Money demand in China and time-varying cointegration

Haomiao Zuo\textsuperscript{a,}\textsuperscript{*}, Sung Y. Park\textsuperscript{b}

\textsuperscript{a} The Wang Yanan Institute for Studies in Economics, Xiamen University, Xiamen, Fujian 361005, China
\textsuperscript{b} Department of Economics, The Chinese University of Hong Kong, Shatin, N.T., Hong Kong

\textbf{ABSTRACT}

Many studies analyze the money demand using a (fixed coefficient) cointegrating regression model, which may not be appropriate to deal with the money demand of a transition economy like China. This paper investigates this issue using a time-varying cointegration approach based on the quarterly data from 1996 to 2009. We find some interesting results: (i) the estimates of the income elasticities are between 0.60 and 0.75, which are comparable with the previous studies; (ii) the estimated interest rate elasticity supports the argument that the overall effect of the interest rate on the money holding is weak although there are some mild evidences that it has been strengthened in recent years; (iii) the substitution effect of equity asset dominates the wealth effect, especially, during the bullish market period. Our result is robust to the alternative choices of the scale or opportunity cost variables and shows that omission of the stock prices in the money demand function would possibly yield a misspecification problem.

\textbf{1. Introduction}

Traditionally, the long-run money demand is of great concern to both economists and policy makers. The income elasticity measures the speed of monetary expansion in the long-run while the interest rate elasticity represents the sensitivity of household’s willingness to hold money with respect to the change of monetary policy. Moreover, the central bank’s effort for controlling money supply and selecting valid policy instruments crucially depends on the relationship between the quantity of money and some key indicators of real economy. Numerous efforts have been made to investigate the above issues for both developed and developing countries in the literature, for examples, Judd and Scadding (1982), Ericsson (1998) and Sriram (2001).

Equipped with a cointegration approach introduced by Engle and Granger (1987) and further developed by Johansen (1988) and Johansen and Juselius (1990), many recent studies perform cointegration tests to find an evidence of the long-run stability of the money demand function. Applications along this line include Hafer and Jansen (1991), Hoffman and Rasche (1991), McNown and Wallace (1992) for the U.S.; Bahmani-Oskooee and Bohl (2000) for Germany; Adam (1991), Johansen (1992) for the U.K.; Muscatelli and Spinelli (2000) for Italy; Bahmani-Oskooee and Shabsigh (1996) and Bahmani-Oskooee (2001) for Japan, among many others. A general consensus reached by these studies is that both the broad money (M2) and narrow money (M1) are cointegrated with disposable income and interest rates as well as some other variables. In addition to the above studies for the developed countries, there are also a few studies investigating the money demand in the developing economies, such as Baharumshah (2004) for Malaysia.

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Corresponding author. Tel.: +852 2609 8001; fax: +852 2603 5805.
E-mail addresses: zuohaomiao@gmail.com (H. Zuo), sungpark@cuhk.edu.hk (S.Y. Park).

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Thornton (1996) for Mexico; Bahmani-Oskooee (1996) for Iran, among many others. In particular, the developing economies are significantly different in terms of economic openness, financial market maturity and macroeconomic environment, which poses new challenges in modeling and estimating the money demand function. For example, the issue of stability of the long-run money demand is challenging if a country has institutional changes, which is exactly the case of China.¹

Since Chinese economy is transforming from a centrally planned economy to a market oriented one, the financial system as well as monetary policy are under transition (He, 2005; Prasad & Rajan, 2006). As pointed out by Prasad and Rajan (2006), the China’s economic reform beginning in late 1970s took incremental and experimental processes, particularly, in the financial sector.

The reform has been processed from local and small-scaled experiments to global and large-scaled experiments. This is if the result of the experiments turned out to be satisfactory, then they have been implemented elsewhere.² Such changes have the common feature that the process is smooth and gradual without distinct or sharp regime shifts although the reforms themselves are very fundamental. Even though the recent economic growth of China is remarkable, China remains as a developing economy, and, especially, the financial market is still under-developed. For instance, the bond market in particular, the corporate bond market, is relatively small and inactive (Hansan, Wachtel, & Zhou, 2009). Under the strict regulation of the government, only large state owned companies are allowed to issue bonds with high credit rankings similar to those of the government bond. This results in a relatively low and unattractive nominal interest. Moreover, there are no financial futures and derivatives market in China. As far as the reforms of the financial system are concerned, three important aspects of financial institutional changes are: (i) the reform of the banking system; (ii) the exchange rate reform; and (iii) the rapid development of the capital market. The below we briefly summarize these three important aspects.

