5. THE DEMAND FOR MONEY IN CHINA: A REASSESSMENT USING THE BOUNDS TESTING APPROACH

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Abstract

This paper investigates the demand for money in China using annual data covering 1977-2006. To this end, we apply a newly-developed bounds testing technique to overcome the inherent limitations in testing for unit roots prior to testing for the existence of a level relationship between a dependent variable and a set of regressors. Our results clearly identify the long-run money demand relationship among real narrow money (or real broad money), real income, and nominal interest rates for China. The estimated long-run income elasticity and interest semi-elasticity are, respectively, 0.884 (0.915) and -0.034 (-0.002), using the real M1 (M2) equation. Our estimates of the long-run elasticity are consistent with previous studies, but they are towards the lower end of existing estimates. The results of the parameter stability test reveal that both M1 and M2 money demand are stable for China.

Keywords: Money demand, China, ARDL bounds test, CUSUM test
JEL Classification: E41, C22, C52

1. Introduction

The demand for money plays a crucial role in the process of economic development, especially from the macroeconomic perspective that is concerned with how to select and execute an appropriate monetary policy for rapidly-growing economies like that of China. Although throughout much of the past decade many studies have focused on issues pertaining to the money demand function for China, most have examined the relationship between money and income; nevertheless, to be sure, very few have

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investigated the stability of the money demand function and, even if they have done so, they have provided mixed results for China.[3]

This paper's main methodological contribution lies in investigating the demand for money in China using an autoregressive distributed lag (ARDL) bounds testing approach to cointegration proposed by Pesaran and Shin (1999) and Pesaran *et al*. (2001). By combining this approach with the long-run parameters being estimated through a long-run static solution of the estimated ARDL model, the short-run dynamics are then estimated by the error correction model (ECM). We thus make use of the ECM framework to provide both long- and short-run elasticities. We also repeatedly implement the ARDL model, using extended annual and quarterly data with many types of critical variables (the inflation rate, exchange rate, or equity shares, etc.) under different lag lengths to investigate whether the long-run money demand function is stable or not.

The reasons that make studying China's money demand particularly interesting are as follows.[4] First, ever since market-oriented reforms were announced in 1979, China has undergone significant systematic changes. Among these, the ones with the greatest impact have 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 grown at an astounding rate, and money stock has increasingly played a central role in facilitating all kinds of economic activities and market transactions (Zhang and Fung, 2006; Rao and Kumar, 2009). Second, China's slow but steady reforms have induced a dual-track banking control structure. In order to maintain better control over the economy, the banking system has been reformed to separate banking functions, thereby creating new financial instruments. Following the establishment of the central banking system in 1983, the People's Bank of China (PBC) has taken control of the total volume of credit in the economy and has continued to work closely with the State Council when it comes to making important macroeconomic policy decisions. Beyond this, financial reforms have spawned new, far-reaching financial instruments, financial services, and institutions. Turning to one implication of these reforms, we cannot overlook the increased importance of monetary aggregates in policy decision-making.

Previous studies such as that by Mehrotra (2008) have examined the demand for domestic broad money M2 in China during 1994-2005, including outright deflation. This paper establishes a stable money demand relationship in a vector error correction framework, but no interest rate is included in the money demand model due to a lack of data on market-based interest rates. Wang (2008) demonstrates the money demand model in China by incorporating foreign interest rates and the exchange rate on a country-specific, bilateral basis. Wu (2008) investigates an error correction model using a general-to-specific methodology and confirms that a stable broad money demand function exists which appropriately takes asset substitution into

[3] Payne (2003) indicates 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.

[4] See Yu and Tsui (2000) for a detailed discussion.

account. Baharumshah *et al*. (2009) also explore that a long-run relationship exists between M2 and its determinants. Interestingly, Bahmani-Oskooee and Wang (2007) indicate that they employ quarterly data (1983:1 -2002:4) to carry out the empirical analysis on demand for money in China using the ARDL bounds testing approach to cointegration.

In light of our relatively small sample size (1977-2006), we use a somewhat new, and as yet little employed, estimation technique - an ARDL framework. Also worth noting is that we take a fresh look at the stability of China's money demand function with regard to the long-run coefficients in the context of the short-run dynamics using the latest data for the country. Our sample set covers the period after 1991, which few studies have investigated. During this period, further financial reforms in China, including a 16-point financial reform program in 1993, have 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. Finally, as a comparison, robustness tests are made with two alternative models.

A general consensus seems to have been reached whereby the bounds testing approach to cointegration has major advantages that cannot be overlooked. For one, the Monte Carlo simulations studies clearly demonstrate that an ARDL approach outperforms the Engle and Granger (1987), Johansen and Juselieus (1990),5 and Phillips and Hansen (1990) tests, and this is particularly true in the case of small samples (Ghatak and Siddiki, 2001; Pesaran *et al*., 2001).6 The second advantage is that it can be applied irrespective of whether the underlying regressors are purely integrated of degree zero (I(0)), purely integrated of degree one (I(1)), or mutually cointegration (Bahmani-Oskooee, 2002; Dergiades and Tsoulfidis, 2010). However, given that it is possible that the long-run relationship may contain at least one I(0) variable, it is incorrect to talk about cointegration. Lastly and equally important, it corrects for the potential problem of endogenous regressors in the money demand function.

The remainder of this paper is organized as follows. Section 2 briefly outlines the data and model specifications for China. Section 3 describes the bounds testing procedure and the results we obtain. Section 4 provides the concluding remarks and the implications from this study.

