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Is money demand in Taiwan stable?☆

Chung-Shu Wu[a,b], Jin-Lung Lin[a,b,*], George C. Tiao[c], David D. Cho[c,1]

[a] Institute of Economics, Academia Sinica, No. 128, Sec. 2, Academia Rd. Nankang, Taipei 11529, Taiwan, ROC
[b] Department of Economics, National Chengchi University, Taiwan, ROC
[c] Graduate School of Business, University of Chicago, Chicago, USA

Abstract

Is money demand in Taiwan stable? Moreover, is money a luxury goods in Taiwan such that the income elasticity is greater than one? A casual application of Goldfeld type of money demand to the Taiwanese economy answers no to the first question and yes to the second one. This paper rigorously analyzes the money demand in Taiwan and attempts to provide more accurate answers to these questions. We employ both the ARMAX and cointegration models to study the money demand and use the rolling estimation approach to examine the stability of parameter estimates over time. Furthermore, we take into account the impact of stock market on money demand. Our empirical analysis concludes that the money demand in Taiwan is stable and that the income elasticity is less than one. Wrongly including a constant term within a dynamic model with lagged values of the dependent variable as regressors results in unstable estimates over time. In addition, the stock market is confirmed to have a significant impact on the demand of money.

JEL classifications: C22; C32; E41

Keywords: Goldfeld money demand; Cointegration; ARMAX model; Taiwan; Stability; Recursive estimate; Constant term

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* Corresponding author. Tel.: +886-2-27822791, ext: 313; fax: +886-2-27828673.
E-mail address: jlin@sinica.edu.tw (J.-L. Lin).
[1] *Present address:* School of Management, SUNY, Buffalo, NY, USA.

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1. Introduction

The stability of money demand function has long been the central proposition of monetary economics. The money demand function plays a key role in many economic models such as the New Classical approach, the New Keynesian analysis, and real business cycle models. See Sargent and Wallace (1975), Mankiw (1991) and King et al. (1991). In addition, the money demand function has been extensively studied empirically. For a few examples, see Goldfeld (1973), Judd and Scadding (1982), Laider (1985), Hendry and Ericsson (1991), Hoffman et al. (1995) and Nell (1999). In particular, Hoffman et al. (1995) and Hendry and Ericsson (1991) found that the money demand function is stable. However, Goldfeld (1973) and Nell (1999) found an opposite result and claimed that the money demand function may be unstable.

Since 1980, there have been a number of financial liberalization policy changes in Taiwan. Examples include interest rate deregulation, opening of a large number of new private banks, insurance companies and stock exchange companies, and gradual opening of foreign exchange market to outside investors. Hence it is reasonable to suspect that the money demand function might become unstable. Goldfeld (1973) specified a money demand function with lagged money, income and interest rate as explanatory variables. The Goldfeld type money demand or its variants have been extensively used in almost all previous empirical studies of money demand in Taiwan. Applying the Goldfeld type money demand function to Taiwan data would typically result in a declining long-run income elasticity. In response, many local researchers recognized this phenomenon and incorporated the financial deregulation or innovation to improve the specification of the money demand function. For examples, see Wu (1987), Lin (1997), Chen and Hu (1997), Wu (1998) and Ou and Lee (1999). However, it appears to us that the instability problem has not yet been rigorously studied. This paper adopts a rolling estimation approach to the following three models: the Goldfeld model, the ARMAX model and the cointegration model to shed light on the stability question. We also investigate the impact of stock market transactions on the money demand.

The remainder of the paper is organized as follows. Section 2 discusses Goldfeld type of money demand and describes the data. Section 3 presents our econometric approaches and analyzes the empirical results, and Section 4 concludes.

2. Goldfeld type of money demand

According to Baumol (1952) transaction motivation of money demand, the desired real money demand, $m^*=M/P$, where M is the nominal money supply and P denotes the general price index, can be expressed as a function of the real transaction y, and the opportunity cost of holding money, the interest rate i, that is

$$m^* = f(i, y). \tag{1}$$

Since there exist adjustment costs, actual money holdings are assumed to adjust linearly to the gap between desired holdings and last period's actual holdings. Thus,

$$m_t = m_{t-1} + \eta\left(m_t^* - m_{t-1}\right) \tag{2}$$

where η is the coefficient of adjustment.

Substituting Eq. (1) into Eq. (2), we can derive the well known Goldfeld (1973) short-run money demand function,

$$m_t = a_0 + a_1 m_{t-1} + a_2 y_t + a_3 i_t + \varepsilon_t, \tag{3}$$

where ε_t is the error term.

In what follows, all the variables considered in model equations will be expressed in logarithmic metric. It then follows from Eq. (3) that the long-run income (real transaction) and interest rate elasticities of money demand are $a_2/(1-a_1)$ and $a_3/(1-a_1)$, respectively. To evaluate if the Goldfeld type of money demand fits the Taiwanese economy, we estimate the rolling money demand using data from the first quarter of 1978 to the fourth quarter of 1999. The rolling estimation begins at the first quarter of 1987 and ends at the last quarter of 1999. We use M1B as the nominal money demand, real gross domestic product (GDP) as the real transaction, and 1-month time deposit rate as the opportunity cost of holding money. Three seasonal dummies are included in the regression models to control for seasonality effects. As is often observed in other empirical analyses, regression residuals of Taiwan data are detected to have strong serial correlations. We follow the conventional practice by fitting an AR(1) model to residuals and thus the model becomes,

$$m_t = a_0 + a_1 m_{t-1} + a_2 y_t + a_3 i_t + d_1 D_{1t} + d_2 D_{2t} + d_3 D_{3t} + \frac{\varepsilon_t}{1-\phi B} \tag{4}$$

where m_t: log of real M1B; y_t: log of real GDP; i_t: log of interest rates; D_{1t}, D_{2t}, D_{3t}: seasonal dummies; B: backshift (lag) operator.

2.1. Data

All series considered in this paper are seasonally unadjusted quarterly data taken from the AREMOS databank. The real GDP is measured by GDP at the 1996 constant price, the interest rate is 1-month time deposit rate of the First Bank, and money is the average of three end-of-the-month monthly money supply deflated by CPI. Fig. 1a, b and c plot the three series, M1B, GDP and interest rate, respectively, all in their original metric. In earlier part of the sample, interest rate was under government control and maintained at a high level. In the early 1980s, interest rate started declining, and bounced back in 1988 when deregulation and opening of a number of new banks occurred. GDP and M1B maintain a growing trend while the former displays a stronger pattern of seasonality. As mentioned earlier, all series are taken logarithmic transformation for model estimation. Finally, Fig. 1d plots the monthly stock volume, which will be discussed later.

