15.67
$r \leq 1$	0.064	14.45	9.25
$r \leq 2$	0.037	5.20	5.20

Note: 1 Model I: $\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-k} + \mu_0 + \varepsilon_t$ is estimated.

2 Under the null of cointegration rank = r, Model II $\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \alpha(\beta', \beta_1)(x_{t-1}', t)' + \mu_0 + \varepsilon_t$ is assumed.

See Table 23 for critical values.

To conclude the cointegration analysis for the demand for real M1A, we represent the normalized eigenvector associated with the largest eigenvalue and the corresponding weight in Table 19. We notice that the coefficients in panel (c), under the assumption of a linear trend in the cointegrating relation, the obtained income elasticity is not only greater than unity, is also greater than two, as mentioned above, which is rarely seen in the literature for this narrow monetary aggregate.

The time trend is important to the conclusion of a stable cointegrating relationship in the demand for real M1A from single equation methods (Section II and III). To accept this conclusion, we need to accept the huge income elasticity. Based on economic theories and all the evidence we have obtained as well, we conclude that there is no such long-run relationship for the demand for real M1A data.

Table 19

The eigenvector associted with the largest eigenvalue ($\hat{v}_1$) and the corresponding weights ($\hat{w}_1$)

Panel A

Model I: VAR(4)[1] with k = 8, x = (RM1A, y, i^a, umm)

	m	y	i^a	umm
$\hat{v}_1$	1.00	-1.27	-0.76	-0.014
	(0.21)	(0.27)	(0.17)	(0.008)
$\hat{w}_1$	0.01	0.05	0.03	1.15

Panel B

Model I: VAR(3) with k = 5, x = (RM1A, y, umm-i^a)

	m	y	umm-i^a
$\hat{v}_1$	1.00	-1.26	0.003
	(0.14)	(0.24)	(0.26)
$\hat{w}_1$	-0.10	0.001	1.36

Panel C

Model II: VAR(3)[2] with k = 5, x = (RM1A, y, umm-i^a)

	m	Y	umm-i^a	t
$\hat{v}_1$	1.00	-2.33	0.019	0.023
	(0.04)	(0.42)	(0.008)	(0.009)
$\hat{w}_1$	-0.13	0.01	0.83	

See notes to Table 18.
Figures in ( ) are standard errors computed using the procedure suggested by Johansen (1991).

(2) The Taiwanese demand for real M1B data

(A) Misspecification tests

The misspecification tests for this model are reported in Table 20. As in the demand for real M1A data, the normality hypothesis is rejected for the interest rate variables (i^h, umm, umm-i^h) due to excess skewness and kurtosis. Thus, the residuals from real RM1B and y are assumed to follow a Gaussian process and the residuals from the interest rates follow an innovation process.

Table 20
Residual misspecification tests

	VAR(4) with k = 7, x = (RM1B, y, i^h, umm)				
Eq.	S.E.E	SK	EK	τ_1	τ_2 $\chi^2(34)$
ΔRM1B	0.03	-0.28	0.81	4.18	29.73
Δy	0.02	0.13	-0.03	0.30	53.06
Δi^h	0.09	0.47**	3.26**	51.80**	31.32
Δumm	0.63	0.72**	4.07**	84.04**	21.30
	VAR(3) with k = 5, x = (RM1B, y, umm-i^h)				
ΔRM1B	0.03	-0.25	0.72	3.85	45.82
Δy	0.02	0.09	-0.08	0.19	41.09
Δumm-i^h	0.65	0.85**	5.20**	149.61**	26.61

Note: Model $\Delta x_t = \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{k-1} \Delta x_{t-k+1} + \Pi x_{t-k} + \mu_0 + \varepsilon_t$ is fitted to the data.

(B) Testing for reduced rank

From the trace and λ_{max} test results in Table 21, the null hypothesis of no cointegration (r=0) cannot be rejected at either 5% or 10% level for the two data sets. However, the null is rejected at the 20% level, which appears consistent with the results from the residual-based cointegration tests (see Table 7 and 12). Although the acceptance of one cointegration vector in the system relies on the test size of 20%, which usually would be considered too high, Johansen and Juselius 1990 suggests that when the cointegrating relation is quite close to the nonstationary boundary, the powers of the test are likely to be low. Hence it seems reasonable to use a higher critical value higher than the usual 5%.

In order to make final decision of the order of the cointegration rank, we also examine the stationarity of the cointegrating relation ($\hat{\beta}x_t$) constructed from the full-sample point estimates. The ADF-t statistic (= -2.64 for VAR(4), = -3.22 for VAR(3)) (using 9 lagged differences) rejects the unit root hypothesis at the 10% level for VAR(4) model, at the 5% level for VAR(3), suggesting that the cointegrating relations for these two data sets are stationary.

Since we reject r = 0 and we cannot reject the hypothesis of $r \leq 1$ at 20% level, we conclude that there is only one cointegrating vector between the variables in the models. Note in APPENDIX 4, we show that there is no cointegration between umm and i^h. Thus, we rule out the possibility that this cointegrating relationship comes from these two variables alone. When using the umm rate alone, we obtain similar results as those from the model using the spread. Therefore, the result is consistent with the conclusion we drawn in Section III.

Table 21
Trace and λ_{max} tests

VAR(4) with k = 7, x = (RM1B, y, i^h, umm)			
H_2	eigenvalue	trace	λ_{max}
r = 0	0.116	40.45	16.84
r ≤ 1	0.095	23.60	13.73
r ≤ 2	0.052	9.87	7.26
r ≤ 3	0.019	2.61	2.61

VAR(3) with k = 5, x = (RM1B, y, umm-i^h)			
H_2	eigenvalue	trace	λ_{max}
r = 0	0.095	24.60	13.91
r ≤ 1	0.057	10.69	8.23
r ≤ 2	0.018	2.46	2.46

See note to Table 20. See Table 23 for critical values.

Next we report the normalized cointegrating vector[25] and the error correction coefficients in Table 22. The standard errors of the normalized cointegrating vectors are obtained using the procedure suggested by Johansen (1991). We will use the obtained point estimates and their standard errors to test the significance of each regressor.

For the VAR(4) model, the signs of coefficients on income and the own rate are consistent with the prediction of the theories and both are significantly different from zero. The sign on umm rate is consistent with an conjecture that the umm rate is a measure of the opportunity cost of holding real M1B in Taiwan, however, it is not significantly different from zero. Since the constraint ($\hat{\beta}_{i^h} = -\hat{\beta}_{umm}$)[26] cannot be rejected by likelihood ratio test (LR = 1.37), so we reestimate the model. We find that not only that the interest rate spread is significant, but also that the trace test turns out to reject the r = 0 at the 10% level. All

[25] We present the results normalized by the coefficients of real M1B variable such that $\hat{\beta}_{m} = 1.00$. The parameters α and β are not identified since given any choice of the matrix $\xi(rxr)$, $(\alpha\xi)$ and $\beta(\xi')^{-1}$ also produces the same matrix Π, and hence determine the same probability distribution for the variables. The data can only determine the space spanned by the columns in β, and the space spanned by α. For identification of long-run parameters, see Rasche and Hoffman (Chapter 3, 1996).

estimates have smaller standard deviations than unrestricted estimates. The constrained estimated are reported in the Panel B of Table 22.

Turning to the error correction coefficients (α), we can see that equilibrium errors cause real M1B and the umm rate to adjust downward, real GNP and the own rate to adjust upward. We also notice that the weight of the equilibrium errors loading to the change of real balances is very small in unrestricted VAR(4) model (Panel A, Table 22). In both models, the equilibrium error enters the umm equation with largest weights (in absolute value), suggesting that the cointegration relation is important to the adjustment of the unorganized money market rates.

For the VAR(3) model, the income elasticity (= 1.56) is significantly different from zero and unity. The interest semi-elasticity (= -0.017) has a negative sign as conjectured but becomes insignificant. The error correction coefficients suggest that the equilibrium errors cause real M1B to adjust downward, real GNP and the spread to adjust upward.

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[26] The restriction is imposed on the long-run information but not on the short-run dynamics, (see Hoffman, Rasche, and Tieslau 1995).

Table 22

Normalized cointegrating vectors ($\hat{\beta}$) and error correction coefficients ($\hat{\alpha}$)

Panel A				
Unrestricted VAR(4) with k = 7, x = (RM1B, y, i^b, umm)				
	$\hat{\beta}_m$	$\hat{\beta}_y$	$\hat{\beta}_{i^b}$	$\hat{\beta}_{umm}$
$\hat{\beta}$	1.00	-1.45	-0.35	0.05
	(0.23)	(0.32)	(0.07)	(0.07)
Eq.	ΔRM1B	Δy	Δi^b	Δumm
$\hat{\alpha}$	-0.003	0.04	0.16	-0.29
Panel B				
Restricted VAR(4) with $\hat{\beta}_{i^b} = -\hat{\beta}_{umm}$ (LR = 1.37)[1]				
	$\hat{\beta}_m$	$\hat{\beta}_y$	$\hat{\beta}_{i^b}$	$\hat{\beta}_{umm}$
$\hat{\beta}$	1.00	-1.61	-0.02	0.02
	(0.14)	(0.20)	(0.01)	(0.01)
Eq.	ΔRM1B	Δy	Δi^b	Δumm
$\hat{\alpha}$	-0.03	0.04	0.07	-0.72
Panel C				
VAR(3) with k = 5, x = (RM1B, y, umm-i^b)				
	$\hat{\beta}_m$	$\hat{\beta}_y$	$\hat{\beta}_{umm-i^b}$	
$\hat{\beta}$	1.00	-1.56	0.017	
	(0.27)	(0.23)	(0.090)	
Eq.	ΔRM1B	Δy	Δ(umm-i^b)	
$\hat{\alpha}$	-0.12	0.01	0.41	

Note: Figures ( ) are standard deviations. Also see notes to Table 20. LR is the log likelihood ratio test stastistic for $\hat{\beta}_{i^b} = -\hat{\beta}_{umm}$ against H_2 (Eq.3.20), which is asymptotically distributed as $\chi^2_{0.95}(1) = 3.84$.

Table 23
Critical values for trace and λ_{max}

p-r	80%	90%	95%	80%	90%	95%
		Model I Trace			Model II Trace	
1	1.66	2.69	3.76	8.65	10.49	12.25
2	11.07	13.33	15.41	20.19	22.76	25.32
3	23.64	26.79	29.68	35.56	39.06	42.44
4	40.15	43.95	47.21	54.80	59.14	62.99
		λ_{max}			λ_{max}	
1	1.66	2.69	3.76	8.65	10.49	12.25
2	10.04	12.07	14.07	14.70	16.85	18.96
3	16.20	18.60	20.97	20.45	23.11	25.54
4	21.98	24.73	27.07	26.30	29.12	31.46

Note: Critical values are taken from Table 1 and Table 2* in Osterwald-Lenum 1992.