2. Data and Model Specifications

In estimating the money demand model, long-run and short-run dynamics can be estimated by various methods. Previous empirical studies on developing countries

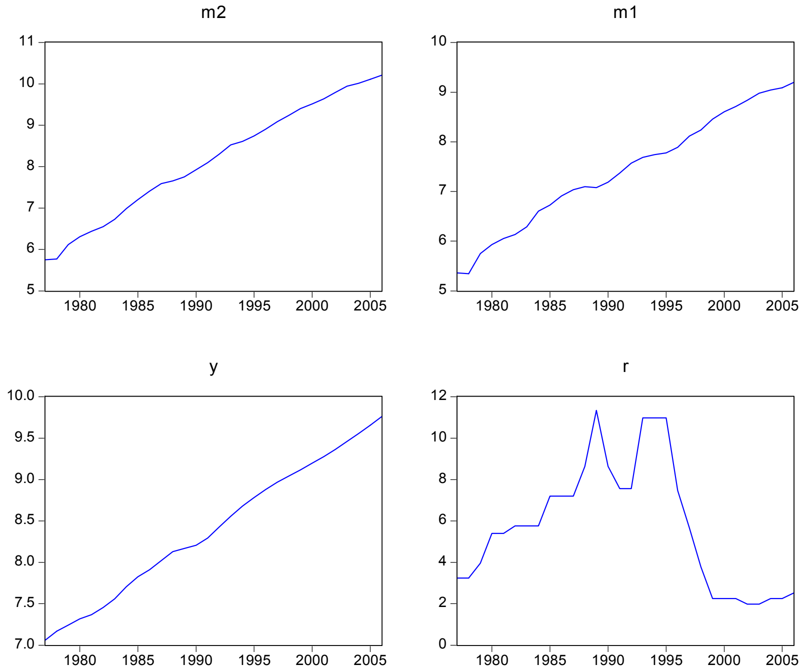
[5] *Badarudin* et al. *(2009) employ the Johansen's cointegration tests to investigate the relationship between bank loans and broad money supply for 10 emerging economies.*

[6] *Notwithstanding the abundance of results from previous empirical research studies, most are based on traditional cointegration tests, but the power of those tests in multivariate systems with a small sample size not only can be severely distorted, but can also be unreliable (Engle and Granger, 1987; Johansen, 1988), not to mention misleading (Mah, 2000). Another potential shortcoming of the above approaches is that they require the underlying time series to be non-stationary and also integrated of the same order.*

indicate that models focusing on narrow money are more reliable when they reflect a weak banking system and low financial sector development. However, over time, broad money accommodates new instruments created as a result of the ensuing development of institutional and financial structures. In line with Pradhan and Subramanian (2003), this paper therefore uses two definitions of the money aggregate, narrow and broad, to model China's demand for money. We select annual data from 1977 to 2006 for real narrow money (*m1*) (currency plus demand deposits held by households and enterprises), real broad money (*m2*) (*m1* plus time and savings deposits held by households and enterprises), real national income (*y*),[7] and the 1-year time deposit rate (*r*) which we obtain from the *AREMOS Economic Statistical Databanks China Statistical Databank.*[8] All real variables are deflated by a GDP deflator (2000=100). All the variables we use are in natural logarithmic form except for the interest rate. Figure 1 presents the data series in graphic form and it is apparent that they are non-stationary and exhibit rather different patterns. While real M1, real M2, and the real national income series show steady growth over much of the last three decades, the time deposit rates display significant variations over time.

Figure 1

Logarithmic values of real M2, real M1, real national income, and the 1-year time deposit rate, 1977-2006 (annual data of the basic model)



[7] *Nominal income measures gross output from agriculture, industry, construction, transportation, and commerce.*

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

We adopt the widely applied Keynesian specification of the demand for money in which real money is a function of real income and the nominal interest rate. Thus, the specification of the semi-logarithmic long-run money demand function is estimated as follows (basic model):

$$m_t = \beta_0 + \beta_1 y_t + \beta_2 r_t + \phi t + \varepsilon_t \ , \qquad (1)$$

where: *m* represents the logged real money balance created by taking a monetary aggregate deflated by the GDP deflator; *y* is the logged real income measured via real national income (scale variables); *r* is the opportunity cost proxied via the 1-year time deposit rate; t is time; and ε is a residual term. The estimated coefficient β_i, $i = 1{,}2$, refers to the elasticity of income and semi-elasticity of the interest rate.[9] The coefficients β_1 and β_2 are expected to have a positive and negative sign, respectively.

Friedman (1956) proposes that research on the money demand function assume that there exists an underlying stationary long-run equilibrium relationship among real money balances, real income, and the opportunity cost of holding real money balances. According to the Cambridge and the Keynesian approaches, there is a direct relationship between real money demand and the level of real income, while there exists an inverse relationship between real money demand and the interest rate. One may think of this in terms of there being an opportunity cost of holding money (Choudhry, 1999; Lee and Chang, 2012). Moreover, to be specific, if the magnitude of the scale variable β_1 is equal to one, then the quantity theory of money applies.

