Is Money Demand in Taiwan Stable? [1]

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Abstract

Is the money demand in Taiwan stable? Is money a luxury goods in Taiwan such that income elasticity is greater than one ? A causal application of Goldfeld type of money demand to Taiwanese economy gives 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 estimate to examine the stability of parameter estimates over time. Further, we take into account the impact of stock market on money demand. Our empirical analysis concludes that money demand in Taiwan is stable and income elasticity is less than one. Wrongly including constant term within a dynamic model with lagged dependent regressor results in unstable estimates over time. Also, stock market is confirmed to have a significant impact on money demand.

1 Introduction

Stability of money demand function has long been the central proposition of monetary models. This function plays a key role in the New Classical view of monetary effects, New Keynesian analysis and some real business cycle models. See Sargent and Wallace(1975), Mankiw (1991), King, Plosser, Stock and Watson (1991). Empirical literature in this area have been abundant. For a few examples, see Goldfeld (1973), Judd and Scadding (1982), Laidler (1985), Hendry and Ericsson (1991), Hoffman, Rasche and Tieslau (1995), and Nell (1999). Some did find a stable money demand function, such as Hoffman, Rasche and Tieslan (1995) and Hendry and Ericsson (1991). Some others, like Goldfeld (1976) and Nell (1999), claimed that there might exist an instability in money demand function. In the case of Taiwan, there have been a number of financial liberalization policy changes since 1980, e.g. interest rate deregulation, opening of a large number of new private banks, insurance companies and stock exchange companies, and gradual opening of money and foreign exchange market to outside investors, and, hence, it is reasonable to suspect that the money demand function might become unstable. Applying Goldfeld type money demand to Taiwan's data would typically lead to the result of declining long-run income elasticity. In response, many local researchers recognized this possible change and incorporated financial deregulation or innovation to improve the specification of the money demand function. For example, see Wu (1987), Lin (1997), Chen and Hu (1997), Wu(1998), Ou and Lee (1999). However, it appears to us that the instability problem has not yet been rigorously studied. In this paper we adopt a rolling estimation approach to three models: Goldfeld model, ARMAX model, and cointegration model to shed light on the stability question. We also investigate the impact of stock market transactions on 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 econometric methodology and analyze the empirical results and Section 4 concludes.

2 Goldfeld Type of Money Demand

According to Baumol's (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 real transaction y, and 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,

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

where η is the coefficient of adjustment.

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

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

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 equation (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 nominal money demand and real gross domestic product (GDP) as real transaction and one-month time deposit rate as 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 correlation. We follow the conventional practice by fitting an AR(1) model to residuals and thus the model becomes,

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

where m_t : log of real M1B

y_t : log of real GDP

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. Real GDP is measured by GDP at 1996 constant price, interest rate is one-month time deposit rate of the First Bank. Money is the average of three end-of-the-month monthly money supply deflated by CPI. Figures 1 reports the figures of all three series in their original metric. During earlier part of the sample, interest rate was under government control and maintained stably 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 maintains a growing trend while the former displays a stronger pattern of seasonal pattern. As mentioned earlier, all series are taken logarithm transformation in model fitting.

2.2 Estimation results

We fit equation (4) using the Maximum Likelihood Estimation procedure of RATS. The resulting estimates of long-run income and interest rate elasticities are plotted in Figure 2.

It can be noted from the figure that both the long-run income and interest rate elasticities have a decreasing trend. The income elasticity declines from 1.5 to around 0.6, 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, Chen and Lin (1982) and Lin (1997). Furthermore, from Figure 3, it can be noted that the velocity of M1B 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 a luxury goods.

Financial transactions, especially those at the stock market, play an important role 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). There exists other financial transactions in Taiwan financial sector such as bond, derivatives, and futures market. Limited by the data availability, we take into account the major and the largest financial transactions which are the stock market transactions. Real stock transaction volume is volume of TAIEX deflated by consumer price index (CPI). The times series plot of real stock transaction volume in Figure 1 shows that the stock transaction volume maintains an increasing trend with large fluctuation.

To verify if the instability of money demand and the income elasticity being greater than one are caused by the missing financial variable, we add stock transaction volume into the model and re-perform rolling estimation, that is

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

where s_t : log of stock transaction volume.

The estimation results are reported in Figure 4. Adding stock transaction volume slightly reduced the income elasticity at early sample period but does not cure the two problems above. Income elasticity still consistently declines from about 1.3 to 0.7 and interest rate elasticity declines from -0.4 to -1.0. As expected, stock transaction volume has a positive effect on money demand.

From the empirical results it can be noted that although long-run income elasticities show reasonable magnitudes after taking into account the stock transactions, there still exists a declining trend. Moreover, interest rate elasticities also have similar results. This implies that the money demand (M1B) function of Taiwan may be unstable, and the instability would not change even by controlling for important financial transactions.