V: ESTIMATION AND TESTS OF LINEAR RESTRICTIONS ON THE LONG-RUN PARAMETERS OF DEMAND FOR REAL M1B IN TAIWAN

1 Estimation of the Long-Run Income and Interest Elasticities

The long-run demand for money plays an important role in the quantitative analysis of the effects of monetary policy. Much of the empirical money demand literature has focused on the search for a stable short-run money demand function. In such cases, the efficient estimators can be used in subsequent stages of the analysis by imposing the estimated cointegrating vectors, for example in a VECM (King, Plosser, Stock and Watson 1991), or in a single-equation error correction framework (Hendry and Ericsson 1991). In this section, we compare estimates of the long-run parameters for the demand for real M1B using four asymptotic efficient estimators, OLS (Engle and Granger 1987, Stock 1987), the fully modified estimators of Phillips and Hansen (1990) (FM) and Johansen's (1988b) VECM maximum likelihood estimator (JOH) and Stock and Watson (1993) DOLS. The last three estimators have an asymptotic distribution that is a random mixture of normals and produce Wald test statistics with asymptotic chi-squared distribution.

In each estimation, again we consider both four-variable system x_t= (RM1B, y, i^h, umm) and three-variable system x_t= (RM1B, y, umm-i^h). Estimates of the cointegrating vectors are normalized such that $\hat{\theta}_m$ = -1.00 and are reported in Table 24. We use the obtained point estimates and their standard errors to test the significance of each regressors.

In the VAR(4) model, the estimates are very close across the estimators and the coefficients on each variable have the expected signs. The income elasticity is around 1.50 and is significantly different from zero and unity at either 5% or 10% (unrestricted JOH in VAR(4)) level for all estimators. The interest semi-elasticity (the own rate $\hat{\theta}_{i^h}$) is significantly from zero only for the Johansen estimator, which is largest among the estimators. $\hat{\theta}_{umm}$ is similar across estimators and significantly different from zero except in Johansen's estimation. Under the constraint β_{i} = $-\beta_{umm}$, which cannot be rejected by the data (LR = 1.37, see Panel B, Table 22), the interest rate spread in Johansen's estimator becomes significant.

In VAR(3), the income elasticity and interest semi-elasticity are almost identical across estimators. Surprisingly, Johansen estimates have the largest standard deviation among all the estimators. The interest variable is not significant in the Johansen estimation at either 5% or 10% level. Under the constraint β_y = -1.50, the interest semi-elasticity from JOH becomes more precisely estimated (with much smaller standard deviation) and turns out to be significant. Since the own rate is not significant across the estimators except JOH, in Table 25, we also report the results obtained in the regressions only using the umm rate. We find that the estimates are similar to those using the spread (the right three columns of Table 24). In all cases, θ_i is significantly different from unity at the 5% level; θ_{umm} is significantly different from zero at either 5% and 10% level.

From Table 24, although all four asymptotic efficient estimators generate qualitatively and statistically close estimates, Johansen estimator generally produces larger standard deviations among all estimators considered. Surprisingly, the analysis based on a complete model (Johansen 1991), where the different variables are jointly modeled and a full information analysis is pursued, does not have efficiency gains in estimating the long-run parameters[27]. This is consistent with the simulation results in Stock and Watson (1993). Our study provides an empirical example that single equation methods can also efficiently estimate long-run equilibria, as advocated in Phillips and Loretan (1991), Stock and Watson (1993).

To conclude the estimation of the long-run demand for real M1B in Taiwan, we present the graphs for the actual real M1B (denoted RLM1B in figures) and the estimated real M1B generated by the estimates of Table 24 to see how well these regressions track the data. Figure 15 and 16 plot the comparison of the estimated values across estimators for the four-variable system (denoted OLS4, DOLS4, FM4, JOH4, REJOH4) and three-variable system (denoted OLS3, DOLS3, FM3, JOH3), respectively. For each estimator, we also compare the estimated real M1B obtained from the four- and three-variable system with the actual real M1B; they are presented in Figure 17 - 20.

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[27] I thank Professor Wooldridge for the comments on this empirical finding.

From Figure 15 - 20, we can see all the estimators except unrestricted JOH (JOH4) track the data well either we use two interest rates or the spread in the regression. We notice that the actual M1B/P is above the estimated value across estimation during 1986-91; this period coincides the boom of the stock exchange market in Taiwan.

Table 24
Estimated Cointegrating relations:
$\mathbf{RM1B}_t = \mu + \theta_y \mathbf{y}_t + \theta_i i_t + \varepsilon_t$

			OLS			
	x = (RM1B, y, i^b, umm)			x = (RM1B, y, umm-i^b)		
$\hat{\mu}$	$\hat{\theta}_y$	$\hat{\theta}_{i^b}$	$\hat{\theta}_{umm}$	$\hat{\mu}$	$\hat{\theta}_y$	$\hat{\theta}_{umm-i^b}$
-6.50	1.52	0.05	-0.021	-6.80	1.55	-0.023
(0.29)	(0.02)	(0.03)	(0.004)	(0.13)	(0.01)	(0.003)
			DOLS[1]			
-6.21	1.50	0.10	-0.025	-6.89	1.55	-0.017
(1.27)	(0.09)	(0.14)	(0.016)	(0.65)	(0.04)	(0.012)
			FM[2]			
-6.42	1.53	0.06	-0.025	-6.81	1.55	-0.017
(0.89)	(0.07)	(0.09)	(0.010)	(0.48)	(0.03)	(0.010)
			JOH			
		VAR(4)			VAR(3)	
-5.05	1.45	0.35	-0.050	-6.92	1.56	-0.017
	(0.32)	(0.07)	(0.075)		(0.23)	(0.092)
Restricted VAR(4) with $\beta_{i^b} = -\beta_{umm}$				VAR(3) with $\beta_y = -1.50$		
-7.57	1.61	0.02	-0.02	-6.83	1.50	-0.016
	(0.20)	(0.01)	(0.01)		(0.28)	(0.004)

Note: 1 DOLS estimation regression:

$$RM1B = \mu + \theta_y y_t + \theta_i i_t + \sum_{j=-k}^{k} d_j L^j \Delta y_t + \sum_{j=-k}^{k} g_j L^j \Delta i_t + \varepsilon_t$$

where i is a vector containing interest rates. We set the number of leads and lags k = 2 and use AR(2) error process to implement DOLS covariance matrix and estimate the standard errors.
2 The frequency zero spectral estimator required for FM are computed using prewhitened quadratic spectral kernel and automatic bandwidth estimator(see Section II and III this Chapter). Figures in parenthesis are standard deviations.

Table 25
Estimated Cointegrating relations:
$RM1B_t = \mu + \theta_y y_t + \theta_{umm} umm_t + \varepsilon_t$

$\hat{\mu}$	$\hat{\theta}_y$	$\hat{\theta}_{umm}$
	OLS	
-6.96	1.56	-0.016
(0.12)	(0.01)	(0.003)
	DOLS	
-7.04	1.56	-0.015
(1.47)	(0.04)	(0.011)
	FM	
-6.96	1.56	-0.015
(0.44)	(0.03)	(0.009)
	JOH	
-7.05	1.57	-0.016
	(0.26)	(0.008)

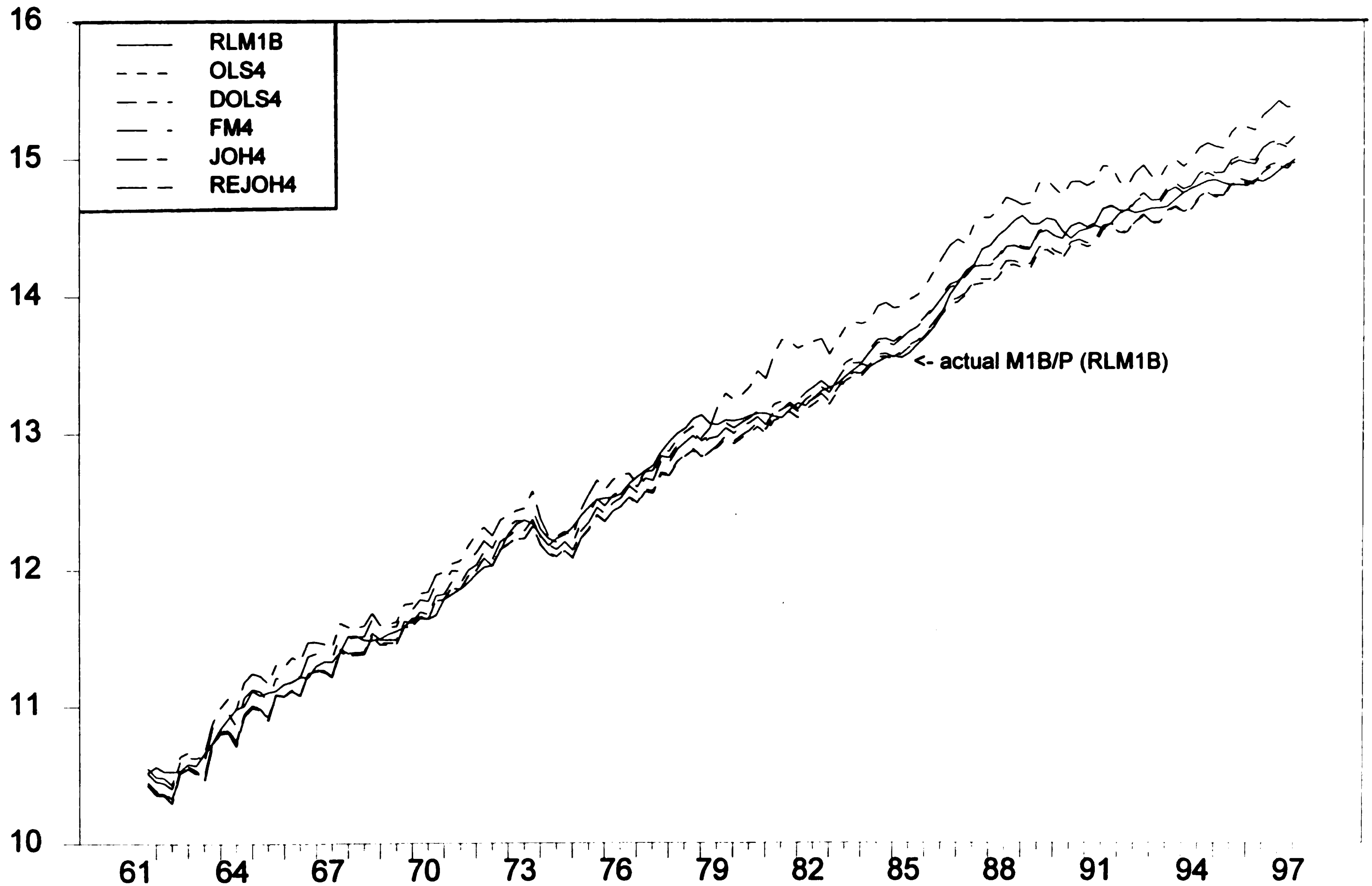

Figure 15: Actual and estimated M1B/P (Four-variable system)