An ARDL model is a general dynamic specification that uses the lags of the dependent variable and the lagged and contemporaneous values of the independent variables, through which the short-run effects can be directly estimated and the long-run equilibrium relationship indirectly estimated. If a long-run relationship exists between *m*, *y*, and *r*, then we write the conditional error correction model (ECM) of interest as follows (basic model):

$$\Delta m_t = \alpha_0 + \theta_1 t + \sum_{i=1}^{p} \alpha_{1i} \Delta m_{t-i} + \sum_{i=0}^{p} \alpha_{2i} \Delta y_{t-i} + \sum_{i=0}^{p} \alpha_{3i} \Delta r_{t-i} + \sigma_1 m_{t-1} + \sigma_2 y_{t-1} + \sigma_3 r_{t-1} + u_t \qquad (2)$$

Here, σ_i ($i = 1{,}2{,}3$) is the long-run multiplier, α_0 is the drift, and t is the time trend. Lagged values of Δm_t and the current and lagged values of Δy_t and Δr_t are used to model the short-run dynamic structure. The bounds testing procedure for the absence of any level relationship among m_t, y_t, and r_t is conducted through an exclusion of the lagged levels variables m_{t-1}, y_{t-1}, and r_{t-1} in equation (2). Thus, we test the null of no long-run relationship between the levels of the variables, defined

[9] *The entire empirical structure is under the assumption that long-run causality goes from real income and nominal interest rate to real balance. However, if the other causality is allowed by theoretical investigations, then if we only change the dependent variable as real income or nominal interest rate of equation (1), we can re-estimate it via bounds testing for the existence of a long-run relationship.*

as $H_0 : \sigma_1 = 0, \sigma_2 = 0, \sigma_3 = 0$, against the alternative $H_1 : \sigma_1 \neq 0, \sigma_2 \neq 0, \sigma_3 \neq 0$, by employing the familiar F-test.

There are two sets of critical values: one set is calculated under the assumption that all of the variables included in the model are I(1), whereas the other is estimated under the assumption that all of the variables are I(0). If the computed F-statistics fall outside the inclusive band, then an inference can be made without a priori knowledge of the order of integration of the variables. However, if the estimated test statistic is higher than the upper bound critical value, then the null hypothesis of no long-run relationship among the levels of the variables is rejected. If the estimated test statistic is lower than the lower bound critical value, then the null hypothesis of no long-run relationship cannot be rejected.

3. Empirical Results

3.1 Results of the Basic Model

This paper uses annual data from 1977 to 2006 to examine the null of no long-run relationship among the variables in the money demand function for China, and both real M1 and real M2 are adopted as the money aggregate. Two steps can be employed in the empirical tests for the ARDL model. The first involves investigating the existence of a long-run relationship between the levels of the variables in the basic model (Equation 2), using the F-test for the upper and lower bounds under new critical values. Owing to the F-test being sensitive to the number of lags imposed on each first-differenced variable, we thus impose one, two, three, and four lag lengths for each first-differenced variable in the basic model. As can be seen, the test outcome varies with the choice of lag order. Table 1 reports the results of the F-test for the existence of a long-run relationship from income and the interest rate to real balances. It is clear from Table 1 that in each of the three cases the estimated F-values are greater than the upper bound critical value, supporting the existence of a long-run relationship among the levels of the variables at the 10% level of significance.

Table 1

Results of F-statistics for the cointegration relationship, 1997-2006 (basic model)

Order of lag	$m1$	$m2$
1	6.990*	1.119
2	1.735	3.384
3	5.602*	2.145
4	4.059	6.815*

*Note: The critical value bounds of $F(m|y,r)$ for upper bound (I(1)) and lower bound (I(0)) with two regressors as 4.19 and 5.06, respectively, at the 10% level of significance. Asymptotic critical values are given in Table CI(v) (unrestricted intercept and unrestricted trend) for Pesaran et al. (2001). * indicates that the F-statistic falls above the 10% upper bound.*

Since lag lengths are selected arbitrarily and without using any criterion, Bahmani-Oskooee and Brooks (1999) and Bahmani-Oskooee and Wang (2007) mention that researchers need an efficient estimation. Thus, we next estimate the long-run coefficients of the money demand function obtained from the long-run static solution of the optimum ARDL model determined by the information criteria, whereby short-run dynamics are estimated by the associated ECM. In the standard rule of the Schwarz Bayesian Criterion (Schwarz, 1978; SBC), the specifications we choose are those of an ARDL (1, 1, 0) for real M1 and an ARDL (1, 0, 0) for real M2.[10] The long-run coefficient estimates are reported in Table 2 using Case V (unrestricted intercept and unrestricted trend) in Pesaran *et al*. (2001).[11] As expected, real income has a significant, positive influence on real narrow money (M1), which is similar to the effect on real M2. Moreover, the long-run income elasticity is significantly smaller than one, which is indicative of a small response of the real money balance to changes in real income - a finding comparable to that of Bahmani-Oskooee (2002) and Pradhan and Subramanian (2003) for the case of Hong Kong and selected developing countries, respectively. This implies that a 1% increase in real income raises real money by around 0.88-0.92%.

Table 2

Estimated long-run coefficients using the ARDL approach for money demand: Selected ARDL model based on the Schwarz Bayesian Criterion, 1977-2006 (basic model)

Model	ARDL (1, 1, 0)	ARDL (1, 0, 0)
Regressor	m1	m2
Trend	0.036** (4.50)	0.059** (3.639)
y	0.884** (27.55)	0.915** (12.71)
r	-0.034** (-2.31)	-0.002 (-0.09)
Wald test	25.83**	1.41

*Notes: Numbers inside the parentheses are the t-values. ** and * indicate significance at the 5% and 10% levels, respectively. The Wald statistic tests the hypothesis that the coefficient of the real income is equal to unity in the long-run equilibrium model.*

The opportunity cost variables, or nominal interest rates, have a negative and statistically significant impact on M1, but an insignificant impact on real M2. We infer the possible reason as being that we use the 1-year time deposit rate for the former, but interest rates of time and savings deposits for the latter, respectively. The semi-elasticity of the interest rate is -0.03 in the long-run real M1 equation. This implies that a 1% increase in the nominal interest rate leads to a 0.03% decrease in real M1, and the interest rate has a larger effect on the demand for real M1 than on that for real M2.

