Stability of money demand function revisited in China

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Abstract

As China's economic reforms have undergone significant structural changes after 1979, it seems difficult to formulate a stable money demand function over the period following 1979. Previous literature on the long-run relationship of money demand in China shows the existence of stable money demand. This paper revisits the stability of the China money demand function over the period after 1979. To employ the unit root tests and the cointegration tests with structural break, the empirical evidence demonstrates that economic and financial deregulation did affect the stability of demand for money in China over the period 1977-2002. Next, the estimated long-run income and interest elasticity is respectively 1.01 (1.11) and -0.14 (-0.08) using the real M1 (M2) equation. In addition, income and interest rate are found to be weakly exogenous. Overall, we do find structural breakpoints mainly in 1980 and 1993, and they look to match clearly with corresponding critical financial and economic incidents.

***JEL classification*:** E41, C22, C52

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1. Introduction

The stability of the money demand function has long been the central proposition of monetary economics. If money demand is not stable, then monetary policy has very little role to play. The success of monetary policy depends on where existing a steady-state relationship between money demand and its determinants (Hacker & Hatemj-J, 2005) Because of its importance, a steady stream of empirical research has been conducted worldwide over the past decades.[1] The extensive literature shows that the instability of the money demand function is a consequence of regime changes. This question has received increasing attention among the industrial countries. Recently applying modern econometric techniques of time series, cointegration and error-correction model, is the most frequent econometric approach in stability of the money demand function literature for developed counties. Examples include Muscatgelli & Spinelli (2000) for Italy, Karfakis & Sidiropoulos (2000) for Greece, Greene (2002) for U.S., Bahmani-Oskooee & Chomsisengphet (2002) for eleven developed countries, Narayasu (2003) for Japan, Hacker & Hatemj-J (2005) for Sweden. For it is reasonable to suspect that money demand might become unstable due to a number of financial liberalization measures since the 1980s, the interest on developing countries has heightened in recent years, including Sriram (2002) for Malaysia, Pradhan and Subramanian (2003) for India, Cheong (2003) for Korea, Gabriel, et al.(2003) for Portugal, Wu et al.(2005) for Taiwan, Bahmani- Oskooee & Rehman (2005) for six Asia developing countries, Bjornladn (2005) for Venezuela.

Over the past several decades a wealth of studies have focused on issues pertaining to the money demand function of China, but most have examined the relationship

[1] See Siriam (1999).

between money and income (Hafer and Kutan, 1993; Hasan, 1999; Xu, 1998; Yu and Tsui, 2000). To be sure, in other words, few have investigated the stability of the money demand function in China.[2] The purpose of this paper is to re-investigate whether economic and financial regime changes in China have broken down the stability of the money demand function by applying the unit root tests and the cointegration tests with structural break. Furthermore, we collect the most up-to-date data and span a period from 1977 to 2002, as these years have been characterized by critical changes in China's economy.

Our sample set covers the period after 1991, for which previous studies do not cover. During this period, further financial reforms in China, including a 16-point financial reform program in 1993, has been on-going. Our empirical investigation also updates the data after 1991 so as to discuss the stability of regime change in China during the 1990s. Unlike previous findings of Chen (1997) and Deng and Liu (1999), this paper shows strong contradictory evidence by following the methodology of Pradhan and Subramanian (2003) and Ramachandran (2004) to the China case. Inconsistent findings among previous studies for a stable money demand function in China may be attributable to the specification of those models, econometric methods, and the length of data span.

We apply a three-step testing procedure to inquire into the implication of the reform process on the stability of money demand. The empirical evidence is based on an elaborate methodology, by first identifying the full systems model of money demand and then reducing it to the single equation framework with a final testing for the structural break with unknown timing. Particularly, the latter issue is investigated by

[2] Chaisrisawatsuk et. al. (2004) indicate that the relative absence of empirical money demand studies for transition economies is due in part to the relative instability of these economies in the transition process itself as well as because of concerns over the reliability and frequency of time series data.

applying the methodology of Gregory and Hansen (1996) to test for cointegration between variables in the models allowing for the possibility of one break in the cointegration vector with unknown timing, which prevents an ad hoc selection of structural breaks. Our empirical study, which shows the possibility of regime shifts, seems to suggest a lack of stability in the demand for money in China, given the dataset from 1977 to 2002 and the model specifications. Hence, our results differ from previous studies by applying the unit root tests and cointegration allowing for structure change.

A study of China's money demand is of interest for three reasons as follows.[3] First, since market-oriented reforms were inaugurated in 1979, China has undergone significant systematic changes. The most influential adjustments of these included the establishment of a primitive market structure and the decentralization of decision-making powers from the central government to the local and grassroots levels. As a result, the economy has rapidly become commercialized, and money stock has already played a central role in facilitating all kinds of economic activities and market transactions.[4]

Second, China's gradual reforms have induced a dual-track banking controlling structure. In order to get better control over the economy, the banking system was reformed in order to separate banking functions and to create new policy instruments. After the establishment of the central bank system in 1983, the People's Bank of China (PBC) now controls the total volume of credit in the economy and works closely with the State Council in making important macroeconomic policy decisions.[5] It is a natural view that the banking system and financial sector will move to a system of monetary

[3] See Xu (1998), Yu and Tsui (2000) and Qin et al. (2005) for a detailed discussion.
[4] See Yu and Xie (1999), p.34.
[5] Beforehand, the reforms of macroeconomic policy were determined solely by the State Council in China. The central bank did not have independent monetary policies, prompting it to essentially support the implementation of the physical output targets contained in the central plan. See Blejer et al. (1991).

control by means of indirect market-based instruments.

In addition, financial reforms have spawned numerous new financial instruments, new financial services, and new institutions. These developments have promoted a substantial increase in the volume of bank deposits. One potential implication of these reforms is the increased importance of monetary aggregates in policy decisions. Due to its potent economic growth and expanding importance of trade and investment in China with countries all over the world, it is pivotal for the relative countries to comprehend PBC's conduct of its monetary policies and their effects on China's macroeconomic indicators, including general price level, production quantity of goods and services, and exchange rates.

In this regard, Chow (1987) applies the quantity theory of money to estimate a simple money demand function for China using annual data from 1952 through 1983. Portes and Santorum (1987) use real and nominal adjustment specifications and test for the homogeneity of money demand with respect to the price level and real income. Feltestein and Farhadian (1987) estimate a money demand function based on Cagan's (1956) model, and Blejer et al. (1991) apply an error correction model to estimate the demand for real balances using data for only the 1980s. Applying the cointegration method, Hafer and Kutan (1994) find a long-run equilibrium relationship between nominal money balances and other macroeconomic variables. Furthermore, Yu and Tsui (2000) construct a monetary services index to replace the simple-sum aggregates in estimating long-term money demand. However, all of them do not deal with the stability issue.[6] This paper fills the gap in the empirical literature on the stability of money demand function by studying the situation of China.