3 Econometric Methodology

Nonstationarity is one of the key question in estimating the money demand equation. Economist's typical approach is to assume autoregressive models for the time series under investigation, test for existence of unit roots and, if confirmative, then proceed the analysis with cointegration analysis, particularly the Johansen maximum likelihood estimate method which is built upon vector autoregression representation. On the other hand, statistics-oriented approach such as ARMAX stresses the importance of white noise residual and are more willing to adopt the moving average error term whenever the model identification points to this way. 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 empirical results.

3.1 ARMAX model

The ARMAX model for an output variable z_t on k inputs $x_{1t}, ..., x_{kt}$ are:

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

where B is the backshift, or lag, operator, $\psi(B)$, $\phi_i(B)$ and $\pi(B)$ are rational polynomials in B, and ϵ_t is the white noise term. For ease of comparing with Goldfeld type of money demand described above 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 Steven Hall that regular and seasonal differencing are likely to wash off structural changes of money demand, we do not take difference unless the residual diagnostic checking suggests so doing.

A model is specified following the steps below. First, in econometric application, choice of $\psi(B)$ and $\phi_i(B)$ are often guided by subject matter considerations such as the money demand equation as in (3). Next, use ACF, PACF and ESACF (Extended Sample Autocorrelation Function) to check if the residuals behave like the white noise. See Tsay and Tiao (1984). If not, appropriate model structures for the residuals are identified. Next, all parameters are estimated simultaneously and outlier effects are diagnostically checked. If outliers are detected, then re-estimate to adjust for outlier effect. Finally, diagnostically check the whiteness of the new residuals. Repeat the modeling process until all diagnostic checks are passed. See Chang, Tiao, and Chen (1988) for details.

3.2 Cointegration Model

Let Y_t be a 4x1 vector generated by a vector autoregressive process of order k,

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

$$\begin{aligned} \text{where } Y_t' &= \begin{bmatrix} m_t & y_t & r_t & s_t \end{bmatrix} \\ D_t &: \text{ vector of dummy variables} \end{aligned}$$

Assume that the characteristic root for Y_t are either on or outside unit circle. As is indicated by Granger Representation, the rank of matrix Π, denoted as r, determines the long-run property of Y_t. If $0 < r < p$ then there exist two pxr matrices α, β, such that $\Pi = \alpha' \beta$. Under the assumption that all 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) has derived the maximum likelihood estimates of β, α and other parameters under the null hypothesis that there is r cointegration vectors and the limiting distribution of the likelihood ratio test statistics of having r cointegration vector.

The appearance of error correction term makes cointegration model different from conventional vector ARIMA models. Difficulty of interpreting cointegration vector arises when there are more than one cointegration vector. In such case, further restrictions need to be imposed to make the structural cointegration vector identifiable. See Hall and Zonzilos (1999) for details. Also, due to large amount of parameters, applying Johansen procedure to short series of small sample size usually leads to unstable estimation results. Small change of model specification might sometimes result in very different estimation results.

3.3 Empirical results

We now turn to the estimation results for the ARMAX modeling approach. We specifically retain lagged money demand as a regressor to compare with Goldfeld money demand. The model specification procedure result in the following model:

$$m_t = a_1 m_{t-1} + a_2 y_t + a_3 i_t + a_4 S_t + d_1 D_{1t} + d_2 D_{2t} + d_3 D_{3t} + \theta_1 \epsilon_{t-1} + \epsilon_t \tag{8}$$

where s_t is the stock transaction volume and D_{1t}, D_{2t}, D_{3t} are three seasonal dummies which appear in all models considered in this paper. Estimation results for the whole sample are put in Table 1.

It is worth noting that the income elasticity is around 0.79 [=.1466/(1-.8151)] while interest rate elasticity is significantly negative, -0.8480 [=-.1568/(1-.8151)]. As expected, the stock market transaction poses a significantly positive effect on money demand with elasticity being 0.09 [=.0171/(1-.8151)]. The MA terms is significant, which justifies its appearance. Mostly importantly, the constant term is insignificant and is, hence, dropped from the model. The ESACF for the residuals is summarized in Table (2), which is consistent with white noise.

Rolling estimates for ARMAX model with sample end points starting from 1987Q1 to 1999Q4 are summarized in Figure 5. From this figure, we find a surprising result– the long-run income elasticity varies stably within the range from 0.86 and 0.80 throughout the whole rolling estimation period. The variation of long-run income elasticity before 1991 is relatively large but stabilizes considerably since then. Similar phenomenon holds for long-run interest rate and stock transaction volume elasticities in that the rolling estimates vary a great deal before 1991 but become very stable after 1991. The relative large variation of money demand elasticity before 1991 can be explained by the change of financial system and regulation then. From these estimates, we can draw a conclusion that money demand in Taiwan is stable, at least after 1991. An important question then arises. What causes the marked difference between Goldfeld type of money demand discussed earlier and ARMAX model of money demand specified in (8)? A simple comparison reveals the differences in model specifications. First, ARMAX model contains an extra variable, stock transaction volume and conventional Goldfeld model does not. Second, ARMAX uses MA(1) model for the residuals while Goldfeld adopts AR(1) model for the residuals. Third, ARMAX has no constant term but Goldfeld model does.