[10] *We ignore the structural breaks and their implications for cointegration tests. However, we argue that there are practical problems in using this test when there are only a limited number of annual observations relative to the number of such possible breaks (see Rao and Singh, 2005).*

[11] *The paper adds a linear trend term in the long-run dynamics when considering other factors that might affect money demand.*

Dreger *et al*. (2007) also find that interest rate elasticity is significantly negative and relatively small, which may reflect the difficulty in controlling money holdings. Hsing (1998) suggests that the demand for real M2 is more (less) sensitive to higher (lower) interest rates. An analysis of the data indicates that nominal interest rates were lower in recent years as a result of financial liberalization and other factors. Thus, interest rate elasticities in absolute value terms were smaller in recent years. In comparison, cash management is more efficient and the demand for real M2 is more sensitive to interest rates in South Korea than in India.

The estimated long-run income elasticity and interest semi-elasticity are, respectively, 0.884 (0.915) and -0.034 (-0.002) using the real M1 (M2) equation. In absolute value terms, the interest rate has a larger effect on the demand for real M1 than on that for real M2. Of particular note is that the income elasticity of money demand for M2 is larger than that for M1, but conversely the interest rate semi-elasticity of money demand for M1 is larger than that for M2. This points out that the growth of money demand exceeds that of real income in China. However, the demand for money grows more slowly than the nominal interest rate. Finally, the linear time trend has a positive and statistically significant impact on money demand. Thus, in the long run we find that ongoing developments in the financial sector might lead to a trend in the money demand.

The short-run dynamics of the demand model for money, presented in Table 3, and the signs of the short-run coefficients are very similar to those of the long-run parameters, except for the nominal interest rate. Being theoretically correct, the coefficients for real income and the nominal interest rate are statistically significant at the 5% level for real M1. The short-run income elasticity of real M1 demand is 1.23, which is significantly higher than the long-run income elasticity, and the short-run nominal interest rate semi-elasticity is found to be -0.01, which is a more inelastic value than its long-run elasticity.

Table 3

Full information estimates from the selected ARDL model, 1997-2006 (basic model)

Regressor	Dependent variable is $\Delta m1_t$ ARDL (1, 1, 0)	Dependent variable is $\Delta\ m2_t$ ARDL (1, 0, 0)
Panel A: Estimated coefficients		
Trend	0.014** (2.23)	0.016 (1.31)
Δy_t	1.233** (2.97)	0.250** (2.33)
Δr_t	-0.013** (-3.37)	$-0.394 * 10^{-3}$ (-0.10)
EC_{t-1}	-0.399** (-3.99)	-0.273** (-1.99)
$\overline{R}^2$	0.61	0.33

Regressor	Dependent variable is $\Delta m1_t$ ARDL (1, 1, 0)	Dependent variable is $\Delta\ m2_t$ ARDL (1, 0, 0)
Panel B: Diagnostic tests		
F-statistics	14.28**	5.50**
J-B	8.56**	2.01
LM, $\chi^2(2)$	0.89	0.51
RESET, $\chi^2(2)$	0.67	1.05

*Notes: Numbers inside the parentheses are t-values. EC_{t-1} is the error correction term, and the F-statistics are the overall F-statistics for the error correction model. LM is the Breusch-Godfrey Lagrange multiplier test for serial correlation. RESET is Ramsey's test for functional form misspecification. J-B is the Jarque-Bera normality test of residuals. * and ** indicate significance at the 5% and 10% levels, respectively.*

Table 3 reports the coefficient estimates of all lagged first differenced variables in the ARDL model when we discuss the M2 demand function. We discover that income elasticity in the short run is 0.25, which is highly significant as reflected by the t-statistic. The interest rate semi-elasticity is negative and significant, supporting our theoretical expectation, and it appears that a decrease in the interest rate in China raises the demand for real M2.[12] Finally, looking at the estimated long- and short-run elasticities for real M2, we find that income policies have stronger effects over time for real M1, while interest policies have weaker effects over time.

The error correction terms (EC_{t-1}) carry the expected negative sign and are statistically significant and similar in magnitude in both the real M1 and M2 money demand models. These results lend credence to the finding of cointegration from both the F-test and the t-test. The magnitude of the error correction term in the basic model (see Table 3) is very near that in Bahmani-Oskooee and Wang (2007) and Baharumshah *et al*. (2009), thus showing the stable adjusted speed in the long run, even when we change the data span and frequency.

In order to investigate the reliability of our conditional error correction model, we apply a number of diagnostic tests, including tests of autocorrelation, Ramsey (1969)'s test, normality and heteroskedasticity in the error term, stability, and the accuracy of the model. Our empirical results suggest that the statistical fit of the model to the data is

[12] Knell and Stix (2005) argue that, insofar as the holding of money involves fixed costs, money demand might not increase on a one-to-one basis compared to increases in income and transactions. Thus, the 'inventory approaches' formally suggest that the income elasticity should lie between 1/3 and 2/3. However, the development of a country's payment system might affect the demand for money. A higher dissemination of payment cards reduces (narrows) money holdings, and one thus expects a negative impact on the income elasticity (Stix, 2004). Fischer (2007) also shows that the elasticity estimates are significant at the 5% level and are clearly less than one. The estimates lie between 0.4 and 0.6, which are consistent with the Baumol–Tobin transactions model with an income elasticity of 0.5.

excellent as indicated by $\overline{R}^2$ (adjusted R^2) — over 60%/30% of the variation in real money M1/M2 is explained by the demand function for money in China.