[6] They capture the effects of economic and financial reforms subjectively by including a dummy variable that takes a value of one for the 1979-1988 period and zero elsewhere.

The factor of owing to structural breaks is a common problem in macroeconomic series as they are usually affected by exogenous shocks or regime change in critical economic events. This paper reviews evidence from China on structural breaks in the long-run demand for broad and narrow money functions. As China's economic reforms have undergone significant structural changes after 1979, it would be difficult to formulate a stable money demand function over the period following 1979. Compared with the rich literature of quantitative studies on different countries, quantitative studies on the stability of money demand in China are still at an early stage.

Continuously, Huang (1994) employs a cointegration test to investigate the long-run relationship of money demand and constructs an error correction model to evaluate the dynamic adjustment process of money demand in China in the reform period (1979 to 1990). Huang (1994) explores the stability of money demand to use recursive regression method which does not required knowledge of the timing of possible breaks.[7] Chen (1997) implements the sup-F test statistic of Hansen (1992) to test the cointegration stability of money demand in China with unknown timing. The empirical analysis shows that no structural breaks are found in the money demand function during 1951 to 1991. The default of Chen (1997)'s analysis is that it just tests whether the structural change exists, but is unable to find the breakpoint timing. Furthermore, his study has no structural break of money demand in China, which is not associated with the acceleration of economic and financial reforms in China after 1979. Similarly, Deng and Liu (1999) combine artificial neural networks with the cointegration and error-correction models to a non-linear model, which indicate that the money demand function is stable.

[7] Traditional breakpoint tests include the Wald test, Likelihood ratio test, Lagrange multiplier test, and Chow test, etc.

The instability of the economic system may unfortunately in fact be reflected in the parameters of the estimated models that, when used for inference or forecasting, can induce misleading results. In contrast to previous studies, our study carefully analyzes the specification of the model and draws a conclusion by taking into account possible structural breaks. The paper is organized as follows. Section 2 describes the basic function of money demand and data used. The empirical results are presented in Section 3, and concluding remarks and policy implications close the article in Section 4.

2. Data and Specification

Previous literature on developing countries indicates that the models on narrow money work better when reflecting a weak banking system and low-financial sector development.[8] However, narrow money over time accommodates the new instruments created as a result of the evolving institutional and financial system structures. Following Pradhan and Subramanian (2003), this paper uses both definitions of money, narrow and broad, to model the demand for money in China.

The sample period is from 1977 to 2002. Annual data for narrow money (M1) (currency plus demand deposits held by households and enterprises),[9] broad money (M2) (M1 plus time and saving deposits held by households and enterprises), real national income (deflated by the consumer price index, 1990=100),[10] and the 1-year time deposit rate are obtained from *International Financial Statistics* (IFS) and *World*

[8] Pradhan and Subramanian (2003) and Ramachandran (2004) offer detailed discussions.
[9] Our justification for using annual data was based on the following. As Davidson and MacKinnon (1993) stated, to avoid any potential biases from using seasonally adjusted data (quarterly or monthly data) when conducting unit root tests, annual data should be used. Along the same lines, Hakkio and Rush (1991) found that when using monthly or quarterly data in a cointegration analysis, increasing the number of observations does not add any robustness to the results.
[10] Nominal income measures gross output from agriculture, industry, construction, transportation, and commerce.

Development Indicators (WDI). We use the 1-year time deposit rate as an opportunity cost of holding real balances.[11] The sample period is determined by the availability of consistent measures of the aggregate in question. All variables used are in natural logarithms except for the interest rate. To avoid possible distortions of the dynamic properties of the model, this paper uses seasonally-unadjusted data. Figure 1 presents graphs of the data series, which appear to be non-stationary and exhibit rather different patterns. While the real M2, the real M1, and the real national income series show steady growth over the last three decades, we observe a downturn during the last part of the 1980s. Moreover, the time deposits display a significant variation over time, which seem to be breaking around the 1990s.

Most previous works are based on the following general specification of the standard semi-logarithmic specification of the long-run money demand function:[12]

$$m_t = \phi_0 + \phi_1 y_t + \phi_2 r_t + \varepsilon_t , \qquad (1)$$

where *m* represents logged real money balances created by taking a monetary aggregate deflated by the CPI; *y* is logged real income measured via real national income; *r* is an opportunity cost proxied via the 1-year time deposit rate, and ε is a residual term. The coefficient ϕ_i, $i = 1,2$, refers to the elasticities of income and interest rate. The coefficients ϕ_1 and ϕ_2 are expected to enter with positive and negative signs, respectively. If a long-run relationship exists between m, *y*, and *r*, then the finding of cointegration is the statistical equivalent of the long-run concept in economics. Friedman (1956) suggests that research on the money demand function assumes that

[11] Poole (1988) argues that long-term interest rate specifications are more robust than those employing a short-term interest rate in money demand function.

[12] See Ericsson et al. (1998), Carlson, et al. (2000), and Cargill and Parker (2004).

there exists an underlying stationary long-run equilibrium relationship between real money balances, real income, and the opportunity cost of holding real money balances. According to the Cambridge and the Keynesian approaches, the relationship between real money demand and the level of real income is direct, and the relationship between real money demand and the rate of interest are inverse. One may think as such that there is an opportunity cost of holding money (Choudhry, 1999).

3. Empirical Investigation

3.1 Unit root test

We first apply Dickey and Fuller's (1981; hereafter ADF) three-model tests, and we follow the determining rule by Doldado et al. (1990, hereafter DJS) to establish the appropriate model. The determining rule by DJS tests for the significance of the trend coefficient in the third model first, followed by testing for the significance of the drift coefficient in the second model. If both outcomes result in being insignificant, then the first model is selected. Moreover, since the estimation might be biased if the lag length is pre-designated without any rigorous determination, this paper adopts the newly-developed Modified Akaike's information criterion (MAIC), as suggested by Ng and Perron (2001), to select the optimal number of lags based on the "principle of parsimony."