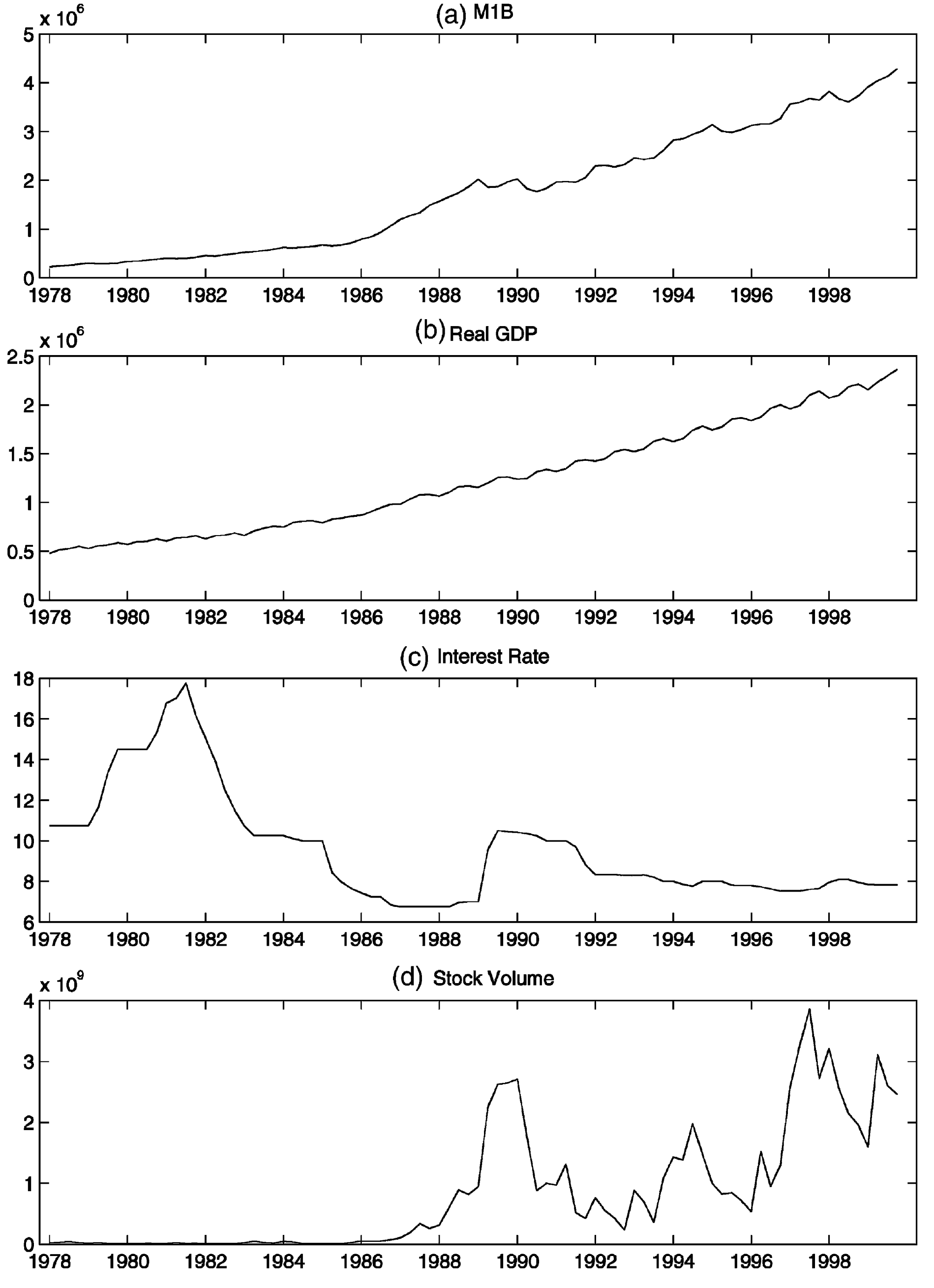

Fig. 1. Time series plots for real variables in original metric.

2.2. Estimation results

Since the AR(1) residuals in a model with lagged dependent variable make the OLS estimate inconsistent, we fit Eq. (4) using the Beach and MacKinnons' (1978) maximum likelihood estimation procedure of RATS. The resulting rolling estimates of long-run income and interest rate elasticities are plotted in Fig. 2. The figure shows that both the long-run income and interest rate elasticities have a decreasing trend. The income elasticity declines from 1.5 to approximately 0.7, whereas the interest rate elasticity decreases from −0.4 to −1.7. The empirical results that income elasticity is larger than unity is commonly found in Taiwan money demand literature, e.g. Liu (1970), Liang et al. (1982) and Lin (1997). Furthermore, Fig. 3 shows that the velocity of M1B, or the ratio between the nominal GDP and money supply, kept declining since 1970s. This phenomenon has led researchers to firmly believe that the income elasticity in Taiwan is greater than unity. However, it is hard to interpret why Taiwanese people regard real money balance as luxury goods.

It is known that adding a wealth variable to the money demand function reduces the income elasticity. Bomberger (1993) and Fase and Winder (1998) are two such examples. However, we do not follow this approach for two reasons. First, since

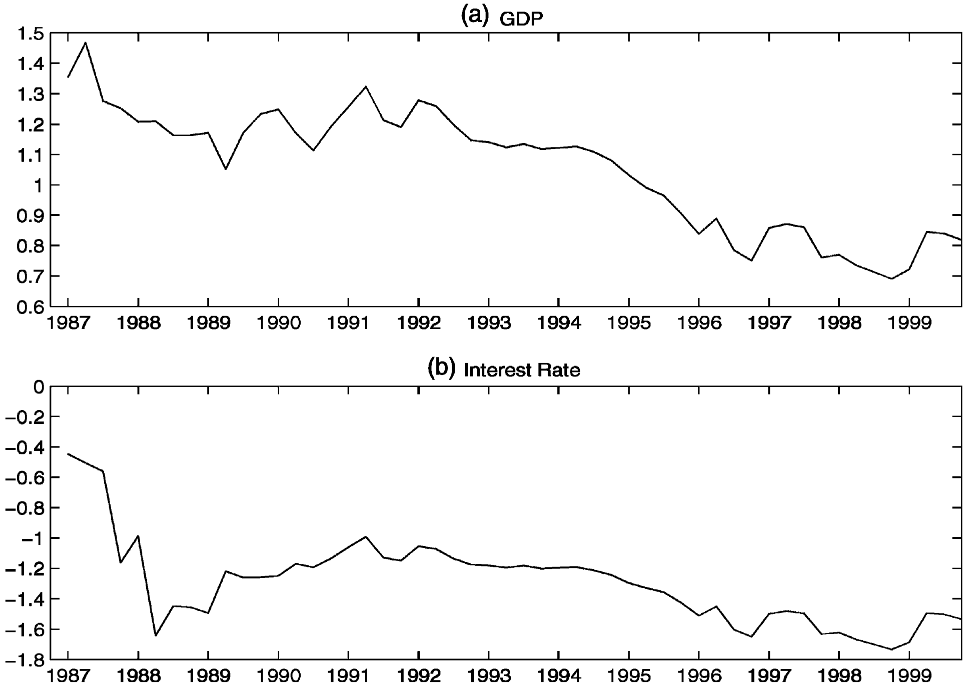

Fig. 2. Long-run elasticity of M1B:AR(1), no stock volume, constant.