To determine the real cause, we have estimated two additional models: Goldfeld model with stock transaction volume included but without constant, and ARMAX model with constant. The results of rolling estimates are reported in Figures 6 and 7. A quick examination would lead to the discovery that it is omission of the insignificant constant that stabilize the parameter estimates. Adding stock variable and replacing AR model for residuals with MA model for residuals does have some effect but not in a fundamental way. This seems a surprising answer, as it is the economist's tradition to include constant even when it is insignificant. Furthermore, constant term has been used as a safe guard against neglecting important variables. The role of constant within a dynamic model is an important issue and deserves further research. To briefly summarize, the ARMAX empirical results support the existence of stable money demand in Taiwan.

Next, let us turn to the empirical results of cointegration analysis. The unit root test results are put in Table 3, which clearly support the existence of unit root. It is well known that the form of deterministic terms of VAR will affect the limiting distribution of the cointegration rank test statistics. More specifically, there are five models depending upon deterministic terms,

- 1: $\mu = 0, \delta = 0$,
- 2: $\mu = \alpha\beta_0', \delta = 0$,
- 3: $\mu_0 \neq 0, \delta = 0$,
- 4: $\mu_0 \neq 0, \delta = \alpha\beta_1'$,
- 5: $\mu_0 \neq 0, \delta \neq 0$,

Let $M_{i,j}, i = 0, 1, \ldots, p, j = 1, 2, 3, 4, 5$ denote model i with j cointegration vectors. Then to obtain proper size, we should start testing the most restrictive $M_{0,1}$ using either *Trace* or L_{max} statistics. If H_0 is rejected, then we should proceed with the order, $M_{0,1}, M_{0,2}, M_{0,3}, M_{0,4}, M_{0,5}, M_{1,1}, \ldots, M_{n,5}$ until when H_0 is not rejected. Table (4) summarizes the rank test results.

The test results lead to $M(0, 3)$ that is, unrestricted constant term with no cointegration vector. However, $M(0, 3)$ is barely unrejected at 90% and λ_{max} statistics suggest one unit root. The univariate statistics for all four series under the assumption of one unit root are put in Table 5. The diagnostic checking statistics indicates the residuals behave like white noise. While the normality test is badly rejected, it is the typical case for the data in Taiwan. It is worth mentioning that removing stock transactions variable from the model would lead to clearly no cointegration results.

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

To further examine the stability of cointegration estimate, we perform the rolling estimation and report the results in Figures 8 to 10. From the figures, we find the estimate display instability at early 1993 but remain stable since then. This result seems to be consistent with the ARMAX modeling results.

4 Conclusion

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

Table 1: Estimation results of ARMAX model

Variable	Parameter	Coeff.	Std.	t-ratio
m_{t-1}	a_1	.8151	.0289	28.17
y	a_2	.1466	.0189	7.39
i	a_3	-.1568	.0199	-7.87
S	a_4	.0171	.0054	3.16
MA(1)	θ_1	-.3365	.1011	-3.33
D_1	d_1	.0235	.0064	3.64
D_2	d_2	-.0494	.0076	-6.49
D_3	d_3	-.0327	.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	.06	.18	.02	.04	.16	-.07	.08	-.08	.07	-.12	-.06	-.06	-.10
	1	-.30	.18	-.08	-.00	.15	.04	.01	-.01	-.02	-.13	.03	.00	-.01
	2	.00	-.03	-.04	-.01	.15	.03	.01	-.01	-.01	-.12	.06	-.00	-.01
	3	.02	.01	-.41	-.01	.15	-.05	-.01	-.02	-.01	-.11	.05	-.00	.01
	4	-.03	.01	-.42	-.05	.16	-.03	-.01	-.01	-.00	-.10	.00	-.04	-.04
	5	.42	.12	.17	.07	.29	-.06	.02	-.05	.05	-.00	.00	-.02	-.01
	6	.28	-.29	.17	-.04	.06	.23	.03	-.05	.07	-.02	-.01	-.02	-.01
Simplified extended ACF (5% Level)														
		Q												
		0	1	2	3	4	5	6	7	8	9	10	11	12
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.

Table 3: Unit root test results

Level	τ_μ	First Difference	τ_μ
M_1B	-1.160	ΔM_1B	$-3.889^{\dagger}$
GDP	-1.450	ΔGDP	$-4.447^{\dagger}$
r	-1.650	Δr	$-5.657^{\dagger}$
S	-1.031	ΔS	$-8.127^{\dagger}$

Table 4: Tests for cointegration rank and deterministic terms

eigen values						
r	n-r	Model 1	Model 2	Model 3	Model 4	Model 5
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						
r	n-r	Model 1	Model 2	Model 3	Model 4	Model 5
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 & Juselius 1995 appendix B, Tables B1-B5

Table 5: Diagnostic checking statistics

Univariate Statistics				
Eq.	Std.dev	Skewness	Kurtosis	Normality χ^2
1	0.0257	0.0377	3.2961	1.477
2	0.0433	2.4854	18.0335	41.440
3	0.0090	0.3248	3.3468	2.251
4	0.3637	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	r	s
Cointegration vectors: β'			
1.000	-0.295	0.538	-0.250
Loading Coef. α'			
0.054	-0.131	-0.017	1.657