The classic ADF tests may nevertheless be suspect, not taking into account that the structural breaks could lead to a wrong decision when the null hypothesis is not rejected. A structural break essentially corresponds to an intermittent shock with a permanent effect on the series (Hendry, 1996). The opposite can also happen if the break occurs at the beginning of the sample (Leybourne et al., 1998). In order to take into account this possible shift in regime in the unit root tests, Zivot and Andrews (1992, hereafter ZA) develop a new category of tests that allow for a structural break. The three models of the ZA tests are expressed as the following equations:[13]

$$\text{Model A: } \Delta Y_t = \mu_1^A + \gamma_1^A t + \mu_2^A DU_t(\lambda) + \alpha^A Y_{t-1} + \sum_{j=1}^{k-1} \beta_j \Delta Y_{t-j} + \varepsilon_t \quad (2)$$

$$\text{Model B: } \Delta Y_t = \mu_1^B + \gamma_1^B t + \gamma_2^B DT_t^*(\lambda) + \alpha^B Y_{t-1} + \sum_{j=1}^{k-1} \beta_j \Delta Y_{t-j} + \varepsilon_t \quad (3)$$

$$\text{Model C: } \Delta Y_t = \mu_1^C + \gamma_1^C t + \mu_2^C DU_t(\lambda) + \gamma_2^C DT_t^*(\lambda) + \alpha^C Y_{t-1} + \sum_{j=1}^{k-1} \beta_j \Delta Y_{t-j} + \varepsilon_t \, , \quad (4)$$

where $DU_t(\lambda)$ is a dummy variable which is equal to 1 and $DT_t^*(\lambda) = t - T\lambda$, 0 otherwise. Furthermore, $\lambda = T_B / T$, and T_B represents a possible break point, where T is the sample size. The breakpoint is searched for over the range of the sample (0.15T, 0.85T), and it can be estimated endogenously. Model A allows for a change in the level of the series, Model B allows for a change in the slope of the trend of a series, while Model C combines both changes in the level and the slope of the trend. Since the appropriate model and the optimal lag lengths are crucial in interpreting the results of the tests, we adopt the findings from the ADF tests to select the model and the lag lengths for the ZA tests.

The results of the ZA tests are presented in Table 2 and shown in Figure 2. All series carry a unit root in the level and reject the null of "non-stationarity" in the first

[13] For the low power of finite-sample problems, ZA (1992) proposes some explanations in Section 5. In ZA, the authors use the Nelson-Plosser dataset to carry on endogenous break unit root tests. We find that the observations (T) of some series (real GNP, nominal GNP, real per capita GNP) have only 62 (same as Table 3 of Perron, 1997), but ZA also adopt asymptotic critical values equally (see p.259, Table 6), in which critical values for the limiting distributions in the theorem are obtained by simulation methods. We follow ZA's approach and adopt asymptotic critical values provided by ZA. Follow-up related empirical papers, like Cakan and Özmen (2002) and Kollias et al. (2004), also apply the asymptotic critical values as a guideline for their empirical results when they use the endogenous break unit root tests.

difference. This insures the I(1) type series for all series considered. The ZA test results indicate that the breakpoint occurs in 1988 for real money balances, in 1992 for real income, and in 1993 for the interest rate.

Besides, we also employ the Bai and Perron (1988, 2003, hereafter BP) methodology to test for multiple structural breaks. Table 3 reports the number of breaks selected by the BP sequential tests. The BP test all suggest, for all variables, at least three breaks. The UDmax and WDmax rejected the null of no breaks. In addition, Sup $F_T(l+1|l)$ statistics and the BP sequential procedure indicate that the SupF(3|2) tests were rejected for the null of 2 breaks against the 3 alternative breaks, and failed to reject the SupF(4|3) breaks test. Table 4 is the BP test results for the location of structural breaks. We find that the breakpoints occur in 1983, 1990, and 1996 for m2, in 1983, 1990, and 1997 for m1, 1983, 1991, and 1996 for real income, and in 1979, 1984, and 1997 for the interest rate.

Finally, according to the empirical results of ZA and BP tests for structural breakpoints, we can summarize the following dates of structural breakpoints and find critical economic and financial incidents for China that can match with the structural breaks of these series.

1. The first date of structural break is 1979. What caused the breakpoint in 1979? Since market-oriented reforms were inaugurated in 1979, the structure and mechanism of the economic and financial system in China have significant changes.

2. The second date is near 1983. China's gradual reforms have induced a dual-track banking controlling structure. In order to get better control over the economy, the banking system was reformed in order to separate banking functions and to create new policy instruments. After the establishment of the central bank system in 1983, the PBC can control the total volume of credit in the economy and make important

macroeconomic policy decisions. For covering the loans of the policy-oriented items, the “loan for grant” reform was implemented in 1983 and 1984.[14]

3. The third date is near 1988. China’s inflation continued at accelerating rates well above 18% in 1987-88. In order to tighten credit to prevent overheating in the Chinese economy, more restrictive monetary and credit conditions were implemented in 1988.

4. The fourth date is near 1992-1993, Ever since the reform of 1979, the most dramatic change in the economy took place in 1992, when the Party officially recognized that a market system was not incompatible with the ideals of socialism and accordingly declared the idea of establishing a “socialist market economy.”[15] In the wake of this decision, the Chinese leadership outlined an extensive reform strategy which explicitly identified financial reform as a key element to effect macroeconomic management. Besides, the government instituted a 16-pint financial reform program in July 1993, for solving the increasing inflation problem which was threatening China’s spectacular growth. This program addressed macroeconomic imbalances stemming from uncoordinated lending, growth in the money supply, and subsequent inflation. The PBC was authorized to conduct domestic monetary policy through open market operations and manipulation of the bank loan rates and the banking system’s reserves.[16]

5. The fifth date is near 1996-1997. What caused the breakpoint? In 1995, the National People’s Congress officially a passed ‘The People’s Bank of China Law’ and ‘The Commerical Bank Law of China’. The objective of these reforms is to promote bank commercialization. Besides, the central bank has lowered the interest rates on the central bank loans six times to stabilize economic development since the early 1996.

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[14] See Huang (1998), p.6.
[15] See Mchran and Quintyn (1996), p.25.
[16] See DaCosta and Foo (2002), p.4.