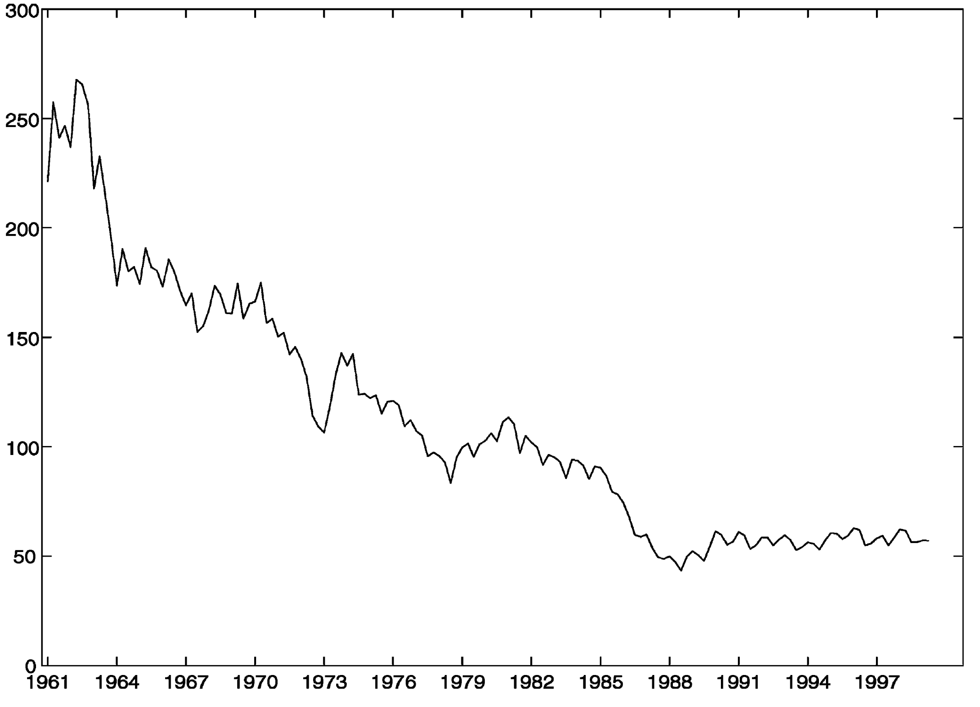

Fig. 3. Velocity of M1B.

survey data of wealth in Taiwan do not exist, any wealth proxy has to be imputed from other variables. There is no general agreement on the best imputing procedure and the accuracy of all existing proxies has been seriously challenged. Second, Wu and Shea (1993) used real estate transaction to construct a wealth proxy and included it in the money demand function. The resulting income elasticity is smaller than that without the wealth variable but is still greater than one.

Financial transactions, especially those at the stock market, play an important role in money demand but only a small part of these transactions is counted in real gross domestic products. See Friedman (1988), Palley (1995) and Choudhry (1996). It is natural to incorporate stock market transactions in the money demand function to improve the specification. See Wu and Shea (1993) and Wu (1995). Even though other financial transactions exist, such as bond, derivatives, and futures, limited by the data availability, we take into account the stock market transactions, which are the major and the largest financial transactions. Real stock transaction volume is the volume of TAIEX deflated by consumer price index (CPI), where TAIEX is the Taiwan stock exchange capitalization weighted stock index compiled by the Taiwan Stock Exchange. The times series plot of real stock transaction volume in Fig. 1d shows that the stock transaction volume maintains an increasing trend with a large fluctuation.

To investigate whether the missing financial variable is the cause of the instability of money demand and the income elasticity being greater than one, we add the stock transaction volume into the model and re-perform the rolling estimation.

$$m_t = a_0 + a_1 m_{t-1} + a_2 y_t + a_3 i_t + a_4 s_t + d_1 D_{1t} + d_2 D_{2t} + d_3 D_{3t} + \frac{\varepsilon_t}{1 - \phi B}, \qquad (5)$$

where s_t: log of stock transaction volume.

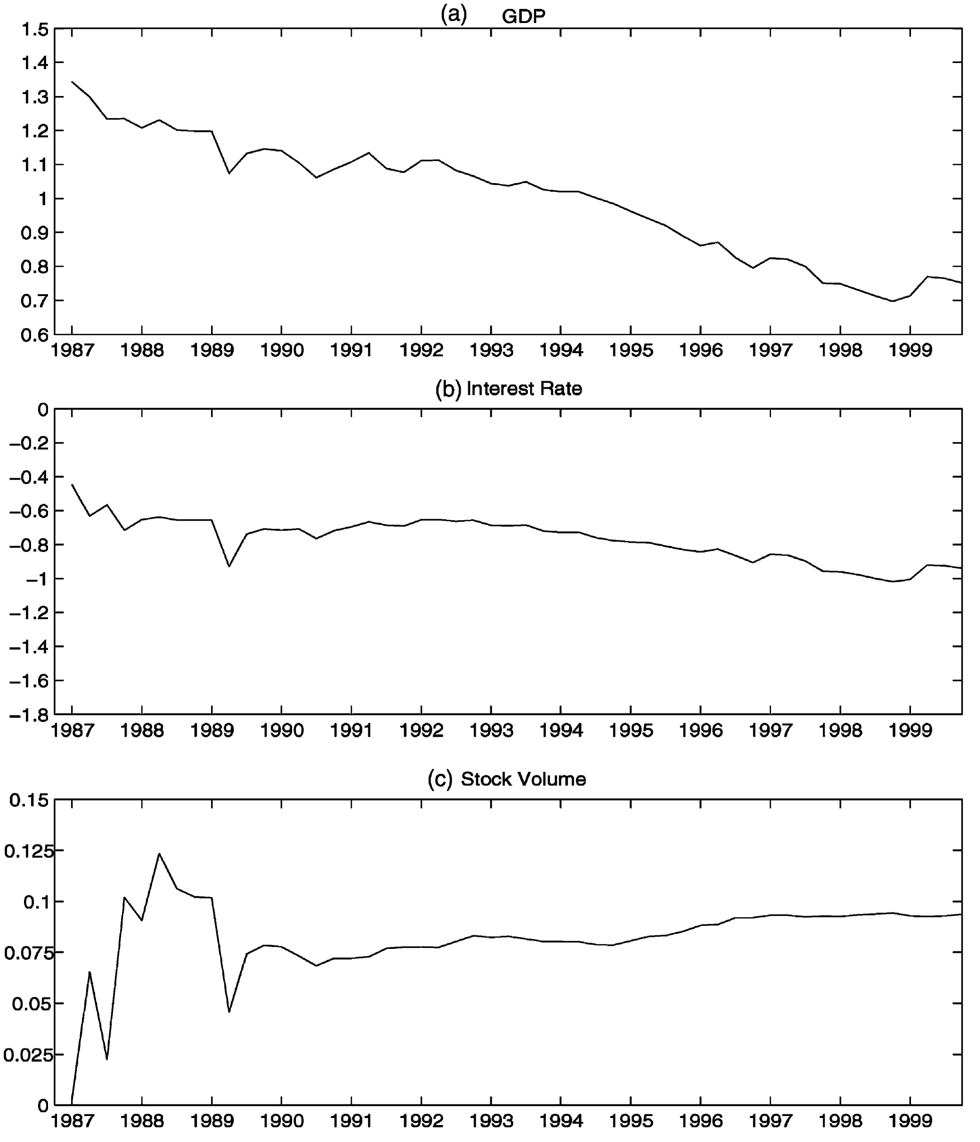

Fig. 4. Long-run elasticity of M1B:AR(1), stock volume, constant.