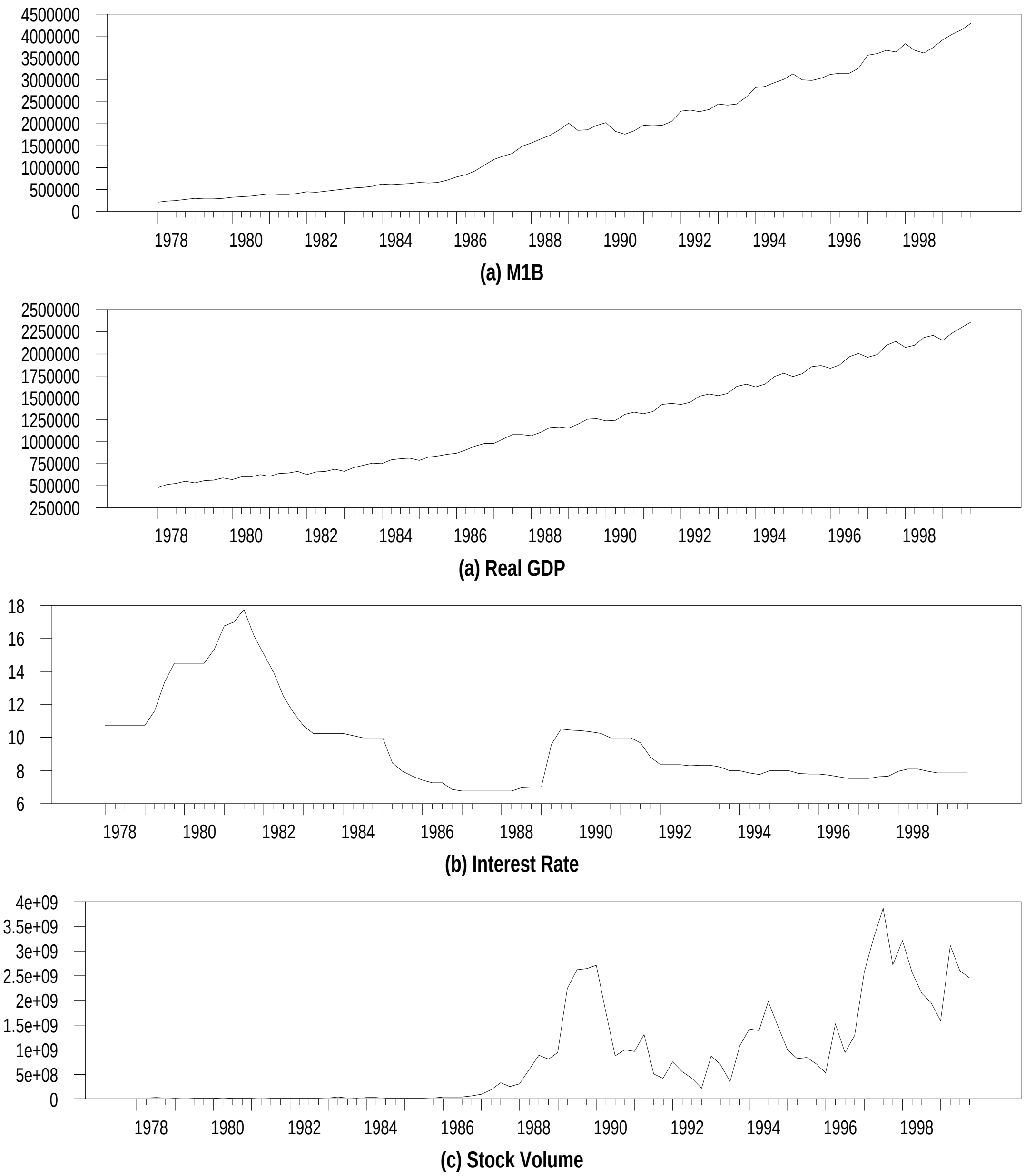

Figure 1: Time Series Plots for Real Variables in Logarithm

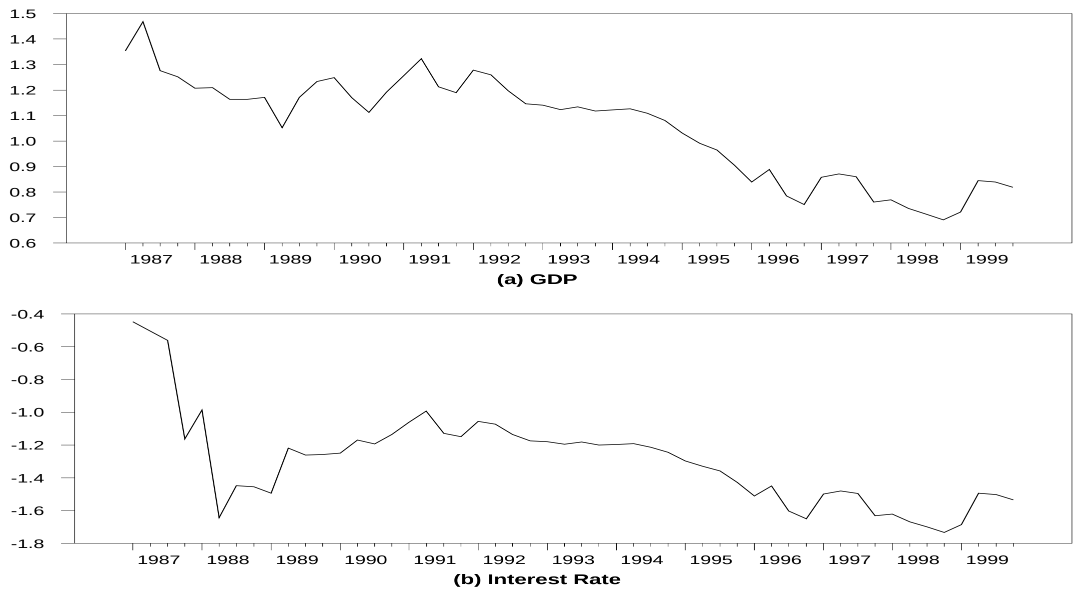

Figure 2: Long-run Elasticity of M1B: No Stock Volume, AR(1), Constant