3.2 Testing the long-run relationship in money demand

Given unit roots, the issue arises as to whether there exists a long-run equilibrium relationship of the money demand function. Johansen (1988) proposes two test statistics for examining the number of cointegrating vectors (namely, the Trace and the L-max statistics). Cointegration implies that the transitory components of a series can be given a dynamic specification by means of the error correction models that are inclined towards a stationary long-run money demand function. Therefore, it is clear that adopting the best strategy to find the long-run relationship between the demand for money, income and the interest rate constitutes the basis for the success of monetary policies. This paper adopts the Schwartz Bayesian information criterion (SBC) to determine the optimal number of lags based on the "principle of parsimony". The SBC suggests one lag for the VAR model.[17] We also correct our statistics for a small sample bias as suggested by Cheung and Lai (1993). The results of Table 5 present the Trace statistics and the L-max statistics, both of which suggest that there exists one cointegrating vector, implying that there exists a long-run relationship. We normalize the cointegrting vector with respect to the real money balances, and then the cointegrating relation is:

$$m2 = 1.110 * y - 0.082 * r \tag{5}$$

$$m1 = 1.013 * y - 0.141 * r\,. \tag{6}$$

From equations (5) and (6), the long-run income elasticity for real money balance is close to one as suggested by the quantity theory of money. The coefficients estimated

[17] Bessler and Binkley (1982) and Geweke and Meese (1980) show that the SBC appears to be superior to other lag length selection methods.

are 1.11 and 1.01, respectively. The long-run demand for real money balance is negatively affected by the own rate of return for money, and the interest semi-elasticity is approximately -0.08 to -0.14. This suggests that higher interest rates in China reduce the demand for broad and narrow money. Moreover, in absolute value, the interest rate has a larger effect on the demand for real M1 than for real M2.

In order to examine the long-run causal relationship, we test for weak exogeneity among the cointegrating relationships (Johansen and Juselius, 1992). Hall and Alistair (1994) and Arestis et al. (2001) interpret weak exogeneity in a cointegrated system as a notion of long-run causality. The null hypothesis is the existence of weak exogeneity. Testing for weak exogeneity in the system as a whole requires a test of the hypothesis that $H_0 : \alpha_{ij} = 0$ for $j = 1,...,r,$ where α is the speed of adjustment parameters (the loading coefficients of the equilibrium error in the dynamic equation for real money) and i contains only zeroes. This test, using a likelihood ratio test, involves the restricted and unrestricted models to ascertain whether the restrictions are valid. If the results of the test are not able to reject the zero null hypothesis, this indicates that the variable is a weak exogenous variable. Therefore, in our empirical model, this variable is the “cause” of the other exogenous variables that are not weak. On the contrary, if the null hypothesis is rejected, this indicates the existence of bi-directional causality between this variable and other exogenous variables that are not weak.

Table 6 shows that weak exogeneity is rejected for real money balances at 5%, which indicate that unidirectional causalities run from real income and interest rates to real money balances in the long run. Based on the result of the weak exogeneity test, a short-run model can be designed with a system of one equation - both narrow and broad money - by considering income and the interest rate as weakly exogenous. The dynamic adjustment to disequilibria occurs via changes in real money. Therefore, we focus our attention only on the single equation specification for money.

3.3 Structural break test

The estimation period for this study covers the somewhat violent time period of financial and economic innovation in China. Consequently, it is important to check the cointegration relationship for structural breaks.[18] We follow the methodology of Gregory and Hansen (1996, hereafter GH) and test for cointegration between variables in the models with regime shifts. The GH test is based on the notion of regime change and is a generalization of the usual residual-based cointegration test.

GH consider three alternative models - a level shift (model C), a level shift with trend (model C/T), and a regime shift that allows the slope vector to shift as well (model C/S). The three models are expressed as the following equations:

model C： $$y_t = \mu_1 + \mu_2 D_t(\tau) + \beta_1 X_t + e_t \qquad (7)$$

model C/T： $$y_t = \mu_1 + \mu_2 D_t(\tau) + \gamma t + \beta_1 X_t + e_t \qquad (8)$$

model C/S： $$y_t = \mu_1 + \mu_2 D_t(\tau) + \beta_1 X_t + \beta_2 X_t D_t(\tau) + e_t \qquad (9)$$

and

$$D_t = \begin{cases} 0, & if \quad t \le T\tau \\ 1, & if \quad t > T\tau \end{cases}$$

[18] Because all of the variables used in this study have unit-root characteristics, using nonlinear models such as the threshold autoregressive model or the smooth transition autoregressive model to deal with the problem of structural change among the variables is also not appropriate. This is because those models' starting points require that all variables conform to the stationarity premise. For more detail, refer to Tersavirta (1994).

Here, $\tau = T_B / T$, and T_B represents a possible breakpoint. GH have developed versions of the cointegration ADF tests of Engle and Granger (1987), as well as the Z_t and Z_α tests of Phillips-Quliaris (1990), whereby all them are modified according to the alternative considered. Taking into account that the date of the change is unknown, they compute the values of $ADF^* = \inf_{\lambda \in J} ADF$, $Z_t^* = \inf_{\lambda \in J} Z_t$, and $Z_\alpha^* = \inf_{\lambda \in J} Z_\alpha$. This model is estimated recursively, allowing the breakpoint τ to vary such that $|0.15T \le \tau \le 0.85T|$.

Table 7 presents the results of the GH test, showing clear evidence of not finding cointegration even when we allow for a structural break in the M1 money demand function. Moreover, the M2 money demand function reveals a structural break in ADF^* . The test does suggest that a structural change in the cointegration vector is important and needs to be taken care of in the specification of money demand function. Hence, the function specification of money demand, enveloping the changing economic and financial incidents, does raise some important questions on the long-run relationship between these series. Furthermore, according to the ADF statistic criterion in the broad money demand function, the structural break years estimated on the basis of the three models are mainly in 1980 and 1993.

The structural break year of 1980 is caused by the reforms of 1979. Prior to 1979, the role of monetary policy was accommodating and was to support the central plan, while maintaining price stability. Chinese monetary policy was to have no allocative function. The reforms undertaken since 1979 have tended to increase the role of monetary policy and have promoted a substantial increase in the volume of bank deposits. The regime changes from 1979 lasted for one period, and their influence on the money demand function had a time lag until 1980.

From the above description, 1993's structure change timing instituted a 16-point financial reform program. The PBC was authorized to conduct domestic monetary policy through open market operations and by manipulation of the bank loan rates and the banking system's reserves, thus functioning more like a central bank. This program made a structural break of the banking and financial systems in China. The empirical investigation is consistent with the research of DaCosta and Foo (2002), who divide China's financial reforms into two distinct periods: pre-1993 and post-1993.

These tests do suggest that structural change in the cointegration vector is important and needs to be taken care of in the specification of China's money demand function. The finding is reassuring since the endogenous estimation procedure produces structure breaks that correspond to recognizable financial and economic events. It implies that, within the context of money demand, households and the government may respond differently when the economy is in a different regime. Hence, the specification of the money demand function, enveloping the changing economic events and financial deregulations, does raise some important questions on the long-run relationship between money, income, and the interest rate.

4. Conclusions and Policy Implications

This paper re-investigates whether economic and financial reforms in China have made the demand function for money unstable for the period 1977-2002. Previous literature on the long-run relationship of China's money demand shows the existence of stable money demand. This paper has revisited the stability of the China money demand function over the period following the reforms of 1979.