Fig. 4 reports the estimation results of Eq. (5). Adding the stock transaction volume reduces the income elasticity at the early sample period but does not fix the above two problems. The income elasticity still declines from approximately 1.3 to 0.7, and interest rate elasticity declines from -0.4 to -1.0. As expected, the positive stock volume elasticity implies that the stock transaction volume has a positive effect on the money demand.

The empirical results suggests that there still exists a declining trend in the long-run income elasticities even after taking into account the stock transactions. Moreover, the same results apply to the interest rate elasticities. To summarize, if one uses the Goldfeld type of money demand to estimate the elasticities, the money demand function in Taiwan appears to be unstable even controlling for financial transactions.

3. Econometric methodology

Non-stationarity is one of the key questions in estimating the money demand equation. Economists' typical approach is to assume autoregressive models for the time series, and to test for the existence of unit roots. If confirmative, then they proceed with a cointegration analysis and, in particular, with the Johansen maximum likelihood estimation method, which is built upon a vector autoregression (VAR) representation. However, statistics-oriented approach such as the ARMAX model stresses the importance of white noise error and is more willing to adopt the mixed autoregressive-moving average error term whenever the model identification points to this way. However, cointegration is less emphasized as it is believed to be sensitive to random level shifts. See Chen and Tiao (1990). In this paper, we perform both approaches and compare the empirical results.

3.1. ARMAX models

The ARMAX model for an output variable z_t on k inputs $x_{1t},\ldots,x_{kt}$ takes the form:

$$\psi(B)z_t = \sum_{i=1}^{k} \phi_i(B)x_{it} + \pi(B)\varepsilon_t, \tag{6}$$

where $\psi(B)$, $\phi_i(B)$ and $\pi(B)$ are rational polynomials in B, and ε_t is the white noise term. For ease of comparing Eq. (6) with the Goldfeld type of money demand described earlier and the cointegration model in the next section, we specifically limit the candidate models to those, which include lagged money demand in $\psi(B)$. As is pointed out by Stephen Hall that regular and seasonal differencing are likely to wash off structural changes in the money demand, we do not take difference unless the residual diagnostic checking suggests so doing.

The ARMAX model is developed by the following steps. First, the choice of $\psi(B)$ and $\phi_i(B)$ are often guided by subject matter considerations such as the money demand equation in Eq. (3). Next, ACF, PACF and ESACF (extended sample autocorrelation function) are used to check whether the residuals behave like a white noise series as suggested by Tsay and Tiao (1984). If not, an appropriate model structure for the residuals can be identified by the ESACF results. Then, all the parameters are estimated simultaneously and the outlier effects are diagnostically checked. If outliers are detected, then one can re-estimate to adjust for the outlier effect. Finally, the residuals are diagnostically checked for whiteness. One can repeat the above modeling process until all the diagnostic checks are passed. See Chang et al. (1988) for details.

3.2. Cointegration model

Let Y_t be a 4×1 vector generated by a vector autoregressive process of order k,

$$\Delta Y_t = \mu + \delta t + \Phi D_t + \Pi Y_{t-1} + \Gamma_1 \Delta Y_{t-1} + \cdots + \Gamma_{k-1} \Delta Y_{t-k+1} + \varepsilon_t \tag{7}$$

where $Y'_t = [m_t \;\; y_t \;\; i_t \;\; s_t]$; D_t: vector of dummy variables.

Assume that the characteristic roots for Y_t are either on or outside the unit circle. As is indicated by the Granger representation, the rank of matrix Π, denoted as r, determines the long-run property of Y_t. If $0<r<4$, then there exist two 4 by r matrices α, β, such that $\Pi=\alpha' \beta$. Under the assumption that all the series in Y_t are at most $I(1)$, then there exist r cointegration vectors, β, such that $\beta' Y_t$ is $I(0)$. Johansen (1995) derived the maximum likelihood estimates for α, β and other parameters under the null hypothesis that there are r cointegration vectors and also obtained the limiting distribution of the likelihood ratio test statistic of having r cointegration vectors.

The appearance of the error correction term makes the cointegration model different in form from the conventional vector ARIMA model. The difficulty of interpreting cointegration vector arises when there are more than one cointegration vector. In such a case, further restrictions need to be imposed to make the structural cointegration vector identifiable. See Hall and Zonzilos (1999) for details. Moreover, due to the large number of parameters, applying Johansen procedure to a small sample usually leads to noisy estimation results. Small changes in model specification sometimes result in very different estimates.

Table 1
Estimation results of ARMAX model

Variable	Parameter	Coeff.	Std.	t-ratio
m_{t-1}	a_1	0.8151	0.0289	28.17
y	a_2	0.1466	0.0189	7.39
i	a_3	−0.1568	0.0199	−7.87
s	a_4	0.0171	0.0054	3.16
MA(1)	θ_1	−0.3365	0.1011	−3.33
D_1	d_1	0.0235	0.0064	3.64
D_2	d_2	−0.0494	0.0076	−6.49
D_3	d_3	−0.0327	0.0063	−5.16

Table 2
ESACF for ARMAX residuals

		Q												
		0	1	2	3	4	5	6	7	8	9	10	11	12
P	0	0.06	0.18	0.02	0.04	0.16	−0.07	0.08	−0.08	0.07	−0.12	−0.06	−0.06	−0.10
	1	−0.30	0.18	−0.08	−0.00	0.15	0.04	0.01	−0.01	−0.02	−0.13	0.03	0.00	−0.01
	2	0.00	−0.03	−0.04	−0.01	0.15	0.03	0.01	−0.01	−0.01	−0.12	0.06	−0.00	−0.01
	3	0.02	0.01	−0.41	−0.01	0.15	−0.05	−0.01	−0.02	−0.01	−0.11	0.05	−0.00	0.01
	4	−0.03	0.01	−0.42	−0.05	0.16	−0.03	−0.01	−0.01	−0.00	−0.10	0.00	−0.04	−0.04
	5	0.42	0.12	0.17	0.07	0.29	−0.06	0.02	−0.05	0.05	−0.00	0.00	−0.02	−0.01
	6	0.28	−0.29	0.17	−0.04	0.06	0.23	0.03	−0.05	0.07	−0.02	−0.01	−0.02	−0.01
Simplified extended ACF (5% level)														
P	0	O	O	O	O	O	O	O	O	O	O	O	O	O
	1	X	O	O	O	O	O	O	O	O	O	O	O	O
	2	O	O	O	O	O	O	O	O	O	O	O	O	O
	3	O	O	X	O	O	O	O	O	O	O	O	O	O
	4	O	O	X	O	O	O	O	O	O	O	O	O	O
	5	X	O	O	O	X	O	O	O	O	O	O	O	O
	6	X	X	O	O	O	O	O	O	O	O	O	O	O

'X' mean significant and 'O' insignificant.