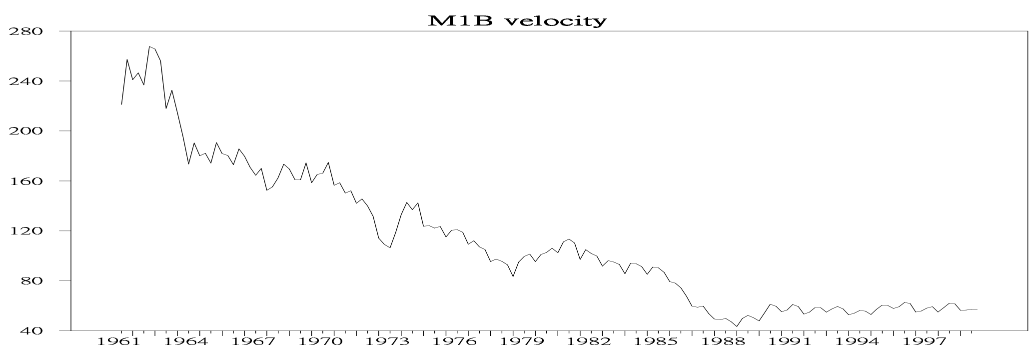

Figure 3: Velocity of M1B

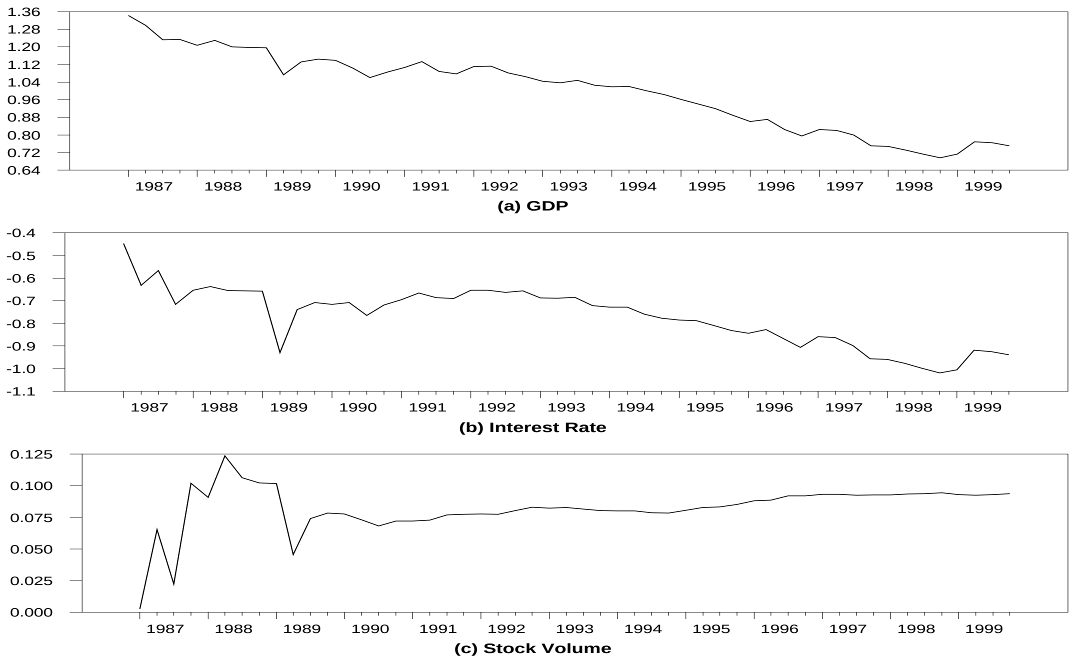

Figure 4: Long-run Elasticity of M1B: Stock Volume, AR(1), Constant

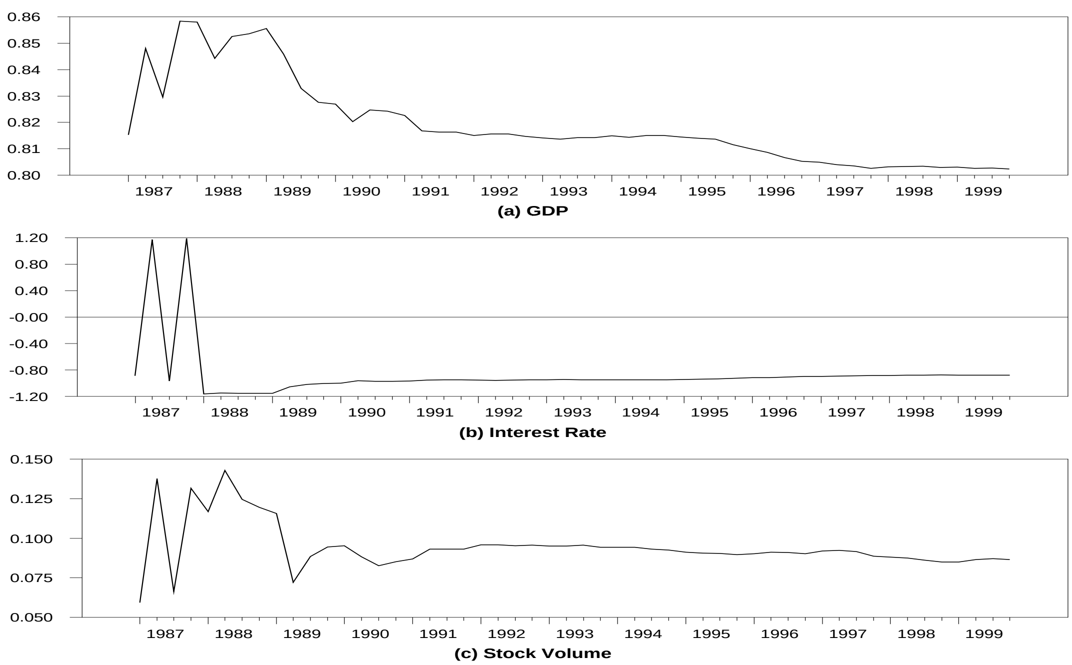