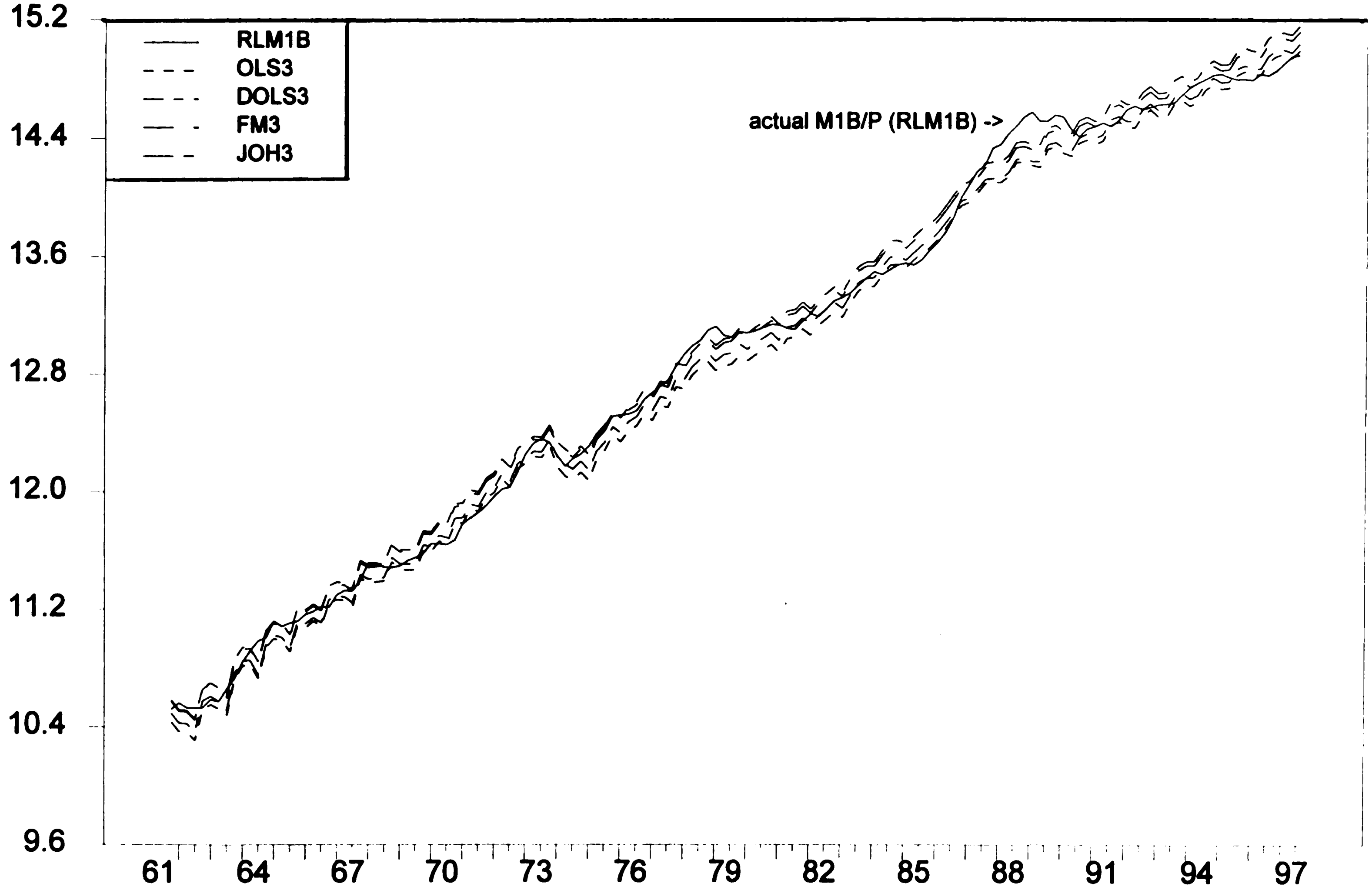

Figure 16: Actual and estimated M1B/P (Three-variable system)