Parameter stability tests are important, since unstable parameters can result in model misspecification, which has the potential to bias the results (Lee *et al*., 2007). We follow Pesaran and Pesaran (1997) and Bahmani-Oskooee and Bohl (2000) and employ the cumulative sum (CUSUM) and cumulative sum of square (CUSUMSQ) tests proposed by Brown *et al*. (1975) to assess parameter constancy. Figures 2-3 also present a graphical representation of the CUSUM and CUSUMSQ tests of the recursive residuals (Brown *et al*., 1975). The CUSUM test is based on the cumulative sum of the recursive residuals based on the first set of n observations. It is updated recursively and is plotted against the break points. A similar procedure is used when carrying out the CUSUMSQ test, which is based on the squared recursive residuals. This means that, if the plot of CUSUM or CUSUMQ stays within a 5% significance level (as portrayed by two straight lines), then the estimated regression coefficients are generally stable over the sample period. To this end, from the test results shown in Figures 2 and 3, regardless of whether the dependent variable is real M1 or M2, the values of both CUSUM and CUSUMSQ are confined within the 5% critical bounds of parameter stability. Therefore, the stability of the error correction model we use in this study is supported, and this confirms the absence of any instability in the coefficients. The fact that the respective error correction models exhibit stability is clear. From this, the inventory approach and the Keynesian speculative demand for money are supported for China based on our findings.

Figure 2

Plot of the CUSUM test and the CUSUM of Squares test for coefficient stability, basic model ($m1, y, r$)

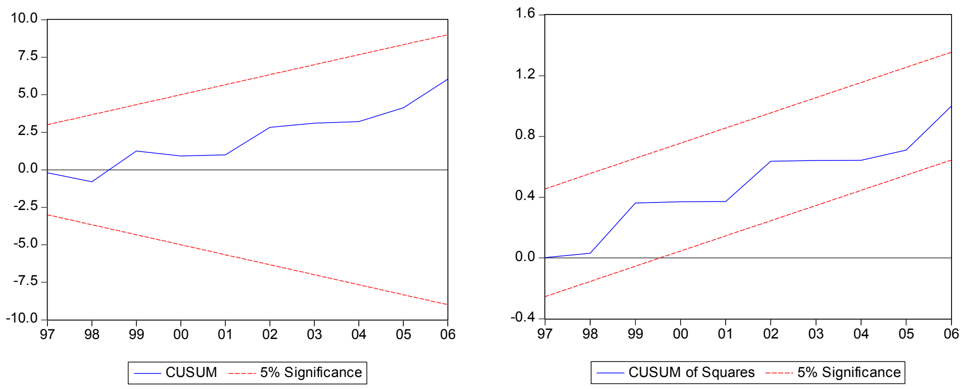
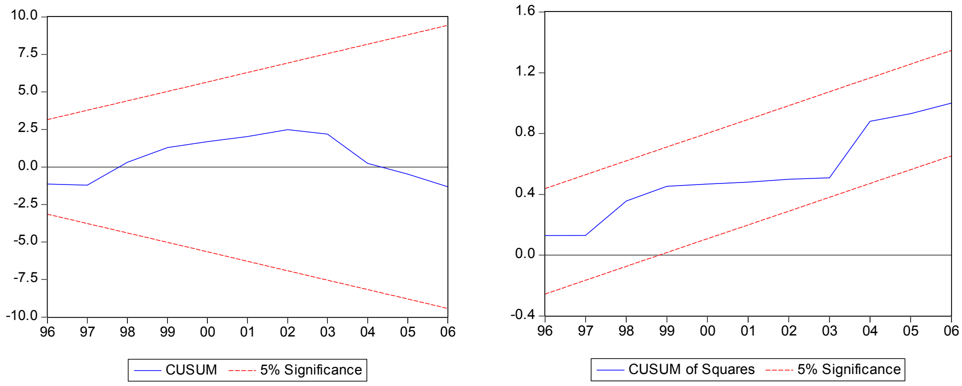


Figure 3

Plot of the CUSUM test and the CUSUM of Squares test for coefficient stability, basic model ($m2, y, r$)



In order to cover any possibility of instability and structural change existing in the cointegrating model, this paper also applies the Hansen (1992) parameter stability test. To examine this issue, we apply three tests of parameter constancy for the I(1) processes proposed by Hansen (1992): the L_C, MeanF, and SupF processes.[13] All three tests have the same null hypothesis of parameter stability, but differ in their alternative hypotheses. Specifically, the Sup F test is useful if we are interested in testing whether there is a sharp shift in the regime, while the L_C and Mean F tests are useful for determining whether or not the specified model captures a stable relationship. We note that Hansen (1992) suggests that these tests may also be viewed as tests for the null of cointegration against the alternative of no cointegration.

Table 4 presents the results of the L_C, Mean F, and Sup F tests with their probability values.