The empirical evidence is based on an elaborate methodology, by first identifying the full systems model of money demand, then reducing it to the single equation framework, and then finally testing for the structural break with unknown timing. This test, which allows for the possibility of regime shifts, seems to suggest a lack of stability in the demand for money, given the dataset from 1977 to 2002 and the model specifications. This evidence, which is in contrast with the finding of Chen (1997) and Deng and Liu (1999), indicates that the stability of the money demand has been broken down by economic and financial reforms. Inconsistent findings among previous studies are found to be attributable to the specifications of the model, econometric methods, and the length of data span. Such as finding is reassuring since the endogenous estimation procedure produces structure breaks that correspond to recognizable financial and economic events.

The paper also shows the six following empirical results:

First, according the empirical results of ZA and BP tests, we did find several structural breakpoints and critical economic and financial incidents for matching with these breakpoints. Second, the long-run income elasticity for real money balance is close to one as suggested by the quantity theory of money. Third, the long-run demand for real money balance is negatively affected by the own rate of return for money. In absolute value, the interest rate has a larger effect on the demand for real M1 than for real M2. Fourth, the dynamic adjustment to disequilibria occurs via changes in real money balance. Fifth, income and interest rate are found to be weakly exogenous. Therefore, the unidirectional causalities run from real income and the interest rate to real money balance in the long run. Finally, we find that the structural break years estimated are mainly in 1980 and 1993, and they could match clearly with corresponding critical financial and economic incidents. Thus, while analyzing the function of money demand of China, one should surely include a structure change into the question.

Of particular interest, noteworthy is that the income elasticity of money demand for M2 is larger than that for M1, but, conversely, the interest rate elasticity of money demand for M1 is larger than that for M2. In the course of history, the sensitivity of desired money- holding in China has changed depending on different determination variables. Regardless of the income elasticity of M2 or M1, both are close to one as suggested by the quantity theory of money. Based on the above, some interesting, specific policy implications cannot go unnoticed. Supposing that the target of monetary policy is M2, then any attempt on the part of the central bank to interrupt the money supply of M2 should not be via the interest rate; on the contrary, it should be through real income in China. When determining monetary policy, since monetary income elasticity is evidently higher than interest rate elasticity, it should be more efficient on the part of the central bank to interrupt the quantity of money through real income rather than by changing the interest rate. Once again, raising interest rates in China tends to reduce the demand for money, particularly for M1.

REFERENCES

Arestis, P., Demetriades, P., and Luintel, K. (2001). Financial development and economic growth: the role of stock markets. *Journal of Money, Credit, and Banking* 33, 16-41.

Bahmani- Oskooee, M. and Rehman, H. (2005). Stability of the money demand function in Asian developing countries. *Applied Economics*, 37, 773-792.

Bahmani-Oskooee, M. and Chomsisengphet S. (2002). Stability of M2 money demand function in industrial countries. *Applied Economics,* 34, 2075-2083.

Bai, J., Perron, P., 1998. Estimating and testing linear models with multiple structural changes. *Econometrica* 66, 47–78.

Bai, J., Perron, P., 2003. Computation and analysis of multiple structural change models. *Journal of Applied Econometrics* 18, 1 –22.

Bessler, D. A. and Binkley, J. K. (1982). On the selection of the order of an autoregression: Some Mote Carlo results. *American Statistical Association 1982 Proceedings of the Business and Economic Statistics*, 340-342.

Blejer, M., Burton, D., Dunaway, S. and Szapary, G. (1991). China: Economic Reform and Macroeconomic Management. *IMF Occasional Paper, 76*.

Bjornland, H. C. (2005). A stable demand for money despite financial crisis: the case of Venezuela. *Applied Economics*, 37, 375-385.

Cagan, P. (1956). The monetary dynamics of hyperinflation. *Studies in the Quantity Theory of Money*, edited by Milton Friedman. Chicago: University of Chicago Press.

Cakan, E., and Özmen, E. (2002). Policy regime change and structural break in the velocity of money: the Turkish evidence. *Applied Economics Letters, 9,* 759-62.

Cargill, T. F. and Parker, E. (2004). Price deflation, money demand, and monetary policy discontinuity: a comparative view of Japan, China, and the United States. *North American Journal of Economics and Finance 15*, 125-147.

Carlson, J. B., Hoffman, D. L., Keen, B. D. and Rosche, R. H. (2000). Results of a study of the stability of cointegrating relations comprised of broad monetary aggregates. *Journal of Monetary Economics 46*, 345-383.

Chaisrisawatsuk, S., Sharma, S. C. and Chowdhury, A. R. (2004) Money demand stability under currency substitution: some recent evidence. *Applied Financial Economics*, 14, 19-27.

Chen, B. (1997). Long-Run Money Demand and Inflation in China. *Journal of Macroeconomics 19*, 609-617.

Cheong, C. C. (2003). Regime changes and econometric modeling of the demand for money in Korea. *Economic Modelling 20*, 437-53.

Cheung, Y., and Lai, K. (1993). Finite-sample sizes of Johansen's likelihood ration tests for cointegration. *Oxford Bulletin of Economics and Statistics* 55, 313-328.

Choudhry, T. (1999). Does interest rate volatility affect the US M1 demand function? evidence from cointegration. *The Manchester School* 67 (6), 621-648.

Chow, G. C. (1987). Money and Price Level Determination in China. *Journal of Comparative Economics* 11, 319–333.

DaCosta, M. M. and Foo, J. P. (2002). China's financial system: Two decades of gradual reforms. *Managerial Finance 28*, 3-18.

Davidson, R. and MacKinnon, J.G. (1993). *Estimation and Inference in Econometrics*. Oxford, New York, Toronto and Melbourne: Oxford University Press.

Deng, S. and Liu, B. (1999). Modelling and forecasting the money demand in China: cointegration and nonlinear analysis. *Annals of Operations Research* 87, 177-189.

Dickey, D. A., and Fuller, W. A. (1981). Likelihood ratio statistics for autoregressive time series with a unit root. *Econometrica* 49: 1057–1072.

Doldado, J., Jenkinson, T., and Sosvilla-Rivero, S. (1990). Cointegration and unit roots. *Journal of Economic Surveys* 4, 249-273.

Engle, R. F., and Granger, C. W. (1987). Cointegration and error correction: Representation, estimation and testing. *Econometrica* 55, 251–276.

Ericsson, N. R., Hendry, D. F. and Prestwich, K. M. (1998). The demand for broad money in the United Kingdom, 1878-1993. *Scandinavian Journal of Economics* 100(10), 289-324.

Feltenstein, A., and Farhadian, Z. (1987). Fiscal policy, monetary targets, and the price level in a centrally planned economy: an application to the case of China. *Journal of Money, Credit, and Banking* 19, 137–156.