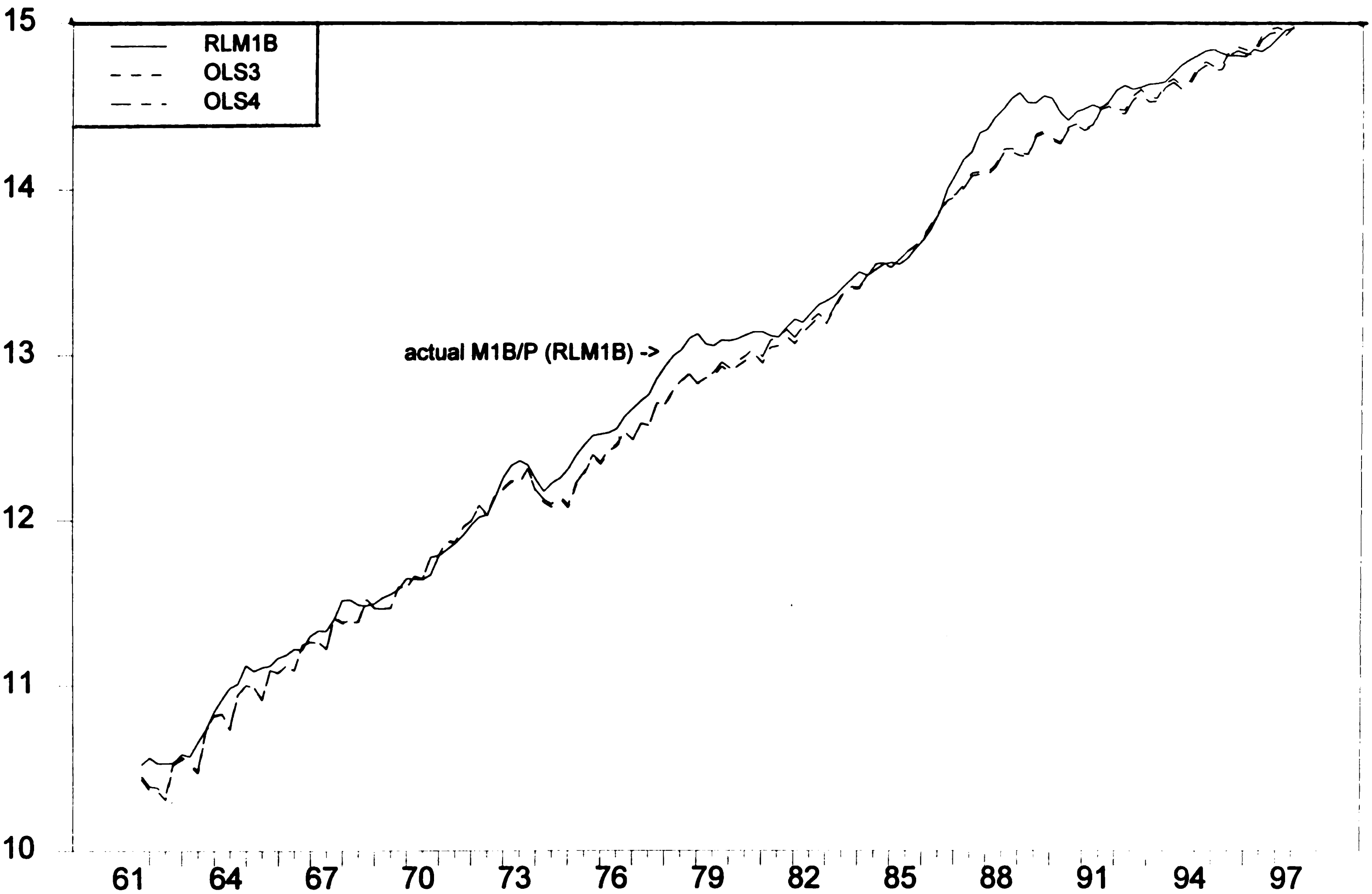

Figure 17: Actual and estimated M1B/P: OLS estimation

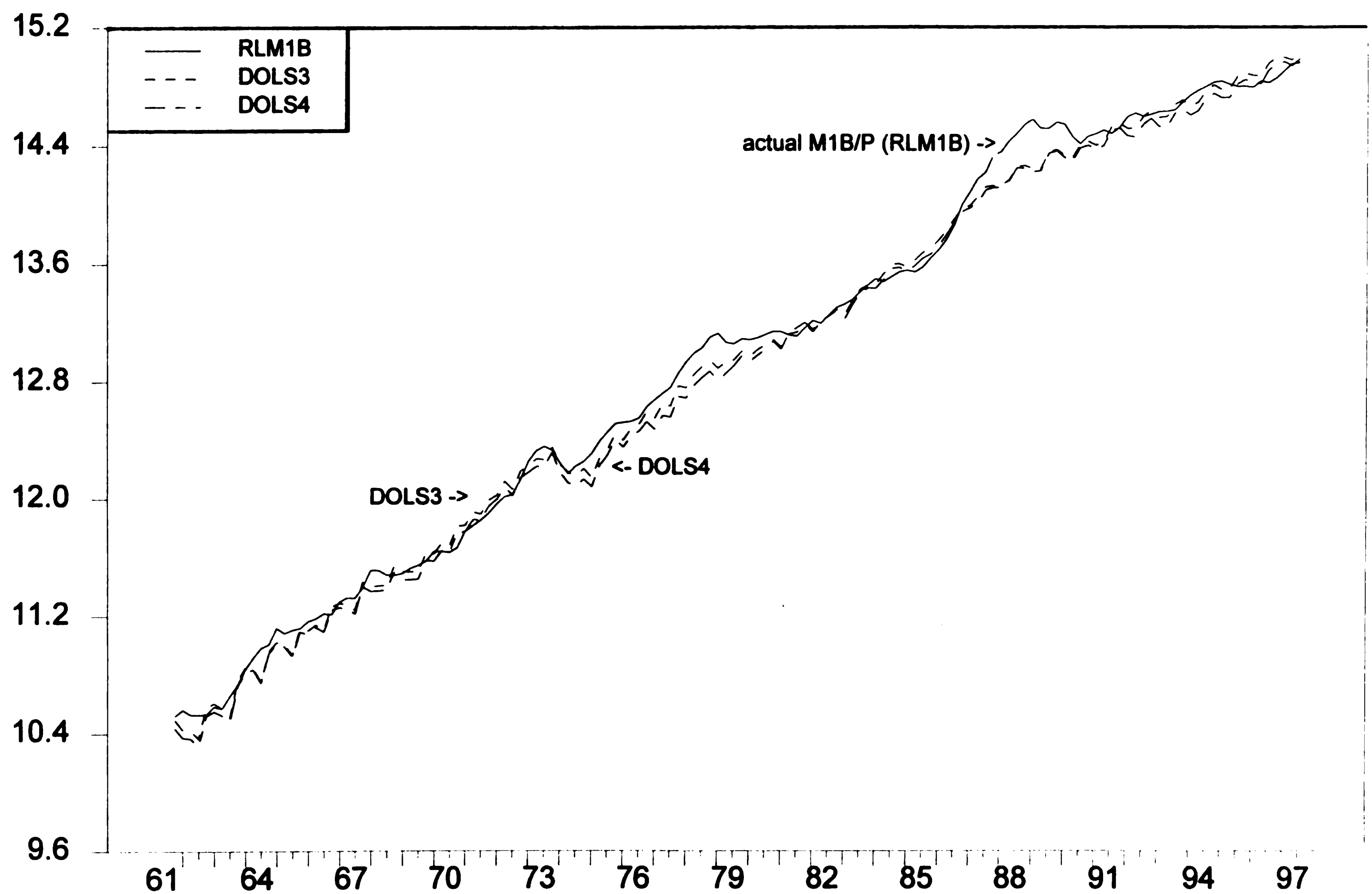

Figure 18: Actual and estimated M1B/P: DOLS estimation

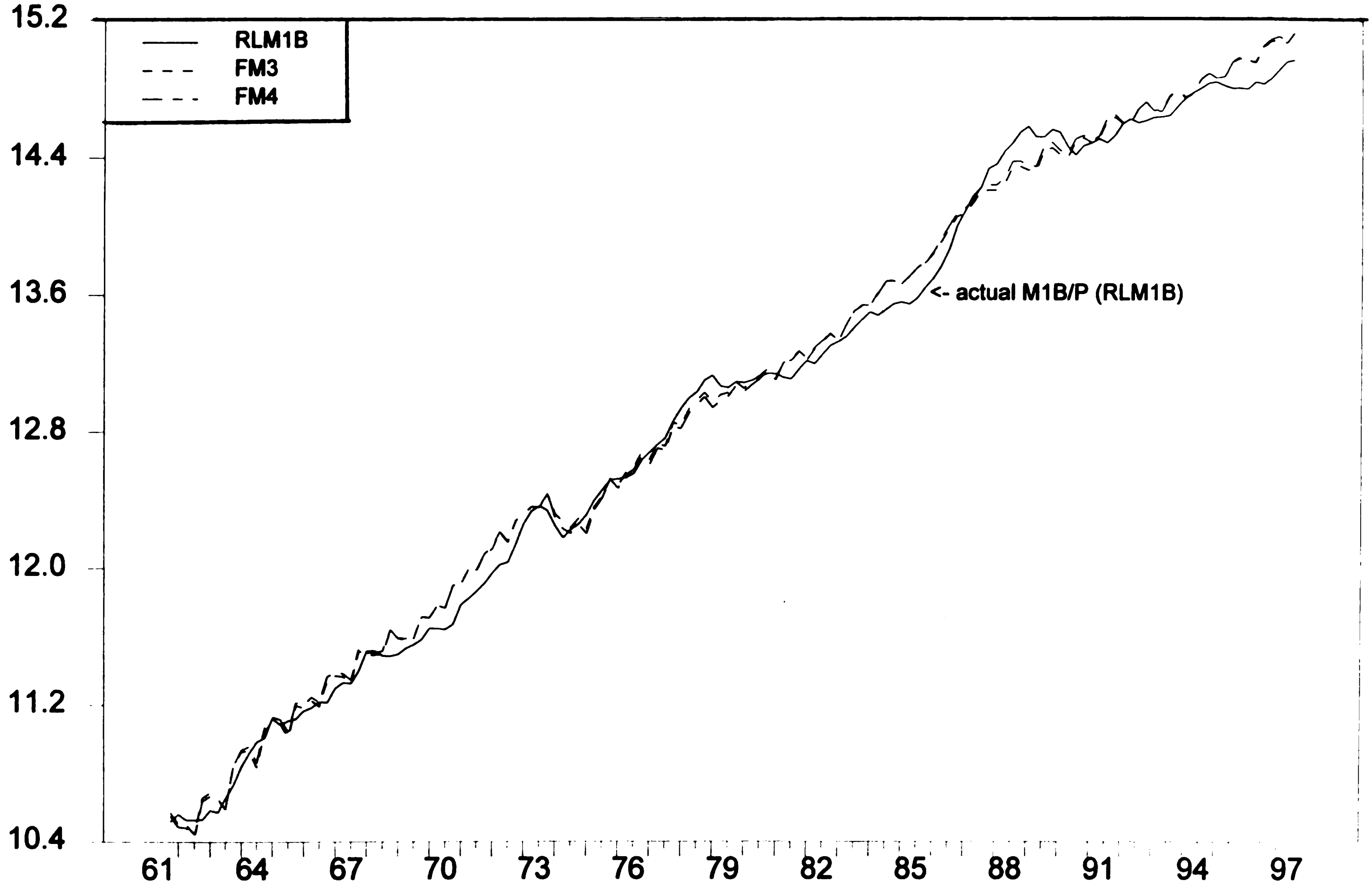

Figure 19: Actual and estimated M1B/P: FM estimation

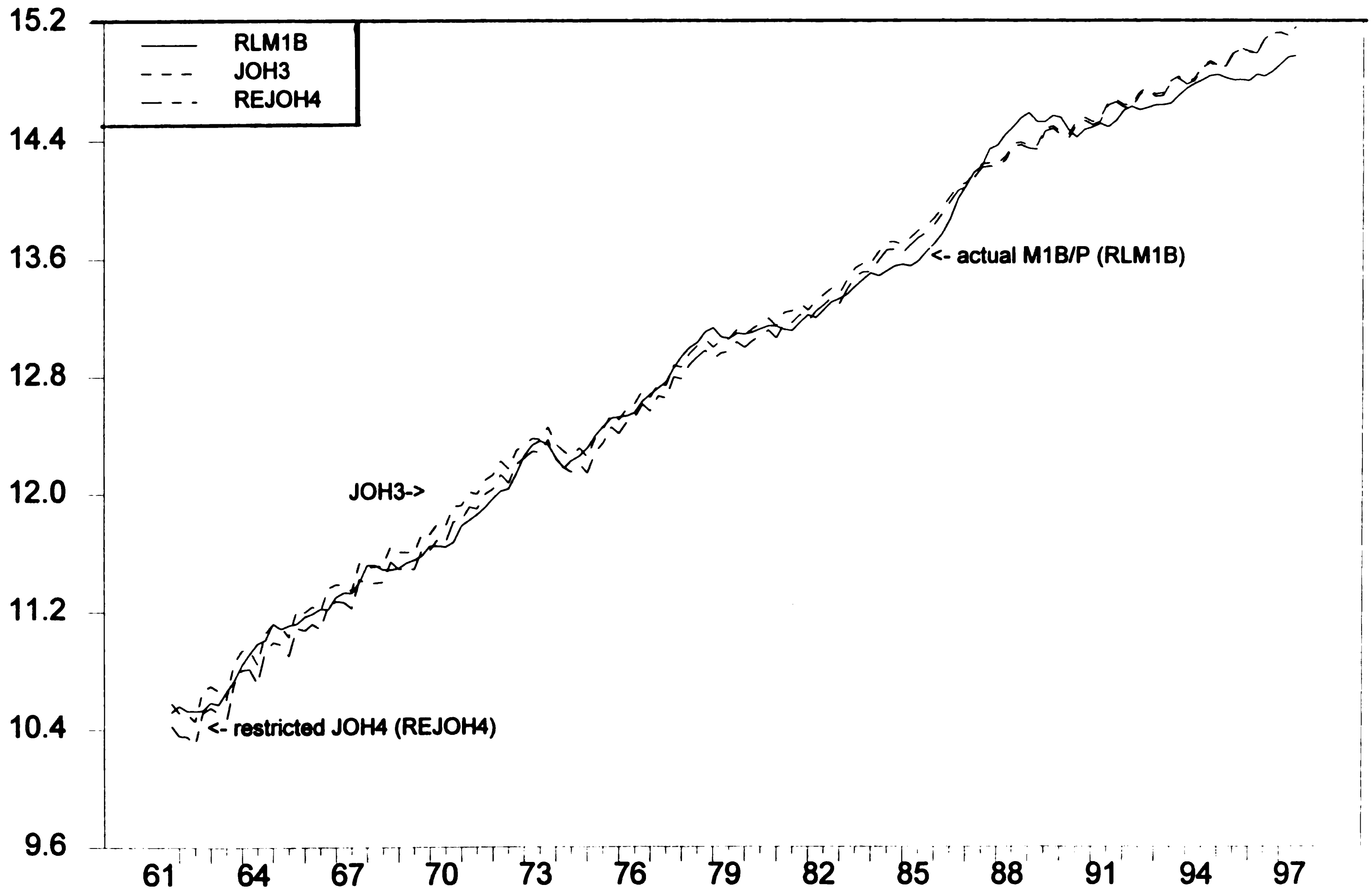

Figure 20: Actual and estimated M1B/P: JOH estimation

2 Tests of Linear Restrictions on Cointegrating Vectors (β) and the Adjustment Coefficients (α)

Because we are not only interested in inference about the cointegrating relations but also in their adjustment coefficients, the Johansen's likelihood ratio tests provide us tests for linear structural hypothesis on α and β. Note that the Wald statistics for linear restrictions on cointegrating vectors constructed from DOLS and FM also have asymptotic χ^2 distribution.

The likelihood ratio test of the linear restrictions about β and α, $H_4 : \beta = H\varphi$ *and* $\alpha = A\psi$ are constructed from solving both the unrestricted eigenvalue problem and restricted eigenvalue problem. (for detail see Johansen and Juselius 1990, Johansen 1991). Since the asymptotic distribution of the maximum likelihood estimator is mixed Gaussian, the hypothesis about cointegrating relations is to use the χ^2 distribution.