First, since the Central Bank Law and the Commercial Bank Law were introduced in 1995, the People’s Bank of China (PBC) has been authorized to implement monetary policy although its freedom and independence have been limited. The reform of banking system has been started from 1997 with carving out of non-performing loans (NPLs) and injecting more capital for four largest state-owned commercial banks (SOCBs) (García-Herrero, Gavilá, & Santabárbara, 2009). Moreover, the partial privatization and minority foreign ownership were introduced to improve the efficiency of the banking system (Berger, Hasan, & Zhou, 2009; Lin & Zhang, 2009; Jia, 2009). By doing these, more and more city commercial banks began to achieve better banking performance (Ferri, 2009). Since 2004, the commercial banks were not restricted by the ceiling of the lending rates or the floor of deposit rates although the floor of lending rates and the ceiling of deposit rates were not yet relaxed. Moreover, the interbank market and the bond market rates had been deregulated since 1996, and the Shanghai Interbank Offered Rate (SHIBOR) was introduced in 2007 to serve as the benchmark interest rate in the Chinese money market. All of these indicate that interest rate liberalization is on the way although the process is rather slow (Koivu, 2009).

Second, the (market-oriented) floating exchange rate regime pegging to the basket of multiple currencies was introduced in 2005. This reform allowed the exchange rate to be more flexible and more sensitive to market conditions.

Third, the most significant change in the financial sector is the rapid development of the stock market. The market capitalization was 1.8 trillion RMB ($ 0.22 trillion USD) with only 3.4 million investor accounts at the beginning of 1998. At the end of 2008, the investor accounts grew to over 100 millions with total market capitalization 12 trillion RMB ($ 1.8 trillion USD). Due to the financial under-development and relatively attractive returns from the stock market, it served as the main alternative choice for domestic investors to the saving deposits. Traditionally, the saving deposits are the main investment channels for domestic investors in China because of financial system repression (Prasad & Rajan, 2006). In total, household deposits reached 24.7 trillion in RMB (or 3.6 trillion USD) at the end of the first quarter of 2009, which were very large amount compare to the annual GDP that is around 30.1 trillion in RMB ($ 4.4 trillion USD in 2008). Not only the scale of savings is large, but also its proportion to the disposable income is very high. Chamon and Prasad (2008) estimate the saving ratio (the ratio of savings to disposable income) for urban households and report that it increased from 17% in 1995 to 24% in 2005. Although the saving ratio is high, the low or even negative real interest rate during those periods generates quite low returns to the savings. Hence, households have strong incentives to withdraw their money from the banks and invest them in the stock market. As a result, large amount of money tends to transfer from the banks to the stock market during the bull market periods. Since the saving deposits are measured based on the usual category of the broad money (M2) whereas money accounts in the stock market are not, there is clearly the substitution effect for the broad money demand. Wu (2009) shows that share of equity assets in household holdings of financial assets varies with the stock market cycles. During the recent two bullish markets in 2001 and 2007, these ratios were about 15% and 20%, respectively. However, during the bearish market in 2005, it dropped to slightly about 5%. As Wu (2009) pointed out, the expanding investment possibility is likely to influence household investors’ decision in choosing between money and other financial assets, and therefore, affecting the aggregate money demand as a whole.

Many papers study China’s money demand in the process of economic transition and financial reform. The list includes Hafer and Kutan (1994), Huang (1994), Qin (1994), Chen (1997), and Bahmani-Oskooee and Wang (2007). Hafer and Kutan (1994) find the existence of a long-run stable relationship for nominal money demand. In the broad money (M2) case, the elasticities of income and GDP deflator are 1.33 and 1.52, respectively. This finding indicates that the velocity of money decreases as income rises. This declining velocity of money is also observed in other developing countries (Bordo & Jonung, 1987). Furthermore, they report that the interest rate elasticities are 0.13 and 0.15 for M0 (currency in circulation) and M2, respectively. In another study, Huang (1994) reports that the income and deflator elasticities for M2 (nominal) are 2.12 and 1.56, respectively, while Chen (1997) shows that the income elasticities for M0 and M2 are 1.50 and 1.93, respectively. Furthermore, Bahmani-Oskooee and Wang (2007) provide the advantages and potential limitations of this approach. Lau, Qian, and Roland (2001) label this as a dual-track approach.