Friedman M. (1956). The quantity theory of money: a restatement. In: Friedman, M. (Ed.). Studies in the Quantity Theory of Money. University of Chicago Press, Chicago.

Garbriel, V. J. C. R. de A., Lopes, A. C. B. da S. and Nunes, L. C. (2003). Instability in contegration regressions: a brief review with an application to money demand in Portugal. *Applied Economics*, 35, 893-900.

Geweke, J. and Meese, R. (1981). Estimating regression models of finite but unknown order. *International Economic Review*, 22, 55-70.

Greene, C. A. (2002). Was money demand in the USA unstable before 1982? An application of the sup-F stability test. *Journal of Economics and Business 54,* 465-481.

Gregory, A. W., and Hansen, B. E. (1996). Residual-based tests for co-integration in models with regime shifts. *Journal of Econometrics 70*, 99–126.

Hacker, H. C. and Hatemj-J, A. (2005). The effect of regime shifts on the long-run relationships for Swedish money demand. *Applied Economics,* 37 (15), 1731-1736.

Hafer, R.W.and Kutan, A. M. (1993), Price deflation, money demand, and further: evidence on money, output, and prices in China, *Journal of Comparative Economics* 17, 701–709.

Hafer, R. W. and Kutan, A. M. (1994). Economic reforms and long-run money demand in China: implications for monetary policy. *Southern Economic Journal* 60(4), 936-945.

Hakkio, C.S. and Rush, M. (1991). Cointegration: how short is the long run? *Journal of International Money and Finance 10 (4)* 571-81.

Hall, S. G., and Alistair, M. (1994). The relevance of p-star analysis to UK monetary policy, *Economic Journal* 104, 597-604.

Hansen, B. E. (1992). Tests for parameter instability in regressions with I(1) processes. *Journal of Business and Economic Statistics*, 321–35.

Hasan, M. S. (1999), Monetary growth and inflation in China: A reexamination. *Journal of Comparative Economics*, 27, 669–685.

Hendry, D. F. (1996). A theory of co-breaking. Paper presented at the time series meeting organized by NBER-NSF.

Huang, G. (1994). Money demand in China in the reform period: An error correction model. *Applied Economics* 26, 713–719.

Huang, Y. (1998). Challenges for China's Financial Reform. China Update 1998 Conference papers, NCDS Asia pacific Press, Canberra,

Johansen, S. (1988). Statistical analysis of cointegration vectors. *Journal of Economic Dynamics and Control 12*, 231-254.

Johansen, S., and Juselius, K. (1992). Testing structural hypotheses in a multivariate cointegration analysis of the PPP and the UIP for UK. *Journal of Econometrics 53,* 211-244.

Karfakis, C. and Sidiropoulos, M.(2000). On the stability of the long-run money demand in Greece, *Applied Economics Letters*, 7, 83-86.

Kollias, C., Naxakis, C. and Zarangas, L. (2004). Defence spending and growth in Cyprus: a causal analysis. *Defence and Peace Economics, 15 (3)*, 299–307.

Leybourne, S., Mills, T., Newbold, P. (1998). Spurious rejections by Dickey-Fuller tests in the presence of a break under the null. *Journal of Econometrics* 53, 211-244.

Mchran, H. and Quintyn, M. (1996). Modernizing China's banking and financial sectors. *The Journal of Social, Political, and Economic Studies 21*, 25-32.

Muscatelli, V. A. and Spinelli, F. (2000). The long-run stability of the demand for

money: Italy 1861-1996. *Journal of Monetary Economics 45*, 717-739.
Nagayasu, J. (2003). A re-examination of the Japanese money demand function and structural shifts. *Journal of Policy Modeling 25*, 359-375.
Ng, S. and Perron, P. (2001). Lag length selection and the construction of unit root tests with good size and power. *Econometrica* 69(6), 1519-1554.
Perron, P. (1997). Further evidence on breaking trend functions in macroeconomic variables. Journal of Econometrics, 80, 355–385.
Phillips, P. C. B., Quliaris, S. (1990). Asymptotic properties of residual-based tests for coitegration. *Econometrica 58*, 165–193.
Poole, W. (1988). Monetary policy lessons of inflation and disinflation. *Journal of Economic Perspectives, 2,* 73-100.
Portes, R. and Santorum, A. (1987). Money and the consumption goods market in China. *Journal of Comparative Economics, Sep*., 154-63.
Pradhan, B. K. and Subramanian, A. (2003). On the stability of demand for money in a developing economy: some empirical issues. *Journal of Development Economics 72*, 335-351.
Qin, D., Quising, P., He, X. and Liu, S (2005). Modeling monetary transmission and policy in China. *Journal of Policy Modeling 27*, 157-175.
Ramachandran, M. (2004). Do broad money, output, and prices stand for a stable relationship in India? *Journal of Policy Modeling*, 26, 983-1001.
Sriram, S. S. (1999). Survey of literature on demand for money: theoretical and empirical work with special reference to error-correction models. Washington: International Monetary Fund, *IMF Working Paper No. WP/99/64.*
Sriram, S. S. (2002). Determinants and stability of demand for M2 in Malaysia. *Journal of Asian Economics* 13, 337-56.
Swanson, N. R. (1998). Money and output viewed through a rolling window. *Journal of Monetary Economics 41*, 455–473.
Terasvirta, T. (1994). Specification, estimation, and evaluation of smooth transition autoregressive models. *Journal of the American Statistical Association, 89,* 208-218.
WDI (2004). World Development Indictors, World Bank: Washington D. C.
Wu, C. S., Lin, J. L., Tiao, G. C. and Cho, D. D. (2005). Is money demand in Taiwan Stable? *Economic Modelling 22,* 327-346.
Xu, Y. (1998). Money demand in China: A disaggregate approach. *Journal of Comparative Economics 26*, 544-564.
Yu, Q. and Xie, P. (1999). Money aggregates management: problems and prospects in China's economic transition. *Contemporary Economic Policy 17*, 33-43.

Yu, Q. and Tsui, A. K. (2000). Monetary services and money demand in China. *China Economic Review* 11, 134-148.

Zivot, E., and Andrews, D. W. (1992). Further evidence on the great crash, the oil-price shock, and the unit-root hypothesis. *Journal of Business and Economic Statistics* 10, 251–270.