3.3. Relationship between ARMAX and cointegration models

It is interesting to find the relationship between the ARMAX and the cointegration models. To be more specific, under what conditions would these two approaches generate the same results? Phillips (1991) has derived the optimal inference in cointegrated systems. If all other variables are strongly exogenous for money and there is only one

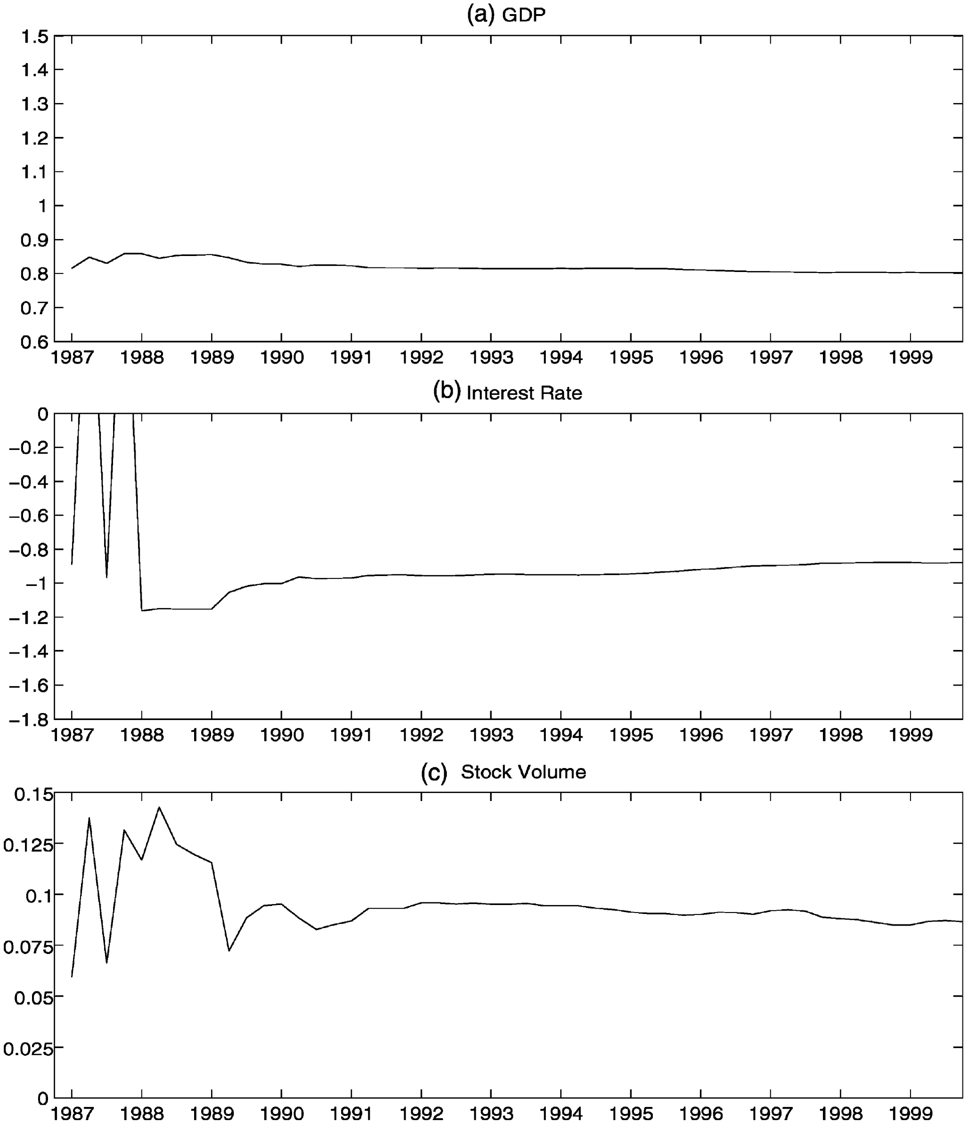

Fig. 5. Long-run elasticity of M1B:MA(1), stock volume, no constant.

cointegration vector, then the OLS estimates and the maximum likelihood estimates will have the same asymptotic distribution. However, parameter estimates in ARMAX model and in the cointegration model would converge to the same limiting distribution. If either of these two conditions fails to hold, then the limiting distribution of estimates in ARMAX model will contain unit root distribution and other components. However, while the equivalence holds in large sample, the discrepancy could be large for small sample. Furthermore, MA terms play an important role in ARMAX modeling but Johansen's MLE