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¹ Hansan, Wachtel, and Zhou (2009) analyze three aspects of institutional changes in China: financial development, legal institution and political system.
² Prasad and Rajan (2006) provide the advantages and potential limitations of this approach. Lau, Qian, and Roland (2001) label this as a dual-track approach.
(2007) show that the long-run income elasticities for M0 and M2 are 1.28 and 1.69 and the interest rate elasticities are \(-4.52\) and \(-1.54\), respectively. However, their estimates of the interest rate elasticity for M2, and foreign interest rate and nominal effective exchange rate elasticities for both M1 and M2 are not statistically significant. Table 1 provides a survey of some recent studies on the money demand in China.

Although the above studies provide useful information about the income and interest rate elasticities, their results are not quite consistent. Even some of the estimated interest rate elasticities turn out to have different signs. This is not surprising because China experiences a gradual institutional transition. Under such situation the above models could confront a (smooth) structure change problem. Since the above models are based on the assumption of the parameter constancy in the regression function, the magnitude or sign of the estimated parameters could be dramatically changed depending on the choice of particular samples. In order to detect the parameter inconsistency problem, Bahmani-Oskooee and Wang (2007) employ CUSUM and CUSUMSQ tests proposed by Brown, Durbin, and Evans (1975) in combination with cointegration analysis, and they find that M1 money demand function is stable but this is not the case for M2. Hafer and Kutan (1994) take care of the parameter inconstancy by introducing time dummy variables. However, Hansen (1992) points out that a stability test based upon a priori specified break date is problematic since an exogenous date choice is conditional on the data, which invalidates the conventional critical values. Thus, he proposes a new testing procedure to overcome this problem. Even though the above mentioned tests such as CUSUM, CUSUMSQ and Hansen (1992)’s tests can be quite helpful in detecting the date of the structural change, these tests do not provide any guidance of further steps under alternative hypothesis. Unfortunately, for an economy under the gradual institutional transition like China, the parameter inconstancy due to the rapid development of financial institutions and markets is a stylized fact rather than merely assumption. Indeed, in addition to Bahmani-Oskooee and Wang (2007) and many others, Baharumshah, Mohd, and Yol (2009) also argues that the money demand function for M2 is unstable in China unless taking stock prices into account.

In this paper, a smooth time-varying cointegrating regression approach is considered to take care of the parameter inconstancy problem in China money demand function. The traditional cointegration method is based on the assumption of the constant cointegrating vectors. This assumption is partly responsible for the empirical failure of the traditional cointegration approach in many cases as pointed by Park and Hahn (1999). For example, the traditional cointegration test cannot reject the null of no cointegrating relationship among variables even though there exists a time-varying long-run relationship. To handle this problem, Park and Hahn (1999) propose the smooth time-varying coefficient cointegrating regression method and prove that the estimators of the time-varying coefficients (TVC) are consistent, asymptotically efficient and normally distributed. The main advantages of our study include: (i) we do not need assumptions on possible dates of structural breaks, which is particularly applicable in China where reform is fundamental yet gradual with few sharp changes so that choices of break point date could be arbitrary; (ii) while the traditional approach splits the whole sample into multiple sub-samples, we utilize the information in the entire sample to detect structural breaks.