(1) Tests of linear restrictions on β

First, we test the unitary income elasticity hypothesis for the VAR(4) model

$$H_3^1 : \beta_m = -\beta_y$$

The likelihood ratio (LR) for H_3^1 against H_2 is computed as LR = 0.85, which is compared $\chi^2_{0.95}(1)$ = 3.84 so the unit income elasticity hypothesis cannot be rejected at the 5% level. The inability of rejection is also consistent with t test constructed from Johansen estimates, however it is inconsistent with t-ratios constructed by DOLS and PHFM. As pointed out in Hoffman, Rasche and Tieslau (1995), the difference in inference between the Johansen and the single equation estimation (FM here) follows from the ex ante and ex post normalization that distinguishes the two estimators.

Next the hypothesis that the coefficients for the own rate of M1B and the umm rate are equal with opposite sign

$$H_3^2: \ \beta_{i^b} = -\beta_{umm}$$

is tested. The likelihood ratio for H_3^2 is computed as LR = 1.37. Thus we fail to reject H_3^2 at the 5% level, which is consistent with the Wald test from FM (Table 12). Since we fail to reject both H_3^1 and H_3^2, we go on to test the joint hypothesis of

$$H_3^3: \ \beta_m = -\beta_y \text{ and } \beta_{i^b} = -\beta_{umm} \text{ against } H_2$$

The likelihood ratio is 7.00, which is compared $\chi^2_{0.95}(1)$ = 3.84. Interestingly, H_3^3 is rejected at the 5% level even though both $\beta_m = -\beta_y$ and $\beta_{ib} = -\beta_{umm}$ are rejected as shown above.

Moving to VAR(3) model, we compute the likelihood ratio for $H_3 : \beta_y = -1.50$ versus H_2 as LR = 1.60. Therefore, we fail to reject the null at the 5% level.

(2) Weak exogeneity tests

We examine whether some variables in the system are weakly exogeneous for the long-run parameters α and β, i.e. for some i, $\alpha_i = 0$, where i is an index of variables in the system. If Δx_{II} is weakly exogenous for α and β in the sense that the conditional distribution of Δx_I given Δx_{II} as well as the lagged values of x, contains the parameters α and β, whereas all distribution of Δx_{II} given the lagged x, does not contain the parameters α and β. Furthermore, the parameters in the conditional and marginal distribution are variation free, (Johansen and Juselius 1990, 1992, Johansen 1991a). Urbain (1992) demonstrates that the rejection of weak exogeneity status of the various variables sufficiently invalidates inference conducted in single error correction model (for both long-run and short-run parameters)[28]. In other words, the single equation ECM has parameters which are themselves a function of the parameters of the marginal process. Under such circumstance, one needs to conduct the analysis based on both marginal and conditional model, i.e. the full model.

In VAR(4) model, since the hypothesis $\beta_{i^b} = -\beta_{umm}$ cannot be rejected by the data, we test hypothesis about α in the restricted model under $\beta_{i^b} = -\beta_{umm}$, that is, H_4^i : α_i = 0 and $\beta_{i^b} = -\beta_{umm}$ versus H_3^2: $\beta_{i^b} = -\beta_{umm}$. For VAR(3), we test the hypothesis of H: $\beta_y = -1.50$ and $\alpha_i = 0$ against $\beta_y = -1.50$.

Table 26-27 reports the likelihood ratio (LR), which is asymptotically distributed as a $\chi^2(1)$ distribution under the null hypothesis (for our case r = 1). We can see the test statistic for each hypothesis that the cointegrating relations are not present in the equation determining each variable in the system is rejected at the 5% level for both VAR(4) and VAR(3) model. It indicates none of variables in the system can be treated as weakly exogenous in estimation of the long-run parameters α and β of the demand for real

[28] See proposition 1 and 2 in Urbain (1992) for the case r=1.

M1B in Taiwan. There is a loss efficiency if we attempt to consider partially specified systems where only some of variables are treated as endogenous and model these conditional on the remaining variables. We note that imposing invalid weak exogeneity hypothesis about some variable, for example, causes the estimates for the interest variables to posses wrong signs (the 3^{rd} and 5^{th} column in Table 26, the 3^{rd} and 4^{th} column in Table 27), and the standard deviations of estimates become very large. Thus, when estimating the long-run parameters of the demand for real M1B in Taiwan, we should treat all of the variables (real M1B, y, i^{b}, umm) as endogenous.

3 Estimation of the Long-Run Impact Matrix Π

Finally, we conclude our cointegration analysis of the Taiwanese demand for money by a comparison of the estimated long-run coefficients matrix $\Pi = \alpha\beta'$ and the restricted $\Pi = \alpha\beta$ under the constraint $\beta_{i^{b}} = -\beta_{umm}$ for VAR(4), and under $\beta_{y} = -1.50$ for VAR(3). The estimates for VAR(4) and VAR(3) are reported in Table 28 and 29, respectively.

4 Summary

Using the system method, we establish that there is one cointegrating vector among real M1B, real GNP, the interest rate spread (or the umm rate). The income elasticity is around 1.50 and significantly different from zero and unity. The semi-interest elasticity is around -0.017 and significantly different from zero under $\beta_{i^b} = -\beta_{umm}$. Using the sample mean value of the umm rate (24.12%) the interest elasticity is calculated as -0.41. These results are robust to the choice of estimators. Furthermore, all the variables should be treated as endogenous in estimating the long-run Taiwanese demand for real M1B. The rejection of weak exogeneity of the variables in the system suggests that it is invalid to conduct inference from the single equation conditional ECM only.

Table 26
Weak exogeneity tests: VAR(4) model

H^4	$\beta_{i^b} = -\beta_{umm}$	$\beta_{i^b} = -\beta_{umm}$ $\alpha_m = 0$	$\beta_{i^b} = -\beta_{umm}$ $\alpha_y = 0$	$\beta_{i^b} = -\beta_{umm}$ $\alpha_{i^b} = 0$	$\beta_{i^b} = -\beta_{umm}$ $\alpha_{umm} = 0$
LR	1.37	5.87**	7.36**	8.97**	3.99**
$\hat{\lambda}_1$	0.11	0.07	0.06	0.06	0.08
$\hat{\beta}_m$	1.00	1.00	1.00	1.00	1.00
	(0.14)	(0.25)	(0.17)	(0.25)	(0.23)
$\hat{\beta}_y$	-1.61	-1.70	-1.49	-1.28	-1.62
	(0.20)	(0.37)	(0.33)	(0.39)	(0.33)
$\hat{\beta}_{i^b}$	-0.02	0.010	-0.008	0.005	-0.04
	(0.01)	(0.002)	(0.01)	(0.02)	(0.06)
$\hat{\beta}_{umm}$	0.02	-0.010	0.008	-0.005	0.04
	(0.01)	(0.002)	(0.01)	(0.02)	(0.06)
$\hat{\alpha}_m$	-0.03	0.00	-0.07	-0.03	-0.04
$\hat{\alpha}_y$	0.04	0.03	0.00	-0.01	0.03
$\hat{\alpha}_{i^b}$	0.07	-0.003	0.04	0.00	0.11
$\hat{\alpha}_{umm}$	-0.72	-0.20	0.83	0.39	0.00

Note: $\hat{\lambda}_1$ is the largest eigenvalue obtained in the restricted eigenvalue problem.

Table 27
Weak exogeneity test: VAR(3) model

	β_y = -1.50	β_y = -1.50 α_m = 0	β_y = -1.50 α_y = 0	β_y = -1.50 α_{umm-i^h} = 0
LR	1.60	12.31**	12.31**	12.31**
$\hat{\lambda}_1$	0.08	0.00	2.0e006	4.0e-006
$\hat{\beta}_m$	1.00	1.00	1.00	1.00
	(0.28)	(154.18)	(59.37)	(41.90)
$\hat{\beta}_y$	-1.50	-1.50	-1.50	-1.50
	(0.28)	(154.18)	(59.37)	(41.90)
$\hat{\beta}_{umm-i^h}$	0.016	-0.02	-0.03	0.018
	(0.004)	(0.96)	(6.80)	(7.73)
$\hat{\alpha}_m$	-0.11	0.00	-2.7e-004	-7.0e-004
$\hat{\alpha}_y$	-0.01	2.5e-005	0.00	-9.1e-005
$\hat{\alpha}_{umm-i^h}$	0.75	2.4e-003	4.7e-003	0.00

Note: $\hat{\lambda}_1$ is the largest eigenvalue obtained in the restricted eigenvalue problem.

Table 28

VAR(4): Estimated long-run coefficients matrix: $\hat{\Pi} = \hat{\alpha}\hat{\beta}'$

H_2 : $\Pi = \alpha\beta'$			
RM1B	y	i^b	umm
-2.8e-004	4.0e004	9.6e-005	-1.4e-005
0.0406	-0.0579	-0.014	0.0020
0.2079	-0.2962	-0.0710	0.0103
-0.1798	0.2561	0.0614	-0.0890
H_3^2 : $\Pi = \alpha\beta'$ and $\beta_{i^b} = -\beta_{umm}$			
-0.0321	0.0518	-7.0e-004	
0.0446	-0.0719	9.7e-004	
0.0661	-0.1067	0.0014	
-0.7216	1.1643	-0.0157	

Table 29

VAR(3): Estimated long-run coefficients matrix: $\hat{\Pi} = \hat{\alpha}\hat{\beta}'$

H_2 : $\Pi = \alpha\beta'$		
m	y	umm-i^b
-0.1204	0.1847	-0.0020
0.0131	-0.0203	2.3e-004
0.4070	-0.6334	0.0070
H_3^2 : $\Pi = \alpha\beta'$ and $\beta_y = -1.50$		
-1.3482	-0.0211	
-0.1502	-0.0023	
8.9325	0.1400	

CHAPTER 5

CONCLUSIONS

This dissertation has examined the stability of the long-run (equilibrium) demand functions for four alternative new definitions of money in Taiwan since 1961:4. Unlike many money demand studies in the Taiwanese literature which use a single bank deposits rate as a proxy for the opportunity cost of holding money balances, we employ the interest rate spread between the own rate of money and the unorganized money market rate to incorporate the characteristics of the financial structure in Taiwan.