Table 4

Hansen (1992) tests for parameter instability, 1997-2006 (basic model)

Model	L_c	Mean F	Sup F
$(m1, y, r)$	0.233 [0.20]	2.796 [0.20]	0.101 [0.20]
$(m2, y, r)$	0.454 [0.13]	5.200 [0.04]##	7.548 [0.20]

Notes: The probability of parameter instability is in the parentheses. A relationship is stable if the estimated probability is $\geq 20\%$. *## denotes a probability level smaller than 5%.*

[13] *While Lc statistic is based on a sequential search over the entire sample, Mean F and Sup F use only 70% of the observations. All the tests have the same null hypothesis of parameter stability, but differ in their alternative hypothesis (Lee and Chang, 2005).*

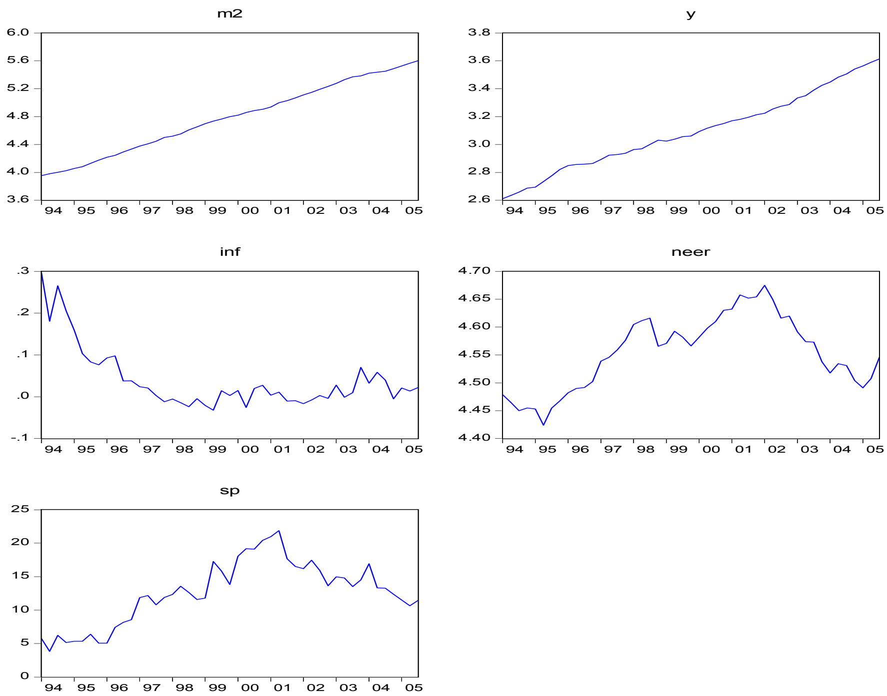
These values imply that we are not able to reject the null of parameter stability with cointegration at even the 20% level except for the model ($m2, y, r$) based on the Mean F tests, when the constant term and trend are included in the specification. The tests show signs of stability, and these results also provide further support in favor of our previous findings such as those in relation to the F-test for the existence of a long-run relationship.

3.2 Some Further Analysis of the Money Demand Function - Annual/Quarterly Data of the Extended Model

The exchange rate and the role of currency substitution are important factors in the money demand function for transition economies (Payne, 2003). Arango and Nadiri (1981) suggest that domestic exchange rate appreciation lowers the domestic currency value of foreign assets, thereby dropping the demand for real money balances. However, currency substitution in favor of domestic assets could be induced once the domestic currency appreciation leads to expectations of further appreciation (Bahmani-Oskooee and Pourheydarian, 1990). The above impact of the nominal effective exchange rate on monetary demand is ambiguous.

Figure 4

Logarithmic values of real M2, real national income, the inflation rate, the nominal effective exchange rate, and the real stock price, 1994Q1-2005Q3 (quarterly data of extended model)



Researchers have argued that the inflation rate can be defined as the opportunity cost of holding money instead of goods (Hosein, 2007; Mehrotra, 2008), whereby inflation affects money demand function in a way predicted by Zhang and Wan (2002). Payne (2000) also finds that the opportunity cost variables of interest rates and inflation rates are statistically insignificant while the exchange rate has a negative impact on money demand.

In order to examine the sensitivity of our findings to the empirical model's specification, including the sample frequency, we follow Mehrotra (2008), whereby the relationship among real money balances, real income, the inflation rate, and the exchange rate may be discussed as below (the extended model):

$$m_t = \beta_0 + \beta_1 y_t + \beta_2 \pi_t + \beta_3 neer_t + \phi t + \varepsilon_t \ . \quad (3)$$

Equation (3) specifies real money balances as a function of real income y, the expected inflation rate π, and the nominal effective exchange rate $neer$. The omission of the nominal interest rate variable reflects a lack of data on market-based interest rates.

Table 5 shows the annual data of the extended model, and we also take one, two, three, and four lags in different lag lengths. While we only consider y and π in the money demand function, a long-run relationship among the variables in the real M1 or M2 money demand function is found when the order of the lag is 1 for M1 and 2 for M2, respectively. Interestingly, once the nominal effective exchange rate $neer$ is joined in the models, more evidence appears that there is an F-statistic that is greater than the critical value, supporting the existence of a long-run relationship among real M1 (M2), real income, the inflation rate, and the exchange rate in an annual data framework.

Table 5

Results of F-statistics for the cointegration relationship (annual data of extended model), 1977-2006

Model	(m, y, π)		$(m, y, \pi, neer)$	
Order of lag	$m1$	$m2$	$m1$	$m2$
1	32.145*	1.065	20.749*	12.834*
2	2.322	5.617*	4.412	22.322*
3	3.017	3.531	7.172*	3.635
4	2.252	0.477	51.872*	2.576

*Note: The critical value bounds of $F(m|y,\pi)$ for the upper bound (I(1)) and lower bound (I(0)) with two regressors are 4.19 and 5.06, respectively, at the 10% level of significance. The critical value bounds of $F(m|y,\pi,neer)$ for upper bound (I(1)) and lower bound (I(0)) with three regressors are 3.47 and 4.45, respectively, at the 10% level of significance. Asymptotic critical values are given in Table CI(v) (unrestricted intercept and unrestricted trend) for Pesaran et al. (2001). * indicates that the F-statistic falls above the 10% upper bound.*