Figure 5: Long-run Elasticity of M1B: MA(1), Stock Volume, No Constant

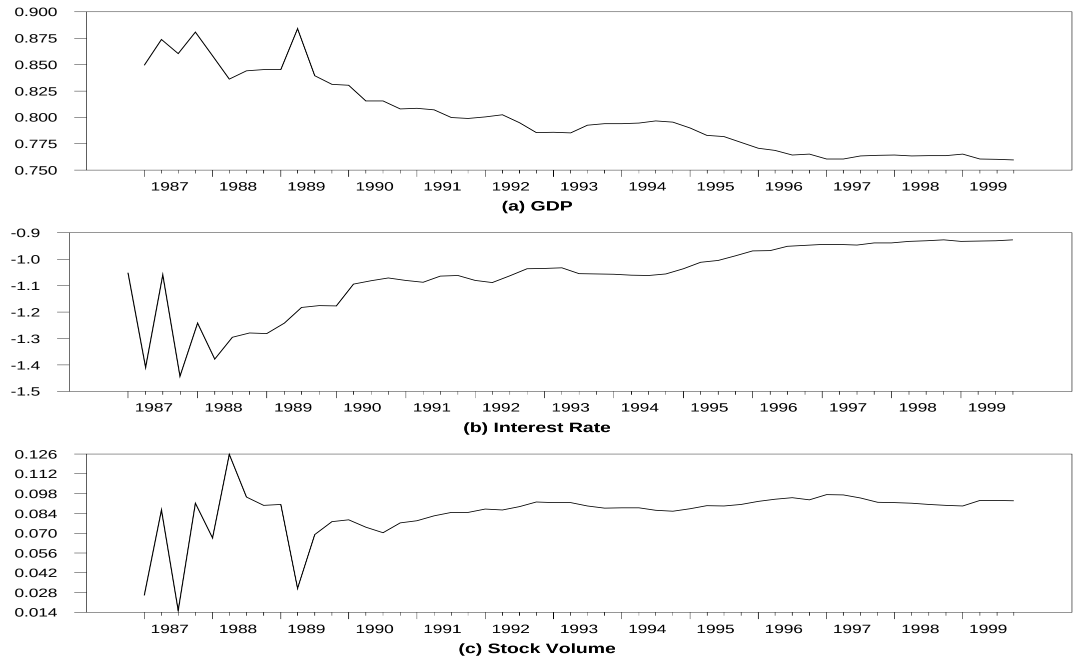

Figure 6: Long-run Elasticity of M1B: AR(1), Stock Volume, No Constant