| t1.1 | Table 1 | A selective survey of some recent studies on money demand in China.
| t1.2 | | Author(year) | Sample(F) | Scale variable | Money aggregates | Interest rates | Other variables | Method | Stability test |
| t1.3 | | | | | | | | |
| t1.4 | Yi (1993) | 52–89(a) | realNI(pop) (0.7–0.9) | realNI(pop) | M0 | log(UP) (0.8–0.9) | RSL, IM (–1.63–1.14) | DLAG, CUSUM |
| t1.5 | Hafer and Kutan (1994) | 52–88(a) | realNI(1.13) | realNI(1.00) | M6 | NIDEF(2.48) | EL, JJ | Hansen |
| t1.6 | Qin (1994) | 78.1–92.4(a) | realGDP(1.00) | realGDP(1.00) | M0 | NIDEF(0.01) | RLS, IM (–1.63) | ADL, CUSUM |
| t1.7 | Chen (1997) | 52–91(a) | realNI(1.5) | M0 | NIDEF(0.01) | RLS, IM (–1.63) | CUSUM, CUSUMSQ |
| t1.8 | Deng and Liu (1999) | 51–91(a) | realNI(1.5) | M0 | NIDEF(0.01) | RLS, IM (–1.63) | CUSUM, CUSUMSQ |
| t1.9 | Bahmani-Oskooee and Wang (2007) | 80.1–94.12 (m) | realGDP(1.28) | realGDP(1.28) | M1 | NIDEF(0.01) | ADL, ECM | CUSUM, CUSUMSQ |
| t1.10 | | 83.1–02.4(a) | realGDP(1.69) | realGDP(1.69) | M2 | NIDEF(0.01) | ADL, ECM | CUSUM, CUSUMSQ |
| t1.11 | Mehrotra (2008) | 94.1–05.3(q) | realGDP(1.73) | realGDP(1.73) | M2 | NIDEF(1.181), E(0.967) | JJ | RLS |
| t1.12 | Wu (2009) | 94.1–08.1(q) | realGDP(0.74) | realGDP(0.74) | M2 | NIDEF(1.181), E(0.967) | JJ | RLS |
| t1.13 | Baharumshah, Mohd, and Yol (2009) | 90.4–05.3(q) | realGDP(0.65) | realGDP(0.65) | M2 | NIDEF(1.181), E(0.967) | JJ | RLS |
| t1.14 | | 90.4–07.2(q) | realGDP(1.06) | realGDP(1.06) | M2 | NIDEF(1.181), E(0.967) | JJ | RLS |

full sample instead of part of it. This has an advantage especially in case of China in which the sample size for time-series data is usually small; and (iii) when a time-varying cointegrating relationship is detected, we can still interpret it as a long-run relationship where the coefficients may help us to learn more about mechanism of the economy behind it.1

The organization of the paper is as follows. Section 2 provides a brief explanation of the money demand model, the estimation method and cointegration test statistics. The results of the estimation of TVC cointegrating regression along with the fixed coefficient (FC) cointegrating regressions are reported in Section 3. Finally, Section 4 offers some concluding remarks.

2. The model

The traditional long-run money demand function usually takes the following form

\[ \frac{M}{P} = f(S, OC), \]

(1)

where \(M/P\) represents the demand for real money balances defined by the ratio of the selected monetary aggregate in nominal terms \(M\) and the overall price index \(P\). \(f(\cdot, \cdot)\) denotes a function that usually takes log-linear (or semi-log-linear) form, \(S\) denotes a scale variable reflecting the economic activity or transactions need, and \(OC\) is a variable for the opportunity cost.

For monetary aggregates, we choose the broad money (M2) since the People’s Bank of China (PBC) considers it as an important monetary policy instrument and announces its annual target for the growth rate of M2 until 2008. For the scale variable \(S\), real gross domestic product (GDP) is usually chosen as a proxy, and we also use real industrial value added (IVA) as an alternative measure following Koivu (2009). As for opportunity cost variables, Sriram (2001) argues that institutional changes, regulation policies as well as development of the financial markets deserve special attention. In practice, the domestic interest, foreign interest, exchange and inflation rates are the most frequently used variables for the opportunity cost. Since China has been under strict capital flow regulation, which include, for examples, a restriction on holding foreign currencies and a pegged nominal exchange rate system with respect to U.S. dollar from 1995 through 2005, we neither introduce foreign interest rates nor exchange rates as determinants of the money demand. Another important opportunity cost variable is the real interest rate obtained by subtracting annual expected inflation from nominal saving deposit rate. Since saving deposit rate is set by the PBC and rarely changes during certain periods, a large portion of variation in real interest rate is attributed to changes in inflation rate. For this reason we also consider quarterly expected inflation rate as another measure of the opportunity cost since it could serve as the proxy of yield on real assets.

Due to the rapid development of the stock market in recent years, a few studies introduced the stock prices as an additional determinant of the demand for real money (Choudhry, 1996; Baharumshah, 2004; McCornac, 1991; Baharumshah, Mohd, & Yol, 2009). Using the postwar data of U.S., Reynard (2004) concludes that further studies should pay more attention to the financial market development. Also, Baharumshah, Mohd, and Yol (2009) investigate the role of the stock prices to the money demand in China. They show empirically that real stock prices have a positive effect on the money demand in the long-run whereas its short run influence is negative. In view of these findings, we extend the traditional money demand function by including the stock prices:

\[ \frac{M}{P} = f(S, OC, SP), \]

(2)

where \(SP\) denotes real stock prices.