The second robustness test is that we change the frequency of the data. We employ quarterly data over the 1994Q1-2005Q3 period, seasonally adjusted, to carry out the empirical analysis. This is the period for which data on all variables were available. On the one hand, the real money stock M2 ($m2$), real income (y), the inflation rate (π), and the nominal effective exchange rate ($neer$) are offered by Mehrotra (2008).[14] On the other hand, differing from Mehrotra (2008), we follow Wu (2008) and Baharumshah et al. (2009) by taking proper account of asset substitution, and thus the real stock price is included in the money demand function. We therefore employ the stock price sp (Shanghai Stock Exchange Composite Index, from CEIC Data Company Ltd., offered by Wu) in one of the model specifications. Tables 6-8 report these re-estimated results.

Table 6 presents quarterly data of the extended model with F-statistics for a long-run relationship. When compared with Table 1 for which annual data are taken, the cumulative evidence shows stronger long-run relationships among the variables $(m2, y, r)$. This confirms again that there is evidence for which cointegration is supported in the extended model $(m2, y, \pi, neer)$ and $(m2, y, \pi, neer, sp)$, in particular for the former model $(m2, y, \pi, neer)$, as can be seen from Table 6. We again demonstrate that long-run relationships exist in the money demand function for China through the above two robustness tests.

Table 6

Results of F-statistics for the cointegration relationship (quarterly data of extended model), 1994Q1-2005Q3

Model	$(m2, y, r)$	$(m2, y, \pi)$	$(m2, y, \pi, neer)$	$(m2, y, \pi, neer, sp)$
Order of lag				
1	3.485	4.897	6.177*	6.033*
2	5.472*	4.804	4.848*	4.285*
3	5.999*	3.639	4.825*	3.932
4	5.334*	5.347*	14.915*	8.666*

*Note: Asymptotic critical values are given in Table CI(v) (unrestricted intercept and unrestricted trend) for Pesaran et al. (2001), as in Table 5. The critical value bounds of the F-statistic for the upper bound (I(1)) and lower bound (I(0)) with four regressors are 3.03 and 4.06, respectively, at the 10% level of significance. * indicates that the F-statistic falls above the 10% upper bound.*

Table 7 reports the estimated long-run coefficients of the annual data of the extended model, employing the SBC in selecting the lag length for each first-differenced variable. The results of model (m, y, π) reveal that y is significantly and positively associated with both M1 and M2, while π significantly and negatively affects real M1, but not real M2. The positive (negative) effect of real income (the inflation rate) on the money stock supports the theoretical expectation that as real income (the inflation rate) rises, the demand for money rises (falls). This also indicates that people prefer to substitute physical assets for money balances. Hence, when we add the nominal effective exchange rate in the extended model $(m, y, \pi, neer)$, it is clear that in both money demand specifications, real income, the inflation rate, and the exchange rate are highly significant.

[14] *Data sources and definitions are specified in Table 2 of Mehrotra (2008).*

Table 7

Estimated long-run coefficients using the ARDL approach for money demand: Selected ARDL model based on the Schwarz Bayesian criterion, 1977-2006 (annual data of extended model)

Model	(m, y, π)		$(m, y, \pi, neer)$	
	ARDL(1,1,1)	ARDL(1,0,0)	ARDL(1,1,1,0)	ARDL(1,0,1,1)
Regressors	m1	m2	m1	m2
Trend	0.046** (11.32)	0.058** (4.34)	0.036** (5.18)	0.049** (9.84)
y	0.813** (51.14)	0.921** (15.45)	0.802** (58.24)	0.829** (87.07)
π	-0.018** (-3.01)	-0.004 (-0.53)	-0.021** (-3.64)	-0.009** (-3.25)
$neer$			0.175** (2.02)	0.377** (6.95)
Wald test	138.56**	1.77	206.13**	319.77**

*Notes: Numbers inside the parentheses are the t-values. ** and * indicate significance at the 5% and 10% levels, respectively. The Wald statistic tests the hypothesis that the coefficient of the real income is equal to unity in the long-run equilibrium model.*

Table 8 provides similar findings for the variables y and $neer$ based on quarterly data for the extended model.

Table 8

Estimated long-run coefficients using the ARDL approach for money demand: Selected ARDL model based on the Schwarz Bayesian criterion, (quarterly data of extended model), 1994Q1-2005Q3

Model	$(m2, y, r)$	$(m2, y, \pi)$	$(m2, y, \pi, neer)$	$(m2, y, \pi, neer, sp)$
	ARDL(1,0,0)	ARDL(1,0,0)	ARDL(1,0,0,2)	ARDL(1,0,0,2,0)
Trend	-0.159 (-0.58)	-0.021 (-0.53)	0.021** (7.78)	0.024** (5.83)
y	2.109** (2.58)	2.283** (2.06)	0.684** (5.72)	0.544** (2.90)
r	-0.045 (-0.91)			
π		-4.964 (-0.74)	-0.335 (-1.52)	-0.386 (-1.54)
$neer$			0.502** (7.70)	0.593** (5.05)
sp				-0.003 (-0.93)
Wald test	1.84	1.34	6.94**	5.91**

*Notes: Numbers inside the parentheses are the t-values. ** and * indicate significance at the 5% and 10% levels, respectively. The Wald statistic tests the hypothesis that the coefficient of real income is equal to unity in the long-run equilibrium model.*

The positive effects of real income as well as the effective exchange rate on both real M1 and real M2 indicate that increasing national income and the depreciation of the domestic currency again raise the demand for money for both specifications of models. Finally, the real income variable not only carries its positive coefficients in all extended models, but is also highly significant in all cases. Thus, it appears that in China money is held mainly for transaction purposes. However, the long-run real stock price is estimated to be insignificantly negative, which implies that it does not support our contention that real stock prices matter in the analysis of real money demand for China.