After considerations have been undertaken to account for data that exhibit both stochastic and deterministic trends and seasonality as well, using various econometric techniques we find that there is a strong evidence for the existence of stationary linear combinations of real balances, real GNP, and the short-term interest rates in each money demand equation spanning 1961:4-1997:3 in Taiwan. However, the stability tests detect apparent parameter nonconstancy in broadly defined money equations (real M1BP, real M2). On the one hand, parameter instability may be explained by regime shifts over the sample period, including the oil crisis in 1973, 1979, the shift from fixed to floating exchange rate regime in 1978, and the financial deregulation since 1980. On the other hand, the absence of constant parameter may be explained by a misspecification in econometric model thereof while the "true" long-run relationship remains constant in spite of these institutional changes. Although results from single equation methods (Phillips and Ouliaris 1990, Gregory and Hansen 1996, and Hansen 1992b) suggest that real M1A, real GNP, and the short-term interest rate (umm or the spread) are cointegrated with stable parameters while a deterministic trend remains in the cointegration relations, the system method (Johansen and Juselius 1990, Johansen 1991) suggests otherwise. Based on economic theories and an implausible huge income elasticity as well, we doubt this long-run relationship. Results reveal that the most stable long-run relationship occurs in the demand for real M1B data (real M1B, real GNP, the short-term interest rate (the umm rate or the spread)), indicating that the narrow M1B aggregate is the preferable measure with which to consider the long-run economic impacts of changes in monetary policy in Taiwan. Moreover, it also provides a useful anchor for study of error correction or short-run dynamic and can also be used in common trend exercises designed to reveal the underlying source of nonstationarity of an economic system. The contrasting results on the issue of stability from the demand for two seemingly homogenous monetary aggregates (real M1B, and real M1BP) are explained by the difference in liquidity characteristics in these two measures of money.

Unlike insignificant role for an interest rate as the opportunity cost of holding money in a developing country setting, we find that the demand for real M1B is interest (the umm rate) elastic. The unorganized money market rate has shown to be an empirically relevant variable in the study of M1B demand in Taiwan, which may have important policy message as documented in Wijnbergen (1982, 1983a,b, and 1985). The estimated equilibrium income elasticity is significantly different from unity and is on the order of 1.50. The estimated equilibrium interest elasticity is on the order of -0.41. The transactions money demand model derived in CHAPTER 1, which possesses steady-state properties of unitary income elasticity and a negative effect of the short-run interest rate and appears consistent with several empirical findings in the money demand literature using the data from developed countries, is inadequate to explain the empirical findings in our study for the long-run demand for real M1B in Taiwan.

APPENDIX 1

APPENDIX 1

THE REDEFINED MONETARY AGGREGATES IN TAIWAN

The central bank in Taiwan (the Central Bank of China, CBC) redefined the monetary aggregates in February, 1997. The definitions of money supply are revised to include the total deposits in the Postal Savings System. For the M2 definition, repurchase agreements and non-residents N.T. (New Taiwan dollar) deposits are also added, while the foreign exchange trust funds and the bank debentures, savings bonds and Treasury bills held by the public are excluded from the original definition. The new aggregates are presented in Table A1

Table A1
New measures of money
Millions of New Taiwan (NT) dollars, seasonally unadjusted, August 1997

M1A	1,721,421
Currency in circulation[1]	493,057
Checking accounts[2]	325,415
Passbook deposits[2]	902,949
M1B	3,728,568
M1A	1,721,421
Passbook savings deposits[2]	2,007,147
M2	14,520,631
M1B	3,728,568
Time and time savings deposits[3]	7,855,219
Postal savings deposits[4]	2,299,346
Foreign currency deposits[5]	399,206
Repurchase agreements issued by monetary institution and Postal Savings System[6]	182,952
Non-resident N.T.(New Taiwan dollar) deposits[7]	55,340

Notes:
1 This component equals currency issued by CBC - cash in vaults held by monetary institutions and the Postal Savings System.

2 These deposits are also referred as monetary deposits, held by enterprises and individuals in monetary institutions. According to Banking Law, savings deposits can only be held by individuals and non-profit organizations.

3 Time deposits include general time deposits and negotiable certificates of deposit (NCDs). Time savings deposits also include deposits replaced by the Postal Savings System.

4 This component includes giro accounts, passbook savings deposits and time savings deposits of the Postal Savings System.

5 It includes foreign currency deposits, and foreign currency certificates of deposits of enterprises and individuals in monetary institutions.

Notes (cont'd)

6 It represents repurchase agreements sold to enterprises and individuals by monetary institutions and the Postal Savings System. Prior to January 1994, the data are not available.

7 It includes demand and time deposits held by foreign non-financial institutions. Prior to January 1994, the data are not available.

Source:
Financial Statistics Monthly, November 1997, the Central Bank of China (the Republic of China, Taiwan).

APPENDIX 2

APPENDIX 2

SOURCES OF FUNDS BORROWED BY PRIVATE ENTERPRISES IN TAIWAN

Table A2 displays the sources of funds borrowed by private enterprises in Taiwan during 1976-1988. From Table A2, we can see that the private enterprises actively borrow from the unorganized money market (U.M.M.) (around one-third of borrowing comes from U.M.M. during 1976-1988), despite the existence of the organized money market since 1976. It appears that there is no sign showing that the importance of the U.M.M. has declined over time regardless substantial financial reforms have been launched to modernize the financial system in Taiwan.

Table A2
Source of funds borrowed by private enterprises in Taiwan

End of year	Financial Institution	Money Market	U.M.M.	%
1976	67.30	0.57	32.13	100.00
1977	63.31	2.89	33.80	100.00
1978	59.98	4.02	36.00	100.00
1979	57.15	8.13	34.72	100.00
1980	54.95	8.91	36.14	100.00
1981	52.80	11.71	35.49	100.00
1982	53.42	13.38	33.20	100.00
1983	54.88	14.31	30.81	100.00
1984	54.71	15.11	30.18	100.00
1985	51.69	13.26	35.05	100.00
1986	50.58	8.41	40.01	100.00
1987	56.22	6.65	37.13	100.00
1988	63.67	5.76	30.57	100.00
Average	56.97	8.97	34.33	100.00

Source: The Efficiency of Allocation of Funds in the Taiwanese Financial System (Table 6, Shea 1991).

APPENDIX 3

APPENDIX 3

TESTING FOR PRICE HOMOGENEITY OF THE MONEY DEMAND FUNCTIONS IN TAIWAN (1961:4-1997:3)

Since economic theories predict that the demand for money is a demand for "real" balances (money holdings measured in constant purchasing power terms), a price level elasticity of the demand for nominal balances is frequently constrained to one. In this appendix, we use the fully modified estimator (FM) of Phillips and Hansen (1990) (see Section III, CHAPTER 4) to estimate θ in the nominal money (M1A, M1B, M1BP, and M2) cointegrating relation

$$m_t = \mu + \alpha_y y_t + \alpha_i i_t + \alpha_{umm} umm_t + \theta P_t + v_t$$

and to test whether $\theta = 1$ using t-ratio constructed from FM estimates, where y, i, umm are the same as defined before; p is GNP deflator (1991 = 100). In Table A3, we report the (FM) estimates, their standard deviations, and Hansen (1992b) parameter stability tests.

As seen in Table A3, the coefficients on price in all equations are not significantly different from one at the 5%

level. Note that the coefficient in M1A equation is not significantly different from zero at either 5% or 10% level.

Table A3
Testing for price homogeneity: $\theta = 1$

$$m_t = \mu + \alpha_y \mathbf{y}_t + \alpha_i i_t + \alpha_{umm} \mathbf{umm}_t + \theta \mathbf{P}_t + \mathbf{v}_t$$

Regression	M1A	M1B	M1BP	M2
L_c	4.69**	1.81**	4.45**	23.18**
MeanF	33.26**	3.11	5.95	60.51**
SupF	280.89**	10.17	27.61**	442.67**
	[67:1]	[74:2]	[67:2]	[67:1]

FM estimates

Regressors	M1A	M1B	M1BP	M2
constant	3.29	-2.30	- 0.64	-10.10
	(0.38)	(0.31)	(0.25)	(0.15)
y	0.45	0.95	0.71	- 2.05
	(0.24)	(0.18)	(0.15)	(0.10)
i	- 0.36	-0.10	- 0.03	- 0.01
	(0.23)	(0.10)	(0.04)	(0.02)
umm	- 0.08	-0.05	- 0.05	- 0.02
	(0.02)	(0.01)	(0.01)	(0.01)
p	2.40	2.05	2.48	1.54
	(1.95)	(1.40)	(1.15)	(0.80)

Note: ***m*** is the log of nominal money supply. p is the log of the GNP deflator.
** indicates significance at the 5% level.
See Table 15-16 for critical values.

APPENDIX 4

APPENDIX 4

TESTING FOR NO COINTEGRATION BETWEEN THE UNORGANIZED MONEY MARKET RATE AND THE OWN RATE OF MONEY IN TAIWAN (1961:4-1997:3)

In this appendix, we use standard residual based cointegration tests Z_t and Z_α (see Section II, CHAPTER 4) to test the hypothesis of no cointegration occurred between the unorganized money market rate (umm) and the own rate on each measure of money (i^a, i^b, i^{bp}, i^{m2}). Test results and the OLS estimates (and standard errors) are reported in Table A4.

From Table A4, we find that the hypothesis of no cointegration is not rejected for each pair of (umm, i) at the 5% level. The results are invariant to the normalization selected.

Table A4
Testing for no cointegration between the unorganized money market rate (umm) and the own rate (i)

	$umm_t = constant + \alpha i^a_t + \varepsilon_t$				
OLS estimates	constant	i^a	$\bar{R}^2$	SEE	DW
	22.78	2.16	0.009	3.13	0.06
	(0.91)	(1.40)			
			Z_t	Z_α	$\hat{M}$
			-1.56	-4.40	1.10
	$i^a_t = constant + \alpha umm_t + \varepsilon_t$				
OLS estimates	constant	umm	$\bar{R}^2$	SEE	DW
	0.44	0.008	0.009	0.19	0.08
	(0.12)	(0.005)			
			Z_t	Z_α	$\hat{M}$
			-1.77	-6.03	0.33
	$umm_t = constant + \alpha i^b_t + \varepsilon_t$				
OLS estimates	constant	i^b	$\bar{R}^2$	SEE	DW
	20.30	2.51	0.32	2.58	0.08
	(0.51)	(0.30)			
			Z_t	Z_α	$\hat{M}$
			-1.98	-6.58	1.16
	$i^b_t = constant + \alpha umm_t + \varepsilon_t$				
OLS estimates	constant	umm	$\bar{R}^2$	SEE	DW
	-1.65	0.13	0.33	0.59	0.04
	(0.38)	(0.02)			
			Z_t	Z_α	$\hat{M}$
			-1.78	-4.42	1.18
			Critical Values		
				Z_t	Z_α
			5%	-3.37	-20.5
			10%	-3.07	-17.0

An * indicates significance at the 10% level.
Critical values are from Table 1 and 2 (n=1, Π=0) in Hansen (1992a).