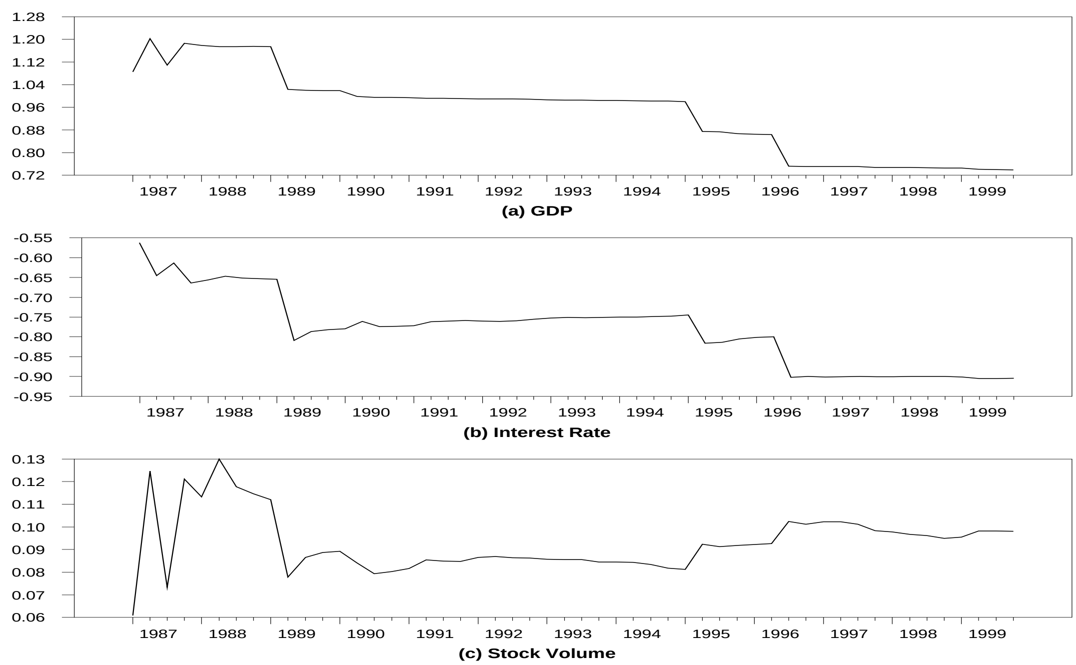

Figure 7: Long-run Elasticity of M1B: MA(1), Stock Volume, Constant

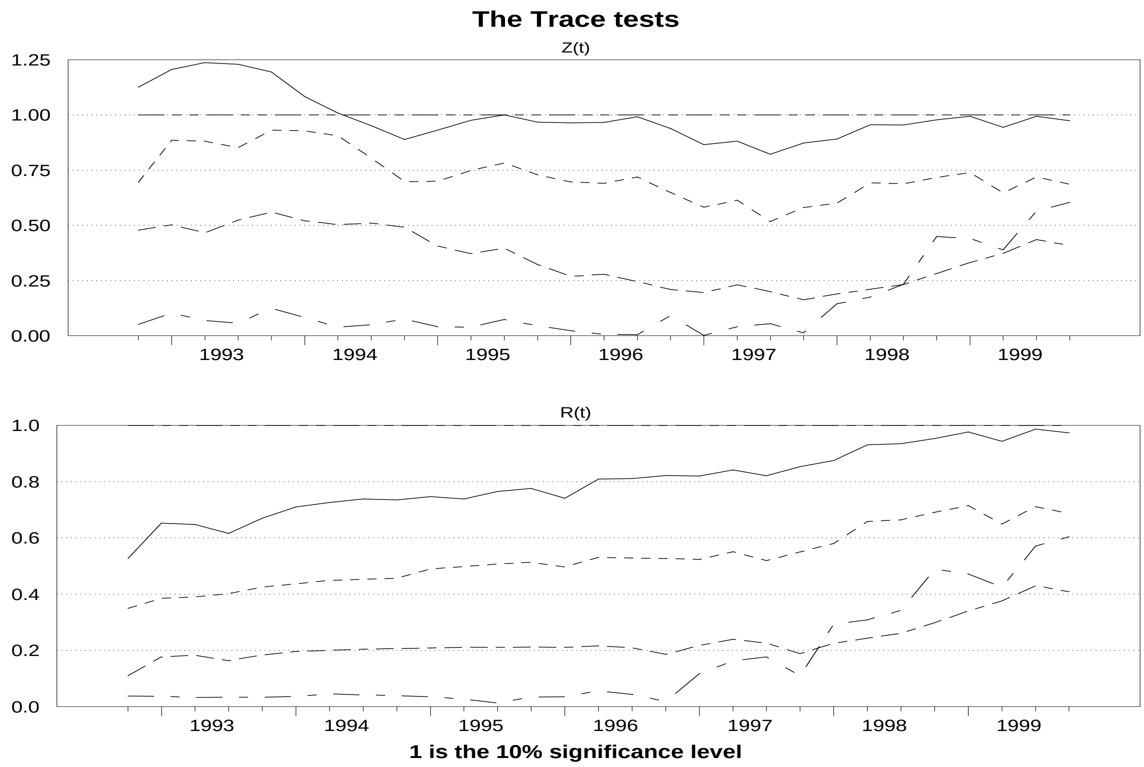

Figure 8: Rolling estimate of trace statistics