Based on the previous studies, we consider the following linear cointegrating regression equation for the money demand

\[ d_{mj} = \pi + \alpha_1 y_{mj} + \alpha_2 r_{mj} + \alpha_3 s_{mj} + \varepsilon_m + \varepsilon_{mj}, \quad m = 1, 2, 3, 4, \quad j = 1, 2, ..., T, \]

(3)

where \(d_{mj}\) denotes demand for real money balance, \(y_{mj}\) is real GDP or real IVA, \(r_{mj}\) is real domestic interest rate, or it can be replaced by expected inflation \(e_{mj}\), \(s_{mj}\) denotes real stock prices, \(\varepsilon_m\) is the seasonal dummy, \(\varepsilon_{mj}\) denotes the error term, and subscripts \(m\) and \(j\) represent the quarter and year, respectively. In the above Eq. (3), \(\alpha_2\) is the interest rate semi-elasticity since the interest rate is measured in percentage (we use the term “interest rate elasticity” or “interest rate semi-elasticity” interchangeably unless indicated), and \(\alpha_1\) and \(\alpha_3\) are the income and stock price elasticities, respectively.

According to Sriram (2001), \(\alpha_1\) is likely to be positive to reflect the transaction or the income effect and \(\alpha_2\) is positive for own rate of interest or negative for the opportunity cost variable. In our case, the sign of \(\alpha_2\) deserves more discussions. Traditionally, money refers to cash and cash equivalents, and the interest rate is the proxy of the opportunity cost. For example, Goldfeld and Sichel (1990) study the demand for M1 (currency plus checkable deposits) in the U.S. For narrow money (M1), the explicit yield is equal to zero. Goldfeld and Sichel (1990) mention that, under this circumstance, the saving deposit rate or some short term interest rates are proxies of opportunity cost. However, for the saving deposits, the saving rates are not variables for the opportunity cost but own rate of interest. High interest rates would provide incentives for household to put more money in the

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saving deposits holding other conditions unchanged. In our case, since we are using the one year saving deposit rate as a variable for interest rate, it is the opportunity cost for the M1 component of M2 while it serves as own rate variable for the saving deposit in M2. Thus the overall sign of \( \alpha_2 \) depends on the relative strength of two effects. Since the ratio of saving deposit in M2 to that in M1 is 1.74 at the end of 2008, we may expect a relatively stronger role of the saving deposits as own rate of interest. If this is the case, the sign of \( \alpha_2 \) is more likely to be positive.

The sign of \( \alpha_3 \) is determined by two effects, say, the substitution and wealth effects. On one hand, when the stock prices increase, equity assets are more likely to be substitutes for the saving deposits. If this substitution effect dominates the wealth effect, \( \alpha_3 \) tends to be negative. On the other hand, when the nominal wealth of households increases and/or the growth of the stock market entails more transaction needs, a large fraction of the investors’ assets would be converted to more liquid alternatives to facilitate consumption or transaction. When this wealth effect dominates the substitution effect, \( \alpha_3 \) becomes positive. Thus the sign of \( \alpha_3 \) is a purely empirical issue.

The model (3) can be consistently estimated using the ordinary least squared (OLS) method when variables are cointegrated. Although the OLS estimator is (super-) consistent, it is usually asymptotically biased and inefficient when there exists an endogeneity problem in Eq. (3). Thus, generally, the statistical tests based on the OLS estimator are invalid. There are many methods for efficient estimation of the coefficient in the previous studies, for examples, fully modified OLS (FM-OLS) (Phillips & Hansen, 1990), the usage of sieve approximation (Saikkonen, 1992; Stock & Watson, 1993), canonical cointegrating regression (CCR) (Park, 1992) among others.

It is quite natural to ask a question whether the long-run stable relationship is constant or time-dependent. It may be hard to determine on \([0, 1]\), \( m \) is the order or \( \alpha \) is a smooth function defined on \([0, 1]\), \( n \) is the number of observations, and \( t \) is the order or observations in the total sample given by \( t = 4(j - 1) + m \). The model (5) is the time-varying coefficients (TVC) version of the fixed coefficients (FC) model (3).