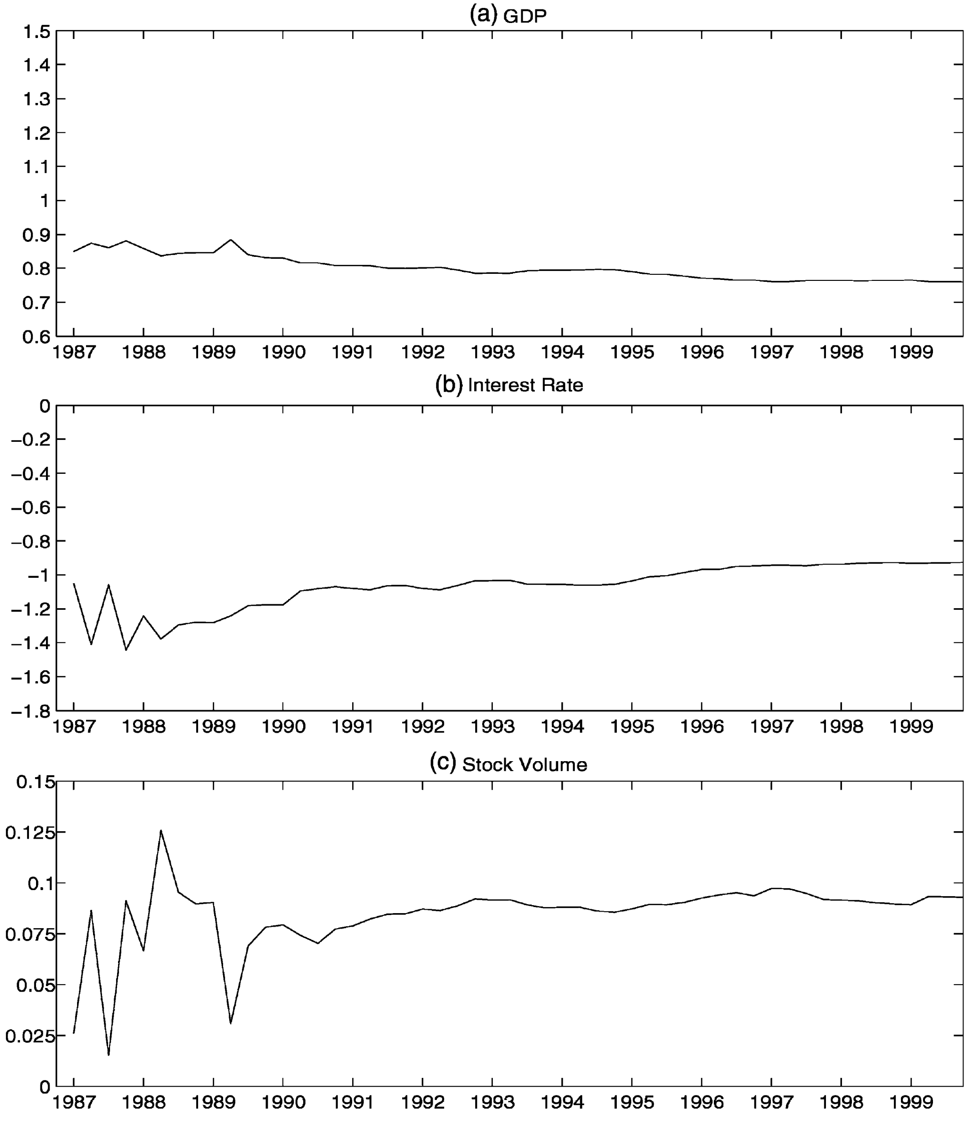

Fig. 6. Long-run elasticity of M1B:AR(1), stock volume, no constant.

is confined to VAR. Finally, there has been some work done extending the Johansen procedure to a VARMA framework. See Hunter and Dislis (1996) and Wagner (1999).

3.4. Empirical results

This subsection presents the estimation results for the econometric modeling approach. First, the empirical results from the ARMAX model are presented. We specifically retain

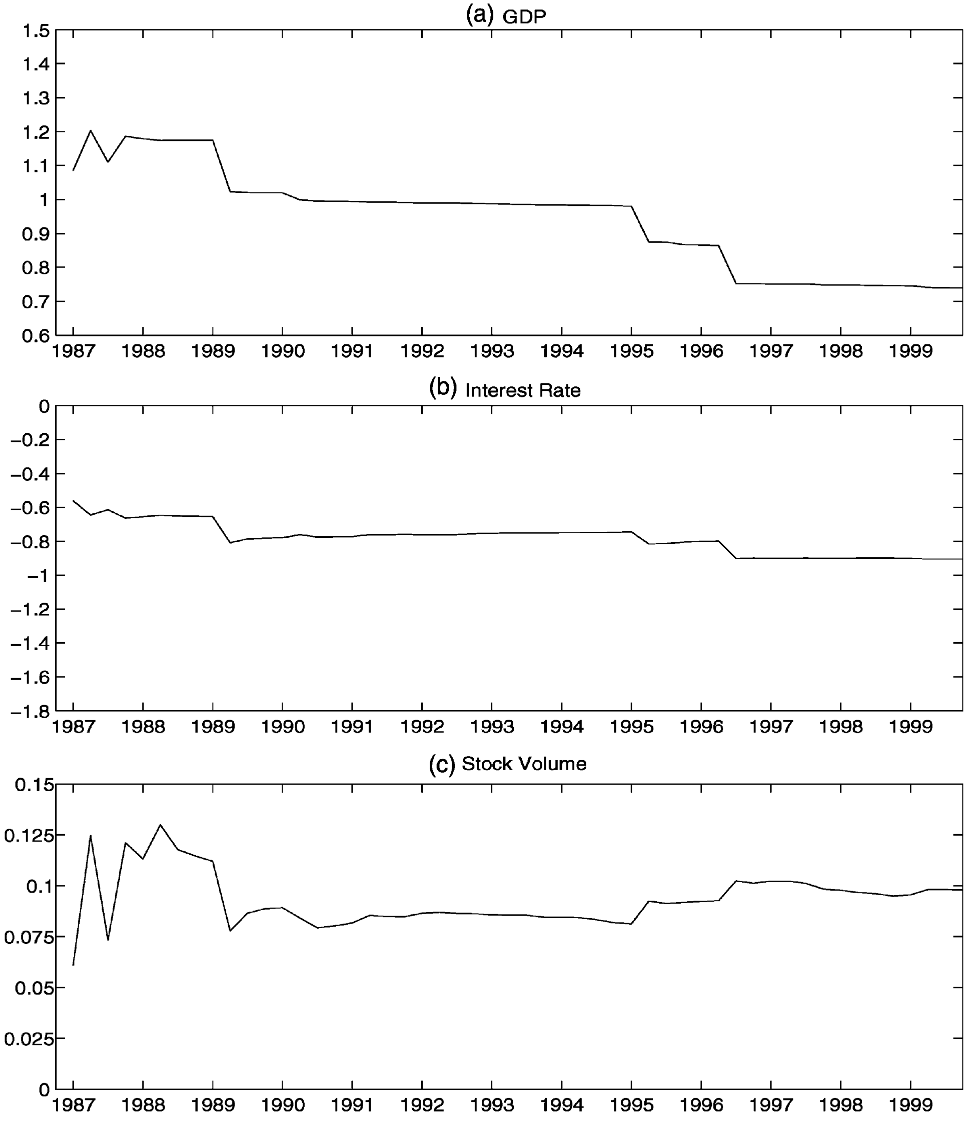

Fig. 7. Long-run elasticity of M1B:MA(1), stock volume, constant.

the lagged money demand variable as a regressor to compare it with the Goldfeld money demand. The model specification procedure results in the following model:

$$m_t = a_1 m_{t-1} + a_2 y_t + a_3 i_t + a_4 s_t + d_1 D_{1t} + d_2 D_{2t} + d_3 D_{3t} + \theta_1 \varepsilon_{t-1} + \varepsilon_t, \tag{8}$$

where $\theta_1 \varepsilon_{t-1}$ is the lag one moving average MA(1) term. Table 1 reports the estimation results for the entire sample.

It is worth noting that the income elasticity is approximately 0.79[=0.1466/(1−0.8151)], while the interest rate elasticity is significantly negative, −0.8480[=−0.1568/(1−0.8151)]. As expected, the stock market transaction poses a significantly positive effect on money demand with elasticity being 0.09[=0.0171/(1−0.8151)]. The MA coefficient is significant, which justifies its appearance. Most importantly, the constant term, a_0, is insignificant and hence is dropped from the model. The ESACF for the residuals is summarized in Table 2, which is consistent with a white noise series.