Table A4 (cont'd)

	$umm_t = constant + \alpha i^{hp}_t + \varepsilon_t$				
OLS estimates	constant	i^{hp}	$\bar{R}^2$	SEE	DW
	19.84 (0.51)	2.05 (0.16)	0.53	2.15	0.11
			Z_t	Z_α	$\hat{M}$
			-2.53	-9.97	1.19
	$i^{hp}_t = constant + \alpha umm_t + \varepsilon_t$				
OLS estimates	constant	umm	$\bar{R}^2$	SEE	DW
	-4.23 (0.50)	0.26 (0.02)	0.53	0.77	0.07
			Z_t	Z_α	$\hat{M}$
			-2.55	-8.33	1.30
	$umm_t = constant + \alpha i^{m_2}_t + \varepsilon_t$				
OLS estimates	constant	i^{m_2}	$\bar{R}^2$	SEE	DW
	13.04 (0.70)	2.37 (0.15)	0.65	1.87	0.20
			Z_t	Z_α	$\hat{M}$
			-2.77	-14.51	0.80
	$i^{m_2}_t = constant + \alpha umm_t + \varepsilon_t$				
OLS estimates	constant	umm	$\bar{R}^2$	SEE	DW
			0.64	0.64	0.24
	-1.93 (0.41)	0.27 (0.02)	Z_t	Z_α	$\hat{M}$
			-3.11*	-17.69*	0.84
			Critical Values		
				Z_t	Z_α
			5%	-3.37	-20.5
			10%	-3.07	-17.0

An * indicates significance at the 10% level.
Critical values are from Table 1 and 2 (n=1, Π=0) in Hansen (1992a).

BIBLIOGRAPHY

BIBLIOGRAPHY

English Literature

Andrews, D.W.K., 1990, "Tests for Parameter Instability and Structural Change with Unknown Change Point", Discussion paper 943, Yale University, Cowles Foundation.

Andrews, D.W.K., 1991, "Heteroskedasticity and Autocorrealtion Consistent Covariance Matrix Estimation", Econometrica 59, 817-858.

Andrews, D.W.K., and J.C. Monahan, 1992, "An Improved Heteroskedasticity and Autocorrelation Consistent Covariance Matrix Estimator", Econometrica 60, 953-966.

Banerjee, A., R.L. Lumsdaine, and J.H. Stock, 1992, "Recursive and Sequential Tests of the Unit-Root and Trend-Break Hypotheses: Theory and International Evidence, Journal of Business and Economic Statistics 10, 271-287.

Barsky, R.B. and J.A. Miron, 1989, "The Seasonal Cycle and the Business Cycle", Journal of Political Economy 97, 503-534.

Baumol, W.J., 1952, "The Transactions Demand for Cash - An Inventory Theoretical Approach", Quarterly Journal of Economics 66, 545-56.

Beaulieu, J.J., and J.A. Miron 1993, "Seasonal Unit-Roots in Aggregate U.S. Data", Journal of Econometrics 55, 305-328.

Beveridge, S., and C.R. Nelson, 1981, "A New Approach to Decomposition of Economic Time Series into Permanent and Transitory Components with Particular Attention to Measurement of the "business cycle"", Journal of Monetary Economics 7, 151-74.

Box, G.E.P. and G.M. Jenkins, 1970, Time Series Analysis, Forecasting and Control (Holden-Day, San Francisco, CA).

Canova, F. and B.E. Hansen, 1993, "Are Seasonal Patterns Constant over Time? A Test for Seasonal Stability", revised working paper (Brown University, Providence, R.I; European University Institute, Domeninco di Fiesole; University of Rochester, Rochester, NY).

Campbell, J.Y., N.G. Mankiw, 1987, "Are Output Fluctuations Transitory", The Quarterly Journal of Economics 102, 857-80.

Chan, N.H. and C.Z. Wei, 1988, "Limiting Distributions of Least Squares Estimates of Unstable Autoregressive Processes", Annals of Statistics 16, 367-401.

Chow, G., 1966, "On the Short-Run and Long-Run Demand for Money", Journal of Political Economy 74, 111-31.

Cochrane, J.H., 1988, "How Big is the Random Walk Component in GNP ?", Journal of Political Economy 96, 893-920.

DeJong, D.N. , J.C. Nankervis, N.E. Savin, and C.H. Whiteman, 1989, "Integration versus Trend Stationarity in Macroeconomic Time Series", Working paper no. 89-99 (Department of Economics, University of Iowa, Iowa City, IA).

Dickey, D.A., and W.A. Fuller, 1979, "Distribution of the Estimators for Autoregressive Time Series with a Unit Root, Journal of the American Statistical Association 74, 427-31.

Dickey, D.A., and W.A. Fuller, 1981, "Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root", Econometrica 49, 1057-72.

Dickey, D., D.P. Hasza, and W.A. Fuller, 1984, "Testing for Unit Roots in Seasonal Time Series", Journal of the American Statistical Association 79, 355-367.

Diamond, P.A., A Search-Equilibrium Approach to the Micro Foundations of Macroeconomics, 1984 (Cambridge, M.A.: MIT Press).

Engle, R.F., and C.W.J, Granger, 1987, "Co-Integration and Error Correction: Representation, Estimation, and Testing", Econometrica 55, 251-76. Reprinted in R.F. Engle and C.W.J. Granger (eds), Long-Run Economic Relationships: Readings in Cointegration, Oxford University Press: New York, 1991.

Engle, R.F. and C.W.J. Granger, 1991, Introduction, Long-Run Economic Relationships: Readings in Cointegration, Oxford University Press: New York, 1991.

Fisher, Irving, 1911, The Purchasing Power of Money, 2nd, revised ed. 1922; New York: Kelly and Millman, 1963, (first published in 1911).

Friedman, M., 1956, "The Quantity Theory of Money - A Restatement", in Studies in The Quantity Theory of Money, University of Chicago Press.

Friedman, M., 1957, A Theory of the Consumption Function, Princeton University Press, Chicago.

Friedman, M., 1968, "The Role of Monetary Policy", American Economic Review (March), 1-17. Reprinted in P.G. Korliras and R.S. Thorn (eds.), Modern Macroeconomics: Major Contributions to Contemporary Thought, New York: Harper and Row, 1979, pp.91-102.

Friedman, M., 1974, "A Theoretical Framework for Monetary Analysis", in Gordon, R., ed. Milton Friedman's Monetary Framework, University of Chicago Press, Chicago.

Fuller, W.A., 1976, Introduction of Statistical Time Series Wiley, New York, N.Y.

Ghysels, E., 1990, "Unit-Root Tests and the Statistical Pitfalls of Seasonal Adjustment: The Case of U.S. Postwar Real Gross National Product", Journal of Business and Economic Statistics 8 (No 2), 145-348.

Ghysels, E., H.S. Lee, J.N. Noh, 1994, "Testing for Unit Roots in Seasonal Time Series: Some Theoretical Extensions and A Monte Carlo Investigation", Journal of Econometrics 62, 415-442.

Goldfeld, S., 1973, "The Demand for Money Revisited", Brookings Papers on Economic Activity 3, 577-646.

Goldfeld, S., 1976, "The Case of the Missing Money", Brookings Papers on Economic Activity 3, 683-730.

Gonzalo, J., 1994, "Five Alternative Methods of Estimating Long-Run Equilibrium Relationships", Journal of Econometrics 60, 203-233.

Granger, C.W.J., and P. Newbold, 1974, "Spurious Regressions in Econometrics", Journal of Econometrics 2, 111-20.

Gregory, A.W. and B.E. Hansen, 1996, "Residual-Based Tests for Cointegration in Models with Regime Shifts", Journal of Econometrics 70, 99-126.

Hamilton, J.D., 1994, Time Series Analysis, Princeton University Press, Princeton, New Jersey.

Hansen, B.E., 1992a, "Efficient Estimation and Testing of Cointegrating Vectors in the Presence of Deterministic Trends", Journal of Econometrics 53, 87-121.

Hansen, B.E., 1992b, "Tests for Parameter Instability in Regressions with I(1) Processes", Journal of Business and Economic Statistics 10(No.3), 321-335.

Hansen, H. and S. Johansen, 1993, "Recursive Estimation in Cointegrated VAR Models", unpublished manuscript, University of Copenhagen, Copenhagen.

Haug, A.A., 1992, "Critical Values for $\hat{Z}_{\alpha}$- Phillips-Ouliaris Test for Cointegration", Oxford Bulletin of Economics and Statistics 54, 473-480.

Hendry, D.F., and N.R. Ericsson, 1991, "Modeling the Demand for Narrow Money in the United Kingdom and the United States", European Economic Review 35, 833-848.

Hoffman, D.L., and R.H. Rasche, 1991, "Long-Run Income and Interest Elasticities of Money Demand in the United States", The Review of Economics and Statistics 78, 665-674.

Hoffman, D.F., C. Tahiri, 1994, "Money Demand in Morroco: Estimating Long-Run Elasticities for a Developing Country", Oxford Bulletin of Economics and Statistics 56, 35-24.

Hoffman, D.L., R.H. Rasche, and M.A. Tieslau, 1995, "The Stability of Long-Run Money Demand in Five Industrial Countries", Journal of Monetary Economics 35, 317-339.

Hylleberg, S., R.F. Engle, C.W.J. Granger, and B.S. Yoo, 1990, "Seasonal Integration and Cointegration", Journal of Econometrics 44, 215-238.

Hylleberg, S., 1995, "Tests for Seasonal Unit Roots: General to Specific or Specific to General?", Journal of Econometrics 69, 5-25.

Jarque, C.M. and A.K. Bera, 1980, "Efficient Tests for Normality, Homoscedasticity and Serial Independence of Regression Residuals", Economic Letters 6, 255-59.

Johansen, S., 1988a, "The Mathematical Structure of Error Correction Models", Contemporary Mathematics, American Mathematical Society 80, 359-86.

Johansen, S., 1988b, "Statistical Analysis of Cointegration Vectors", Journal of Economic Dynamics and Control 12, 231-54. Reprinted in R.F. Engle and C.W.J. Granger (eds.), Long-Run Economic Relationships: Readings in Cointegration, Oxford University Press: New York, 1991.

Johansen, S., 1992, "Cointegration in Partial Systems and the Efficiency of Single Equation Analysis", Journal of Econometrics 52, 389-402.

Johansen, S., 1991, "Estimation and Hypothesis Testing of Cointegration Vectors in Gaussian Vector Autoregressive Models", Econometrica 59, 1551-80.