When compared with Bahmani-Oskooee and Wang (2007) who employ quarterly data (1983Q1-2002Q4) to carry out their empirical analysis, the signs of both the income elasticity and interest rate semi-elasticity in this study are consistent with their study, but the estimated elasticity values are slightly lower in our extended model with annual data. In the quarterly data of the extended model, although the coefficient of the nominal effective exchange rate is positive and significant in the case of Mehrotra (2008), thereby supporting the wealth effect, it is different from Bahmani-Oskooee and Wang (2007), who show a negative but insignificant effect in their model. As with Mehrotra's arguments, the effect of the nominal effective exchange rate is ambiguous, which is similar to the comments by Arango and Nadiri (1981) who suggest that a domestic exchange rate appreciation would lower domestic wealth and then decrease the demand for real money. However, Bahmani-Oskooee and Pourheydarian (1990) argue that once people see a further appreciation in the domestic currency, currency substitution in favor of domestic assets will occur, which is a similar finding offered by Gunter (2004). Our findings match these debates.

The coefficient of the inflation rate both significantly and negatively affects the real money stock for annual data, but is insignificant in terms of the quarterly data of the extended model. As expected, higher inflation gives rise to a higher opportunity cost of holding money, which is different from Mehrotra (2008). However, as Mehrotra points out, positive inflation shocks on money demand actually lead to the wrong sign, because inflation is an opportunity cost of holding money.

While our paper is close to that of Bahmani-Oskooee and Wang (2007) and Baharumshah *et al*. (2009), we fill some gaps as follows. First, the above two studies mainly employ quarterly data to perform an empirical analysis on the demand for money in China using the ARDL approach. Similar to them, apart from selecting extended annual data from 1977 to 2006, we also employ seasonally adjusted quarterly data over the 1994Q1-2005Q3 period to study the demand for money and its determinants. Second, we include the crucial variable of real stock prices, which reflects the wealth effect and asset substitution, in our model. Third, we evaluate the Pesaran *et al*. (2001) model, including Case V (unrestricted intercept and unrestricted trend), but the ARDL model used by the above two studies does not consider the linear time trend.[15] Fourth, we adopt the Hansen (1992) approach in order to cover the possibility of parameter instability and structural change existing in the cointegration model. Finally, to demonstrate that the market in China does not determine the

[15] *Baharumshah* et al. *(2009) employ quarterly seasonally unadjusted data from 1990:4 to 2007:2 to investigate the demand for M2 without a deterministic time trend in China.*

nominal interest rate, we instead use the inflation rate as the proxy variable for the interest rate to carry out a robustness test. The omission of the nominal interest rate variable reflects a lack of data on market-based interest rates (Hosein, 2007; Mehrotra, 2008).[16]

4. Concluding Remarks

This paper examines the long-run demand for money in China using annual data covering the period 1977-2006. By applying a newly-developed bounds testing technique, we are able to identify a long-run relationship between real narrow money (M1) or real broad money (M2), real income, and nominal interest rates. The CUSUM (CUSUMSQ) and Hansen (1992) tests also both confirm the stability of the long-run coefficients of the money demand function. This indicates that the structures of the parameters have not diverged abnormally over the period of the analysis. Therefore, money supply is the appropriate monetary policy instrument for China's central bank.

Based on our basic model, overall the long-run coefficients of real income have a significantly positive influence on real M1, which is similar to the effect on real M2. Besides, nominal interest rates have a negative and statistically significant impact on real M1, which contrasts with an insignificant impact on M2. Beyond this, the error correction terms are statistically significant, and both the M1 and M2 models have similar signs.

Our estimates of long-run elasticity are consistent with previous studies, though they run towards the lower end of existing estimates. We also provide a greater discussion regarding the money demand function using annual/quarterly data (an extended model with annual and quarterly data). By combining the inflation rate, exchange rate, and Shanghai Stock Exchange Composite Index in our extended model, we are again able to find that long-run stable relationships exist in the money demand function for China.

Based on the above, some implications do stand out. Understanding the form of a stable money demand function is important for predicting the effects of monetary policy. For instance, in intermediate macro theory this is required in order to know the slope of the money market equilibrium (LM) curve. This information is also important for the government's money management in formulating appropriate income and interest rate policies.

The estimates of the income and interest rate elasticities of money demand are important pieces of information when formulating money policies on restructuring, especially given that financial reform is a crucial component of reform for China. In China, money is held mainly for transactions purposes, Furthermore, money demand does not appear to be very sensitive to interest rates, possibly reflecting their partial

[16] *In addition, this paper considers four assets held in China: broad money, domestic goods, equity shares, and foreign currency. Our paper differs from Wang (2008) in four respects. First, Wang only reports the long-run elasticities. Second, we include a linear time trend in the cointegrating space. Third, instead of using the Johansen method, we use the bounds testing procedure to cointegration within an ARDL framework. Fourth, no paper uses the Hansen (1992) parameter stability test for the I(1) processes.*

liberalization (Wu, 2008). Once again, raising interest rates in China tends to reduce the demand for money, particularly the demand for real M1. This is indicative of the fact that monetary policy may not necessarily achieve its maximum effectiveness. Thus, policies to stabilize the economy must go beyond monetary policy alone (Akinlo, 2006).

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