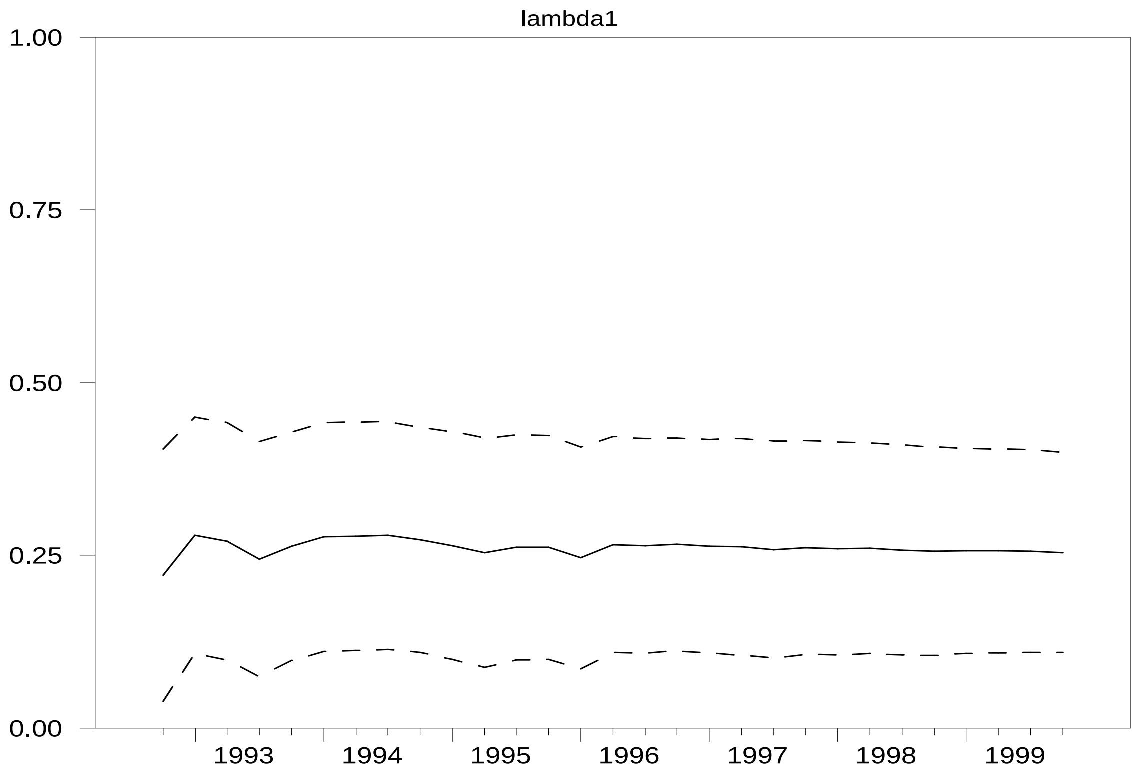

Figure 9: Rolling estimate of eigen values

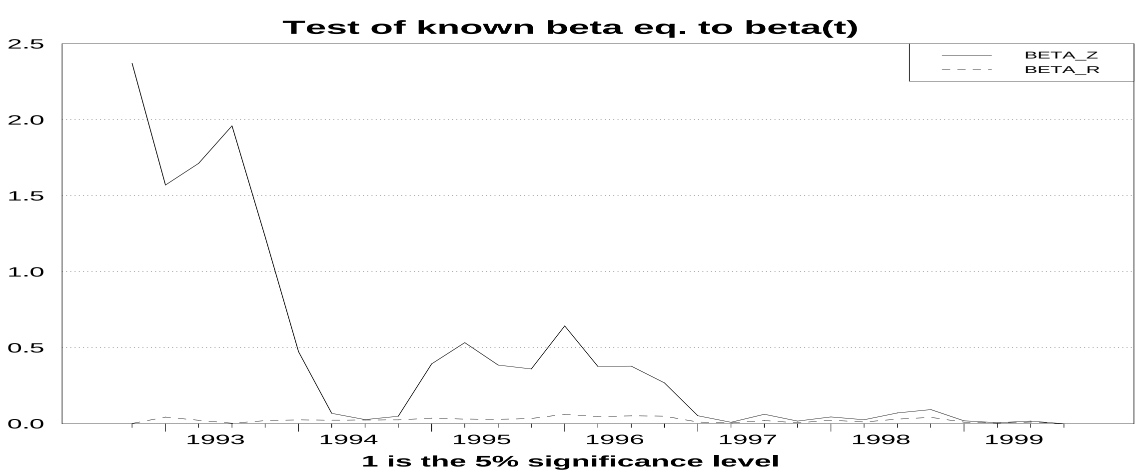

Figure 10: Rolling estimate of β estimate

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