Fig. 5 reports the results of rolling estimations for the ARMAX model. Again, the sample end point for the rolling estimations starts from the first quarter of 1987 to the fourth quarter of 1999. The figure indicates that the long-run income elasticity varies stably within the narrow range from 0.80 to 0.86 throughout the whole rolling estimation period. This is in sharp contrast with the rolling estimation results shown in Fig. 2 and Fig. 4. The variation of long-run income elasticity before 1991 is relatively larger but stabilizes considerably since then. Similar results extend to the long-run interest rate and stock transaction volume elasticities. The relatively larger variation of money demand elasticity before 1991 can be explained by the change of financial system and the financial regulation before 1991. From these estimates, we can draw a conclusion that the money demand in Taiwan is stable, at least after 1991, and that the income elasticity is less than one. An important question then arises: what causes the markedly different results between the Goldfeld type of money demand discussed in Section 2.2 and the ARMAX model of money demand specified in Eq. (8)? A simple comparison reveals the differences in model specifications. First, the ARMAX model contains an extra variable, the stock transaction volume, and the conventional Goldfeld model does not. Second, the ARMAX model uses MA(1) model for the errors while the Goldfeld model adopts an AR(1) model. Third, the ARMAX model does not have a constant term but the Goldfeld model does.

To determine the real cause, we have estimated two additional models: a Goldfeld model with stock transaction volume included but without the constant term, and an ARMAX model with constant. The results of rolling estimation are reported in Figs. 6

Table 3
Unit root test results

Level	τ_μ	First difference	τ_μ
M1B	−1.160	ΔM1B	−3.889†
GDP	−1.450	ΔGDP	−4.447†
I	−1.650	ΔI	−5.657†
S	−1.031	ΔS	−8.127†

† denotes significance at 5% level.

Table 4
Tests for cointegration rank and deterministic terms

r	$n-r$	Model 1	Model 2	Model 3	Model 4	Model 5
Eigen values						
0	4	0.2585	0.2587	0.2543	0.2693	0.2678
1	3	0.1683	0.2481	0.1439	0.1724	0.1623
2	2	0.0783	0.1045	0.0447	0.0926	0.0909
3	1	0.0142	0.0287	0.0195	0.0244	0.0114
Trace test						
0	4	48.062	60.087	42.680	51.866	49.431
1	3	23.242	35.246	18.330	25.822	23.560
2	2	7.951	11.581	5.433	10.113	8.858
3	1	1.187	2.419	1.635	2.049	0.949

Source: Hansen and Juselius (1995) Appendix B, Tables B1–B5.

and 7, respectively. A quick examination would lead to the discovery that it is the omission of the insignificant constant term that stabilizes the parameter estimates. Adding the stock transaction variable and replacing the AR model for the residuals with a MA model do have some effect but not in a fundamental way. This seems a surprising answer, as it is the economist's tradition to include the constant even when it is insignificant. Furthermore, the constant term has been used as a safe guard against neglecting important variables. The role of the constant term within a dynamic model is an important issue and deserves further research. To briefly summarize, the ARMAX empirical results support the existence of a stable money demand in Taiwan.

Next, the empirical results of cointegration analysis are presented. Table 3 reports the unit root test results, which clearly supports the existence of a unit root. It is well known that the form of deterministic terms of VAR will affect the limiting distribution of the

Table 5
Diagnostic checking statistics

Univariate statistics				
Eqs.	S.D.	Skewness	Kurtosis	Normality χ^2
Fig. 1	0.0257	0.0377	3.2961	1.477
Fig. 2	0.0433	2.4854	18.0335	41.440
Fig. 3		0.3248	3.3468	2.251
Eq. 4		0.0721	2.5508	0.338
Multivariate statistics				
Test for white noise				
LB(20)	CHISQ(252)=279.656			P-val=0.11
LM(1)	CHISQ(16)=17.740			P-val=0.34
LM(4)	CHISQ(16)=16.235			P-val=0.44
Normality				
	CHISQ(14)=43.647			P-val=0.00

Table 6
Cointegration vector and loading matrix

m	y	i	s
Cointegration vectors: β'			
1.000	−0.295	0.538	−0.250
Loading Coef. α'			
0.054	−0.131	−0.017	1.657

cointegration rank test statistics. More specifically, there are five models depending upon the deterministic terms,

1: μ=0, δ=0;
2: $\mu = \alpha\beta_0'$, $\delta = 0$;
3: $\mu_0 \neq 0$, δ=0;
4: $\mu_0 \neq 0$, $\delta = \alpha\beta_1'$; and
5: $\mu_0 \neq 0$, $\delta \neq 0$.

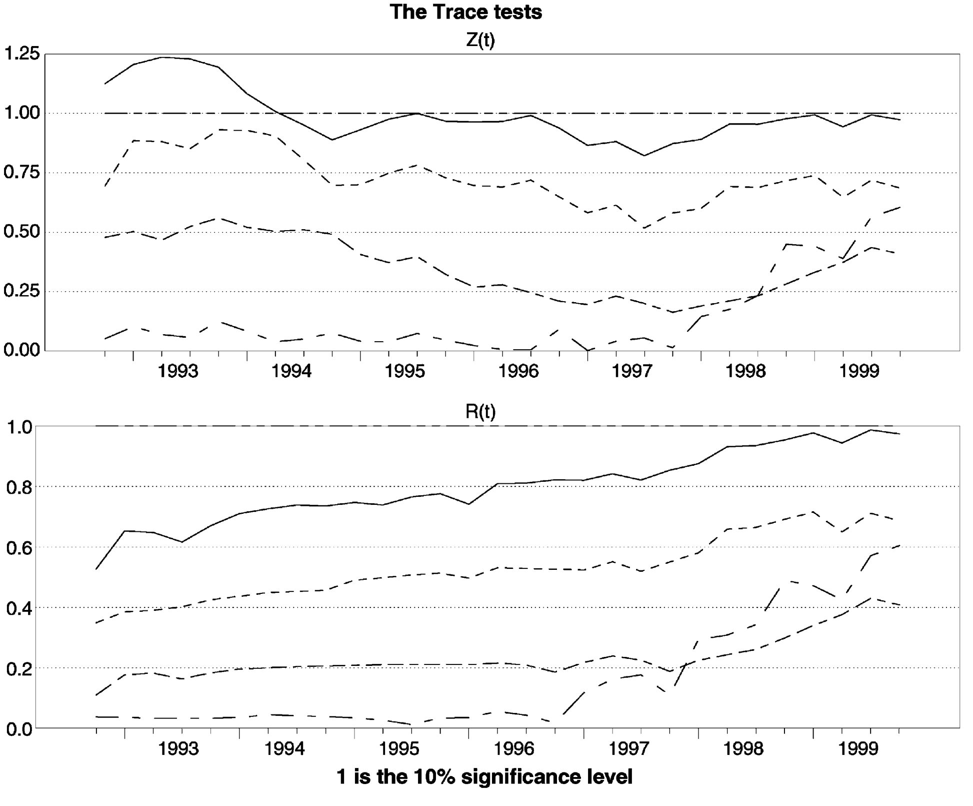

Fig. 8. Rolling estimate of trace statistics.