Park and Hahn (1999) suggest to use the Fourier flexible form to approximate the smooth time-varying parameter \( \alpha_{mj} \)

\[
\alpha_k(r) = \beta_{k,1} + \beta_{k,2}r + \sum_{i=1}^{k} (\beta_{k,2i+1}, \beta_{k,2(i+1)}) \phi_i(r),
\]

where \( \beta_{k,j} \in \mathbb{R}^p \), \( p \) is the dimension of \( \alpha_{mj}, j = 1, 2, \ldots, 2(k + 1) \), and \( \phi_i(r) = (\cos 2\pi ir, \sin 2\pi ir)' \). The \( \alpha_k(r) \) in Eq. (6) can be rewritten as

\[
\alpha_k = \left( \begin{array}{c} f_k \otimes I_p \end{array} \right) \beta_k, \tag{7}
\]

where \( f_k(r) = (1, r, \phi_i(r), \ldots, \phi_k(r))' \) with \( r \in [0, 1] \), \( \beta_k = (\beta_{k,1}, \beta_{k,2}, \ldots, \beta_{k,2(k+1)})' \), \( I_p \) is a \( p \times p \) identity matrix, and \( \otimes \) denotes the Kronecker product. Thus the model (4) is represented by

\[
d_{mj} = \pi + \beta_k' x_{kmj} + \epsilon_m + \epsilon_{kmj}, \tag{8}
\]

where

\[
x_{kmj} = f_k \left( \frac{r}{n} \right) \otimes x_{mj} \quad \text{and} \quad \epsilon_{kmj} = \epsilon_{mj} + (\alpha - \alpha_k) \left( \frac{r}{n} \right) x_{kmj}. \tag{9}
\]

The standard ordinary least square (OLS) estimators can be used to estimate the Eq. (8). However, OLS estimators are asymptotically inefficient and have non-standard limiting distribution so that standard inference procedures cannot be applied in this case. In order to deal with these problems Park and Hahn (1999) use the canonical cointegrating regression (CCR)
transformation. This transformation involves some unknown parameters, mainly the conditional and one-sided conditional long-run variances of the residuals. However, these unknown parameters can be consistently estimated using nonparametric method (Andrews, 1991). After obtaining the estimates of \( \beta_k \) using the Eqs. (8), (7) can be used to recover \( \alpha_0 \). Park and Hahn (1999) show that under certain assumptions

\[
M_{nk}^{1/2} (\hat{\omega}_k^{\prime} - \omega_k) \rightarrow_d N \left( 0, \omega_k^2 I_{pd} \right) \quad \text{as} \quad n \rightarrow \infty,
\]

where \( \Pi(\alpha) = (\alpha_1, \ldots, \alpha_d)' \) and \( \Pi(\hat{\alpha}_k) = (\hat{\alpha}_k(r_1), \ldots, \hat{\alpha}_k(r_d) )' \) for \( r_i \in [0, 1] \), \( i = 1, \ldots, d \), \( I_{pd} \) is a \( pd \times pd \) identity matrix, \( \omega_k^2 \) is the conditional long-run variance of the residuals from the transformed regression. For more details on this method, see Park and Hahn (1999).

The two tests proposed by Park and Hahn (1999), both of which are based on the Wald-type variable addition tests (see Park, 1990), will be employed to check whether the model is correctly specified. The first test performs under the null hypothesis of TVC model against the alternative that the regression is spurious. The test statistic is given by

\[
\tau^* = \frac{RSS_{TVC} - RSS_{sTVC}}{\hat{\omega}_s^2},
\]

where \( RSS_{TVC} \) and \( RSS_{sTVC} \) are the sum of squared residuals from Eq. (5) and (5) augmented with \( s \) additional superfluous regressors and \( \hat{\omega}_s^2 \) is a consistent estimator of \( \omega_s^2 \). Under the null that the TVC model is correctly specified, \( \tau^* \) is asymptotically chi-square distributed with \( s \) degree of freedom. Under the alternative, the statistic would diverge.

As mentioned in Park and Hahn (1999), the fixed coefficient model would become a spurious regression if the true model is TVC model. Thus, the test statistic under the null that the fixed coefficient model is cointegrated takes the following form

\[
\tau^*_f = \frac{RSS_{FC} - RSS_{sFC}}{\hat{\omega}_s^2},
\]

where \( RSS_{FC} \) and \( RSS_{sFC} \) are the sum of the squared residuals from the regression (3) and (3) with \( s \) additional superfluous regressors, respectively. Under the null, the limit distribution is chi-square distribution with \( s \) degree of freedom, otherwise it diverges.