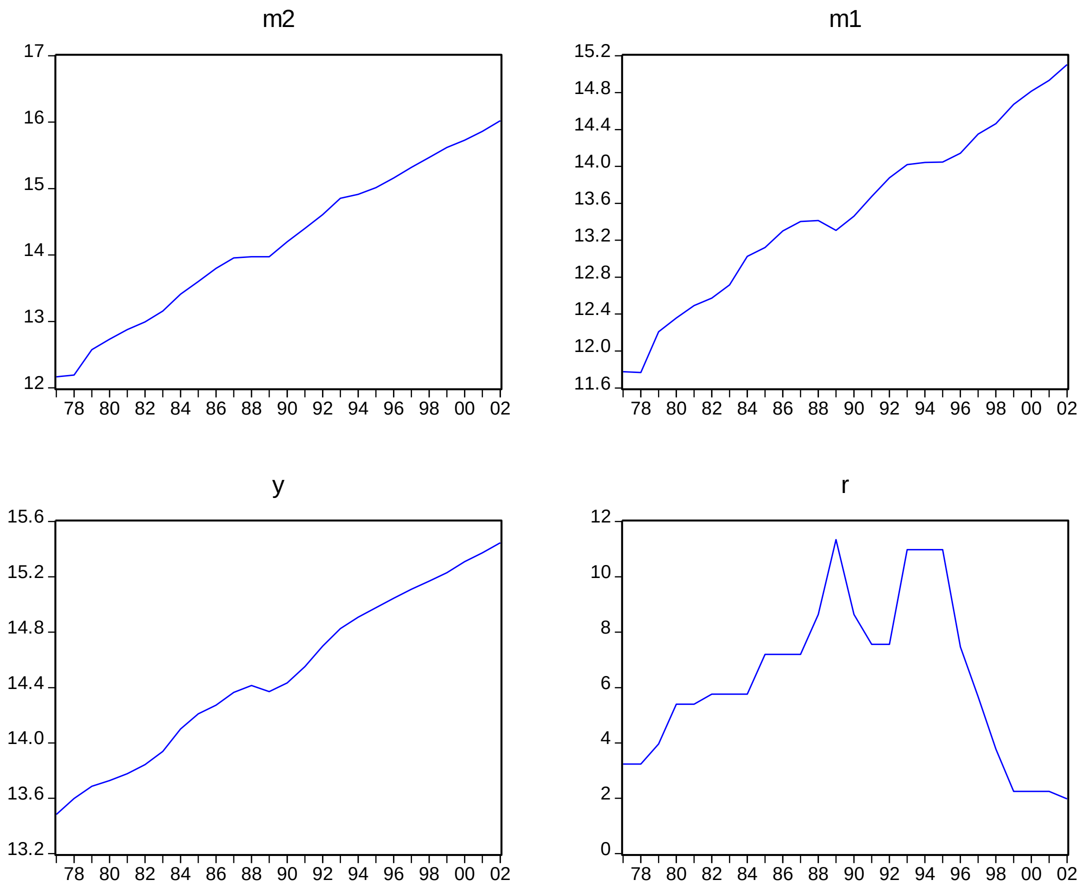

Fig. 1. Logarithmic values of real M2, real M1, real national income, and 1-year time deposit rate, 1977-2002.

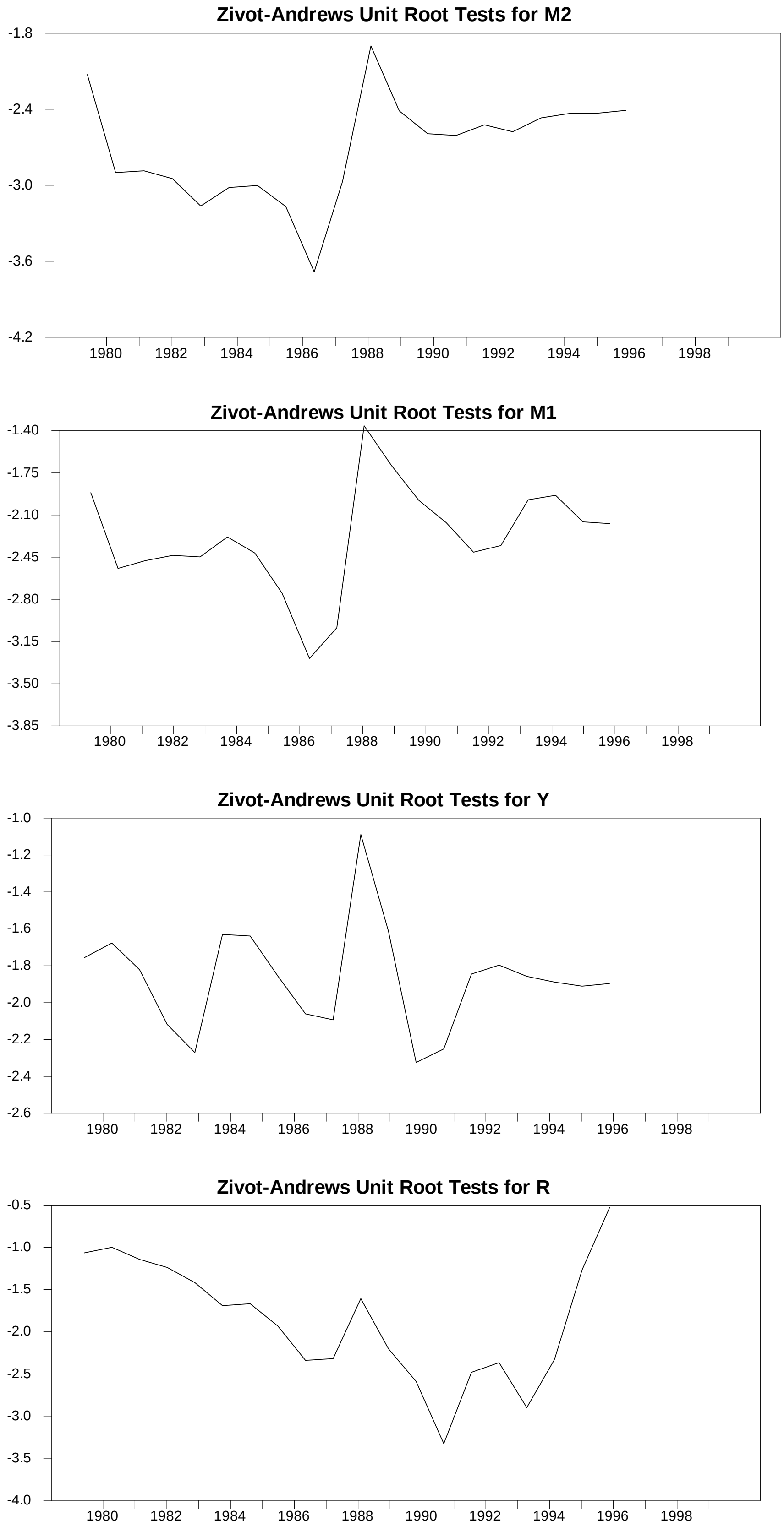

Fig. 2. Plots of Zivot and Andrew (1992) unit-root test.