Let $M(i, j)$, i=0, 1,. . .,n, j=1,2,3,4,5 denote model i with j cointegration vectors. Then to obtain a proper size, we should start testing the most restrictive $M_{0,1}$ using either Trace or $L_{\max}$ statistics. If the null hypothesis is rejected, then we should proceed with the order, M(0,1), M(0,2), M(0,3), M(0,4), M(0,5), M(1,1),. . .,M(n,5) until the null hypothesis is not rejected.

Table 4 summarizes the rank test results. The test results lead to M(0,3), which is a model with an unrestricted constant term and no cointegration vector. However, M(0,3) is not rejected at 90% and $\lambda_{\max}$ statistics suggest one cointegration vector. Table 5 reports the univariate statistics for all four series under the assumption of one cointegration vector. The diagnostic checking statistics indicate that the residuals behave like a white noise series. Even though the normality test is badly rejected, it is a typical case for the data in Taiwan. It is worth mentioning that removing the stock transactions variable from the model would lead to a clear no cointegration result.

Table 6 reports the estimates of the cointegration vector and the loading matrix. Note that all four parameters have the correct signs as theory predicts. The long-run income elasticity of money demand is positive but is as low as 0.295. The interest rate has a negative effect and the stock market transaction has a positive effect.

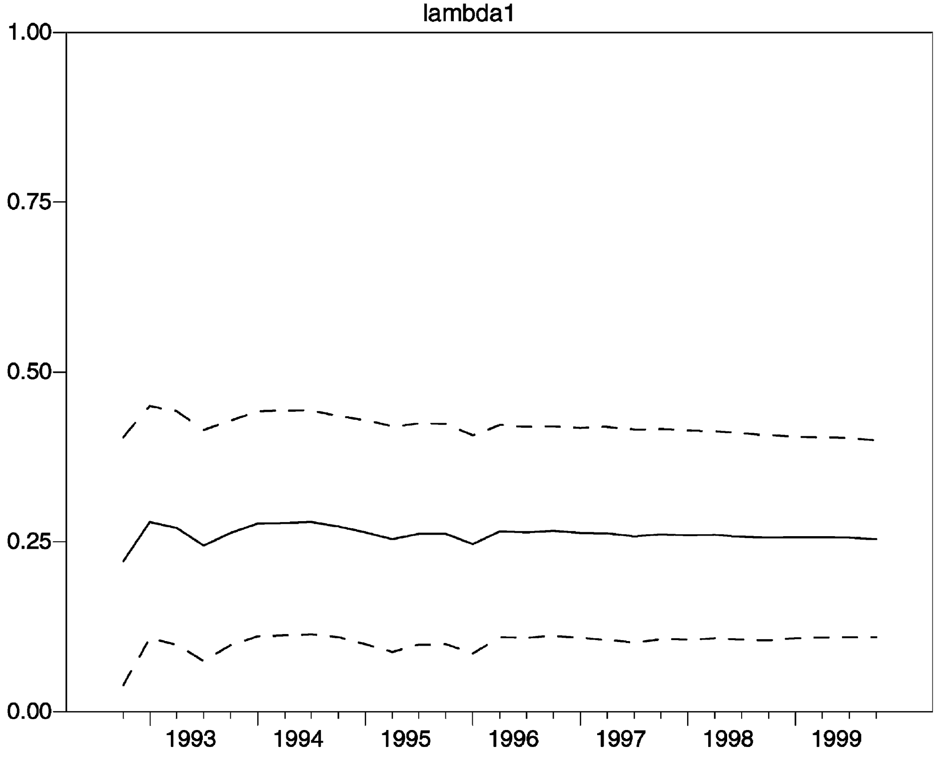

Fig. 9. Rolling estimate of eigen values.

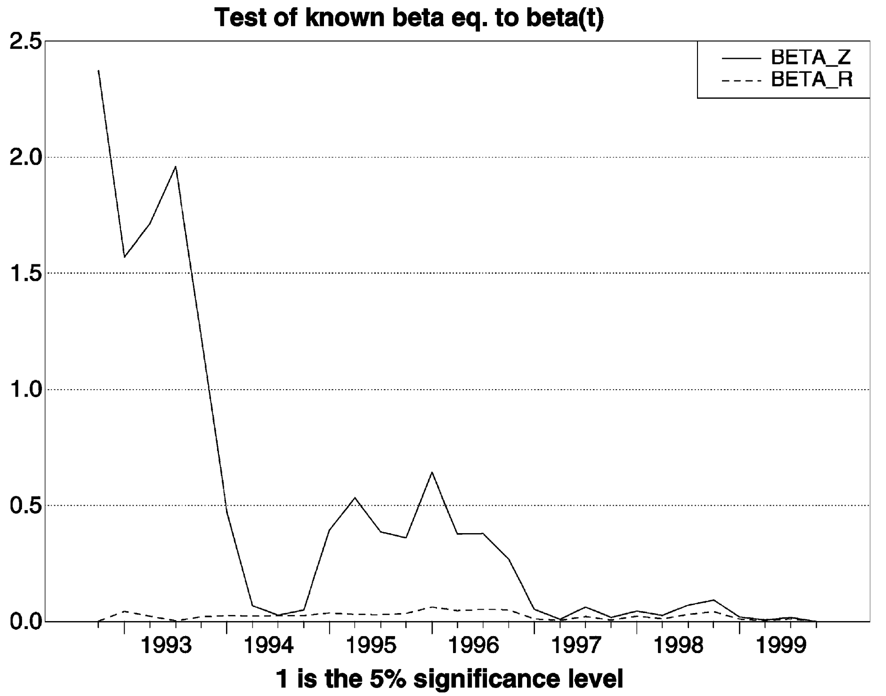

Fig. 10. Rolling estimate of β.

To further examine the stability of cointegration estimate, we perform the rolling estimation and report the results in Figs. 8–10. The figures imply that the estimates display instability at early 1993 but remain stable since then. This result seems to be consistent with the above ARMAX model results.

4. Conclusion

We use the ARMAX modeling and cointegration modeling approaches to analyze the stability of the money demand in Taiwan. Both models confirm the importance of stock market transactions in specifying the money demand. When the money demand function is properly specified, the income elasticity is less than one. Moreover, the stability analysis for both models support the existence of stable money demand function. Wrongly including a constant term within a dynamic model with a lagged dependent regressor results in unstable elasticity estimates over time.

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