### 3. Empirical analysis

#### 3.1. Data description

We consider the quarterly data from 1996Q1 through 2009Q4. For the scale variable, we use real GDP. The National Bureau of Statistics (NBS) of China provides nominal GDP and its growth rates at constant prices. We also consider real industrial value added (IVA) as the scale variable in order to check the robustness of the estimation results. As for the price variable, the consumer price index (CPI) is chosen to deflate GDP, IVA, M2 and stock prices. The Shanghai composite stock index published by the Shanghai Stock Exchange is the proxy for the stock prices in our analysis. The real interest rate is obtained by subtracting expected inflation rate from one-year saving deposit rate, where the expected inflation is an average inflation rate in the previous year assuming that the agents have the belief that their best guess of inflation rate in the next year would be the one at the current year, i.e., the inflation rate follows a random walk process. The stock price index is obtained from the China Stock Market Accounting Research Database (CSMAR). Other macro data used in our analysis are obtained from the China Premium Database in the CEIC. Those databases use the raw data from the People’s Bank of China (PBC), the National Bureau of Statistics (NBS) and Shanghai Stock Exchange (SSE).

All series are plotted in Fig. 1. From Fig. 1, we can observe that real M2, GDP and IVA have increasing time trends and seasonal patterns. Thus, we consider seasonal dummies in the FC and TVC models. In Fig. 2, three different interest rate measures, seven days China Interbank Offering Rate (CHIBOR), three months CHIBOR and one year saving deposit rate (with tax adjustment) are plotted at quarterly frequency. The difference among these measures is that the one year saving deposit rate is set by the central bank, which changes infrequently and the CHIBOR rates are more market oriented interest rates which change with market conditions on a daily basis. Despite their difference, the short term CHIBOR rates follow the similar pattern as official interest rate set by the central bank and seldom deviate much. Furthermore, the interbank rate is vulnerable to short term market condition such as initial public offers in the stock market as pointed out by Hong et al. (2009). For this reason the short term interest rate is not considered in our study. From Fig. 2, we can also observe that nominal interest rate is unusually high before 2000 and decreases sharply thereafter. In Fig. 1, we can observe that the real interest rate is negative sometimes. This is due to high inflation rates since 2004. For the real stock prices we can observe that the periods 1996–2001 and 2006–2008 are two bullish periods in the

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4 In China, interest earned on savings deposits was taxed at a rate of 20 percent since November 1, 1999, which decreased to 5 percent since August 15, 2007. Our interest rate is net of this tax.

5 CSMAR is also included in Wharton Research Data Services. The macroeconomic variables used could also be found in other databases which have access to PBC, NBS and SSE in China. Moreover, the computer programs used are available upon request from the authors.
Chinese stock market, and the period 2001–2006 is a bearish period. Recent bearish periods in 2008 and 2009 are the consequence of international financial crisis.

We perform unit root tests for all series using the Augmented Dickey–Fuller (ADF) and the Kwiatkowski, Phillips, Schmidt and Shin (KPSS) tests (Dickey & Fuller, 1981; Kwiatkowski, Phillips, Schmidt, & Shin, 1992). For ADF test, Bayesian information criterion (BIC) with maximum lag length equals 10 is employed to pick the optimal lag length. For KPSS test, we use spectral GLS-detrended AR method with Modified Akaike information criterion (AIC) as criterion to choose the lag. The results are reported in Table 2. While the null hypothesis of ADF test is the time-series has a unit-root, the null of KPSS test is that the time-series is stationary. At the usual 10% significance level, Table 2 shows that there are strong evidences supporting the presence of unit root in all variables. For the robust check, we also perform unit root tests for the seasonally adjusted series and obtain the same results.

Similar to Smyth and Inder (2004), we also use Zivot and Andrews (1992) and Lumsdaine and Papell (1997) methods to test the unit root hypothesis under possible structural breaks. Both tests treat structural break dates to be endogenously determined. In these tests, one can also consider possible structural breaks in intercept and trend terms. The test proposed by Lumsdaine and Papell (1997) is more general compared to Zivot and Andrews (1992) since it takes care of the possibility of two endogenous break points.