Table 1.

The Results of ADF Unit-Root Tests

	Levels		First Differences	
	$\tau_{\mu}(0)$	$\tau_{\tau}(0)$	$\tau_{\mu}(1)$	$\tau_{\tau}(1)$
$m2$	-0.51[2]	-2.52[0]	-5.09[0]**	-5.37[0]**
$m1$	-0.73[0]	-2.46[0]	-4.84[0]**	-4.90[0]**
y	-0.93[4]	-2.03[0]	-3.08[0]**	-3.00[0]
r	-1.10[0]	-0.12[3]	-3.69[0]**	-4.09[0]**

Note.

1. $\tau_{\mu}(0)$ and $\tau_{\tau}(0)$ are the test statistics for a unit root in the level with a constant, and with both a constant and a trend, respectively. $\tau_{\mu}(1)$ and $\tau_{\tau}(1)$ are the test statistics for a unit root in the difference with a constant, and with both a constant and a trend, respectively.
2. The critical values for the ADF tests are from the Mackinnon (1991) table.
3. The numbers within the square bracket are the appropriate lag lengths for each interest rate based on MAIC as suggested by Ng and Perro (2001).
4. The bold numbers indicate the appropriate model determined by DJS (1990).
5. ** Significant at the 5% level.

Table 2.

The Results of ZA Unit-Root Tests with Structural Break

	Levels	Year of Break	First Differences	Year of Break
$m2$	-4.04 (C)	1988	-6.47** (C)	1990
$m1$	-3.65 (C)	1988	-5.89** (B)	1990
y	-2.57 (C)	1992	-4.99** (B)	1992
r	-3.86 (B)	1993	-5.51** (B)	1996

Note. The critical values for 5% levels are -4.42 and -5.08 for Model B and Model C, respectively, from Zivot and Andrew (1992). The characters within the parenthesis indicate the appropriate model according to the results from the ADF test.
** Significant at the 5% level.

Table 3.

Bai and Perron's test results for structural breaks

Variable	UDmax	WDmax	$\sup F(1\|0)$	$\sup F(2\|1)$	$\sup F(3\|2)$	$\sup F(4\|3)$	$\sup F(5\|4)$	No. of breaks
$m2$	214.02**	296.01**	67.66**	34.34**	33.56**	10.77	5.70	3
$m1$	150.82**	230.38**	41.26**	31.59**	22.00**	10.09	7.21	3
y	138.06**	221.92**	72.73**	37.72**	22.19**	10.86	10.32	3
r	3613.28**	7233.08**	13.83**	25.97**	52.19**	--	--	3

Notes: ** and * indicate significance at the 5% and 10% levels, respectively. – indicates that there are no more place to insert an additional break given the minimal length requirement. The upper bound M is set to be 5 and the trimming percentage is chosen to be 15% in all cases.

Table 4.

Bai and Perron's estimation results of the structural breaks

Variable	Regime 1	Regime 2	Regime 3	Regime 4
$m2$	12.67**(0.13) 1983[1982：1984]	13.85**(0.13) 1990[1988：1992]	14.83**(0.14) 1996[1995：1998]	15.67**(0.09)
$m1$	12.27**(0.15) 1983[1982：1985]	13.29**(0.08) 1990[1988：1991]	14.02**(0.10) 1997[1995：1999]	14.80**(0.12)
y	13.72**(0.07) 1983[1981：1984]	14.34**(0.07) 1991[1990：1993]	14.89**(0.07) 1996[1995：1998]	15.27**(0.04)
r	3.48**(0.23) 1979[1978：1980]	5.62**(0.10) 1984[1979：1984]	8.57**(0.50) 1997[1996：1999]	2.50**(0.33)

Notes: ** and * indicate significance at the 5% and 10% levels, respectively. The values in parentheses are the standard deviations of the estimates and the 95% confidence intervals of the estimated break dates are given in brackets.

Table 5

Johansen's Maximum Likelihood Cointegrating Tests

$H_0 \mid H_1$	L-max Test	5% Critical Value	$H_0 \mid H_1$	Trace Test	5% Critical Value
model: $(m2, y, r)$					
$r \le 0 \mid r = 1$	20.37**	20.22	$r \le 0 \mid r = 1$	31.71**	27.48
$r \le 1 \mid r = 2$	11.17	12.93	$r \le 1 \mid r = 2$	11.33	14.16
$r \le 2 \mid r = 3$	0.16	4.34	$r \le 2 \mid r = 3$	0.16	4.34
model: $(m1, y, r)$					
$r \le 0 \mid r = 1$	31.15**	20.22	$r \le 0 \mid r = 1$	44.15**	27.48
$r \le 1 \mid r = 2$	12.84	12.93	$r \le 1 \mid r = 2$	13.00	14.16
$r \le 2 \mid r = 3$	0.17	4.34	$r \le 2 \mid r = 3$	0.17	4.34

Note.

1. VAR length is 1 for all the models selected based on the smallest number of SBC. ** Significant at the 1% level.
2. The computed Ljung-Box Q-statistics indicate that the residuals are white noise.
3. The 5% finite-sample critical values are constructed from the asymptotic critical values from Osterwald-Lenum (1992) employing the method in Cheung and Lai (1993).

Table 6.
Test for Weak Exogeneity for the Money Demand Function

Model	$\chi^2_{(l)}$ statistic
m2-y-r	
*m*2	7.73[0.01]**
y	1.76[0.19]
r	0.17[0.68]
m1-y-r	
*m*1	7.18[0.01]**
y	0.01[0.975]
r	0.48[0.46]

Note. ** Significant at the 5% level. The Chi-sq tests for weak exogeneity are the LR test. Numbers in brackets are probability values.

Table 7.
Gregory and Hansen (1996) Tests for Regime Shifts

Model	*m*2		*m*1	
	Test Statistics	Breakpoint	Test Statistics	Breakpoint
ADF*				
C	-5.96**	1993	-4.40	1993
C/T	-6.73**	1980	-4.01	1981
C/S	-5.90**	1993	-4.48	1993
Z_t^*				
C	-4.002	1992	-3.284	1980
C/T	-4.507	1980	-3.430	1987
C/S	-5.804**	1987	-4.105	1985
Z_α^*				
C	-20.510	1992	-16.233	1980
C/T	-23.319	1980	-16.666	1980
C/S	-29.866	1987	-22.336	1985

Note. ** Significant at the 5% level. The critical values are from Table 1 of GH (1996).