Johansen, S., and K. Juselius, 1990, "Maximum Likelihood Estimation and Inference on Cointegration - with Applications to the Demand for Money", Oxford Bulletin of Economics and Statistics 52, 169-210.

Keynes, J.M., 1936, The General Theory of Employment, Interest, and Money, Prometheus Books, New York, 1997.

King, R.G., C.I. Plosser, J.H. Stock, and M.W. Watson, 1991, "Stochastic Trends and Economic Fluctuations", American Economic Review 81, 819-840.

Kydland, F.E., E.C. Prescott, 1982, "Time to Build and Aggregate Fluctuations", Econometrica 50(November), 1345-70.

Kwiatkowski, D, P.C.B. Phillips, P.Schmidt, and Y. Shin, 1992, "Testing the Null Hypothesis of Stationarity against the Alternative of a Unit Root: How Sure are We that Economic Time Series Have a Unit Root ?", Journal of Econometrics 54, 159-178.

Laidler, D., 1966, "The Rate of Interest and the Demand for Money - Some Empirical Evidence", Journal of Political Economy 74(December), 545-55.

Lee, C.N., 1994, "The Transactions Money Demand in Taiwan: Analysis of Stability and the Importance of the Innovations in the Common Stochastic Trend", Ph.D. dissertation, Department of Economics, Michigan State University.

Ljung, G., and G. Box, 1979, "On a Measure of Lack of Fit in Time Series Models", Biometrika 66, 265-270.

Lucas, R.E. Jr., 1976, "Econometric Policy Evaluations: A Critique, in the Phillips Curve and Labor Markets", Karl Brunner and Allan H. Meltzer eds. Amsterdam: North-Holland.

Lucas, R.E. Jr., 1988, "Money demand in the United States: A Quantitative Review", in K. Brunner and B.T. McCallum (eds.) Money Cycles and Exchange Rates : Essays in Honour of Allen Meltzer (Carnegie - Rochester Conference Series on Public Policy, 29), North-Holland.

Mankiw, N.G., 1991, "The Reincarnation of Keynesian Economics", NBER working paper no.3885.

Mackinnon, J.G., 1990, "Critical Values for Cointegration" in Engle, R.F. and Granger, C.W. (eds.) Long-Run Economic Relationships : Reading in Cointegration, Oxford, Oxford University Press.

McCallum, B.T., 1989, Monetary Economics: Theory and Policy, Macmillan Publishing Company, New York.

Meltzer, A.H., 1963, "The Demand for Money: the Evidence from the Time Series", Journal of Political Economy 71, 405-22.

Miller, M.H. and D. Orr, 1956, "A Model of the Demand for Money by Firms", Quarterly Journal of Economics, August, 413-35.

Nelson, C.R., C.I. Plosser, 1982, "Trends and Random Walks in Macroeconomic Time Series", Journal of Monetary Economics 10, 139-162.

Osterwald-Lenum, M., 1992, "A Note with Quantiles of the Asymptotic Distribution of the Maximum Likelihood Cointegration Rank Test Statistics", Oxford Bulletin of Economics and Statistics 54(No.3), 461-471.

Perron, P. and T.J. Vogelsang, 1992, "Nonstationarity and Level Shifts with an Application to Purchasing Power Parity", Journal of Business and Economic Statistics 10, (No.3), 301-320.

Phillips, P.C.B., 1986, "Understanding Spurious Regressions in Econometrics", Journal of Econometrics 33, 311-40.

Phillips, P.C.B., 1987, "Time Series Regression with A Unit Root, Econometrica 55, 277-301.

Phillips, P.C.B. and P. Perron, 1988, "Testing for a Unit Root in Time Series Regression", Biometrika 75, 335-346.

Phillips, P.C.B. and B.E. Hansen, 1990, "Statistical Inference in Instrumental Variables Regression with I(1) Processes", Review of Economic Studies 57, 99-125.

Phillips, P.C.B. and S. Ouliaris, 1990, "Asymptotic "Properties of Residual Based Tests for Cointegration", Econometrica 58:165-193.

Phillips, P.C.B., and M. Loretan, 1991, "Estimating Long-Run Economic Equilibria", Review of Economic Studies 58, 407-436.

Poole, William, 1988, "Monetary Policy Lessons of Recent Inflation and Disinflation", Journal of Economic Perspectives 2 (Summer), 73-100.

Rasche, R.H., 1990, "Equilibrium Income and Interest Easticities of the Demand for M1 in Japan", Bank Of Japan, Monetary and Economic Studies 8(No.2), September, 31-58.

Rasche, R.H., D.L. Hoffman, 1996, Aggregate Money Demand Functions: Empirical Applications in Cointegrated Systems, Kluwer Academic Publishers.

Saikkonen, P., "Asymptotically Efficient Estimation of Cointegrating Regressions", Econometric Theory 7, 1-21.

Sargent, T.J., N. Wallace, 1975, "Rational Expecations, the Optimal Monetary Instrument, and the Optimal Money Supply Rule", Journal of Political Economy, April, 241-254.

Saving, T.R., 1971, "Transactions Costs and the Demand for Money", American Economic Review, June, 407-420.

Semkow, B.W., 1992, Taiwan's Financial Markets and Institutions: The Legal and Financial Issues of Deregulation and Internationalization, Quorum Books.

Semkow, B.W., 1994, Taiwan's Capital Market Reform: The Financial and Legal Issues, Clarendon Press, Oxford.

Stein, J.L., 1976, Monetarism, Studies in Monetary Economics, Vol.1, North-Holland Publishing Company, Amsterdam, New York, Oxford.

Stock, J.H., 1987, "Asymptotic Properties of Least Squares Estimators of Cointegrating Vectors", Econometrica 55, 1035-1056.

Stock, J.H. and M.W. Watson, 1988, "Testing for Common Trends", Journal of the American Statistical Association, 83, 1097-1107.

Stock, J.H. and M.W. Watson, 1993, "A Simple Estimator of Cointegrating Vectors in Higher Order Integrated System", Econometrica 61, 783-820.

Tobin, J., 1956, "The Interest Elasticity of Transactions Demand for Cash", The Review of Economics and Statistics, XXXVIII August, 241-247.

Tobin, J., 1958, "Liquidity Preference as Behavior Toward Risk", Review of Economic Studies, February, 65-86

Urbain, J.P., 1992, "On Weak Exogeneity in Error Correction Models", Oxford Bulletin of Economics and Statistics 54, 187-207.

Wijnbergen, S. van, 1982, "Stagflation Effects of Monetary Stabilization Policies - A Quantitative Analysis of South Korea", Journal of Development Economics 10, 133-169.

Wijnbergen, S. van, 1983a, "Interest Rate Management in LDC's", Journal of Monetary Economics 12, 433-452.

Wijnbergen, S. van, 1983b, "Credit Policy, Inflation and Growth in a Financial Repressed Economy", Journal of Development Economics 13, 45-65.

Wijnbergen, S. van, 1985, "Macro-Economic Effects of Changes in Bank Interest Rates - Simulation Results for South Korea", Journal of Development Economics 18, 541-554.

Wooldridge, J.M., 1991, "Notes on Regression with Difference -Stationary Data", Michigan State University. Mimeo.

Zivot, E. and D.W.K. Andrews, 1992, "Further Evidence on the Great Crash, the Oil-Price Shock, and the Unit-Root Hypothesis", Journal of Business and Economic Statistics 10(No.3), 251-270.

Taiwanese Literature (in Chinese)

Chang (1989), C.Y., 1989, Empirical Studies on Demand for Money functions in Taiwan, Tan-Chiang University Press.

Ho, C.Y., 1981, "Bank Rates, Black Market Rates, and the Supply of Funds", Taiwan Economic Forecasting, 13(No.1), 45-64.

Hwang, J.T., B.L. Wu, 1994, "Testing for the Stability of Taiwan's Short-Run Money Demand Function: The Application of Fuzzy Time Series", Taiwan Economic Association Annul Conference Proceedings, 169-190.

Lee, C.N., 1996, "The Transactions Money Demand in Taiwan: Structure and Stability", Quarterly Review, Bank of Taiwan, 47(No.4), 64-86.

Lee, J.C., 1994, Monetary Control in Taiwan: Theory, Evidence, and Policy, Best-Wise Co.,Ltd.

Liang, M.Y., and Chen, M.L., 1985, "Relationships between Bank Rates and Black Market Rates", Chinese Economic Association Annual Conference Proceedings (November), 157-170.

Lin, C.Y., 1997, "The Demand for M2 in Taiwan", Quarterly Journal of the Central Bank of China, 40-70.

Lyo, F.C., 1970, "The Demand for Money in Taiwan", in C.S Chiu (1975) eds. Studies in Monetary Economics in Taiwan.

Shea, J.D., 1979a, "Bank Rates, Black Market Rates, Commissions, Income, and Economic Development", Academia Economic Review, the Institute of Economics Academia Sinica, Taipei, Taiwan, 7(No.2), 25-46.

Shea, J.D., 1979b, "Bank Rates, Black Market Rates, Income Per Capita, and Price", Chinese Economic Association Annual Conference proceedings December, 213-221.

Shea, J.D., 1983a, "Bank Credit, and the Supply of and demand for Funds in the Dual Financial System - Empirical Study in Taiwan", Economic Thesis 37, Chung-Hwa Institute for Economic Research.

Shea, J.D., 1983b, "The Income Elasticity of the Money Demand in Taiwan", Academia Economic Review 11(No.2), September, the Institute of Economics Academia Sinica, Taipei, Taiwan.

Shea, J.D., 1986, "The Effects of Monetary Policy in the Short Run under the Dual Financial System", Academia Economic Review 14(No.2), 111-154, the Institute of Economics Academia Sinica, Taipei, Taiwan.

Shea, J.D., 1991, "The Efficiency of the Allocation of Funds in the Taiwanese Financial System", reprinted in Shea (1995) Studies in Monetary Economics in Taiwan (eds.), 93-115.

Shen, C.H., 1995, "Estimation of the Money Demand Function Using a Correct Information Matrix", Taiwan Economic Review 23:2, 239-258.

Yang, Y.H., 1984, "Effects of the Regulation of Interest Rates under Dual Financial System", Economic Thesis, January, Chung-Hwa Institute for Economic Research.

Yuan, Y.S., 1986, "The Interest Rates in the Unorganized Money Market in Taiwan", Quarterly Review, Bank of Taiwan, vol.37:1, pp46-85, reprinted in Shea (1995)(eds.) Studies in Monetary Economics in Taiwan, 395-441.

Wu, T.M., 1995, "Taiwan's Interest Rates and Inflation: 1907-1986, Taiwan Economic Review 23(No.4), 419-444.