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**Fig. 1.** Time series plot. Note: The figure plots the real M2, real GDP, real IVA and real stock price in logarithm, real interest rate and expected inflation in percentages.

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**Fig. 2.** Different interest rate measures in China.
2.1 Table 2
Unit-root tests.

<table>
<thead>
<tr>
<th>Variables</th>
<th>Demeaned series</th>
<th>Detrended series</th>
<th>T1</th>
<th>T2</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>ADF</td>
<td>KPSS</td>
<td>ADF</td>
<td>KPSS</td>
</tr>
<tr>
<td>Real M2</td>
<td>−0.94[0]</td>
<td>19.18[8]</td>
<td>−2.87[0]</td>
<td>5.91[10]</td>
</tr>
<tr>
<td>Real stock price</td>
<td>−2.43[1]</td>
<td>5.77[1]</td>
<td>−2.74[1]</td>
<td>0.60[1]</td>
</tr>
<tr>
<td>Real interest rate</td>
<td>−2.43[0]</td>
<td>5.18[0]</td>
<td>−2.95[0]</td>
<td>0.78[0]</td>
</tr>
<tr>
<td>10% critical values</td>
<td>−2.60</td>
<td>0.35</td>
<td>0.35</td>
<td>0.18</td>
</tr>
<tr>
<td>55% critical values</td>
<td>−2.92</td>
<td>0.46</td>
<td>−3.50</td>
<td>0.15</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>−4.80</td>
<td>−6.16</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td>−5.08</td>
<td>−6.75</td>
</tr>
</tbody>
</table>

Note: ADF and KPSS are, respectively, the Augmented Dicky–Fuller and Kwiatkowski, Phillips, Schmidt and Shin test statistics for the null hypothesis that the series are nonstationary or stationary, respectively. The numbers in brackets in ADF and KPSS denote the selected lag lengths. ZA and LP are, respectively, the tests proposed by Zivot and Andrews (1992) and Lumsdaine and Papell (1997). The numbers in brackets in ZA and LP denote the selected lag lengths. T1 denotes for allowing structural breaks in intercept terms and T2 allows for structural breaks in both intercept and trend terms.

2.3 points instead of only one (for technical details, refer to Zivot and Andrews (1992), Lumsdaine and Papell (1997), Ben-David, Lumsdaine, and Papell (2003).

We consider two model specifications for these tests: T1 and T2. T1 allows for possible structural breaks in the intercept term while T2 allows for breaks in both the intercept and trend terms. Following Lumsdaine and Papell (1997), we use the data-driven method to select the lag order k. We start with k = 8. If the last lag term is significant, we choose that lag order for the test. Otherwise, we reduce k by 1 and repeat the above procedure. The critical values are obtained from Zivot and Andrews (1992) and Ben-David et al. (2003).

The results are shown in Table 2. From Table 2, we observe that ZA test (Zivot & Andrews, 1992) rejects the null of unit root for real stock prices in case of T1. Under T2, ZA test rejects the null of unit root for expected inflation at 5% significant level. If we use LP test (Lumsdaine & Papell, 1997), the unit root hypothesis of real stock price is rejected for T1. However, if we consider two break points for T2, no evidence of rejecting the null of unit root hypothesis is found.

We found that the evidence of rejecting the unit root hypothesis is relatively weak. Especially, if we consider the LP test with T2, which is the most general set-up in Ben-David et al. (2003), there is no evidence against unit root hypothesis for all the series at 10% significant level. In addition to the results found by ADF and KPSS tests, these results further show the existence of unit root in all variables we considered in our analysis.

3.2. Model estimation and economic interpretation

We estimate four versions of the money demand models: real money balance determined by real GDP, the interest rate and the stock prices (S1); the model when the scale variable is replaced by real IVA (S2); the model when the opportunity cost variable is replaced by expected inflation (S3); the model when the stock prices are excluded (S4). S1 is the benchmark model, and S2 and S3 are for the robust check of our results. S4 is used to analyze the usual money demand function. For each model, we rely on BIC to choose the optimal lag truncation order of trigonometric functions in Eq. (6) when we estimate the regression Eq. (8). In practice, we choose the constant term, the linear trend and the first pair of trigonometric functions for models S1–S3, and we choose the constant term, the linear trend and the first eight pairs of trigonometric functions for model S4. Using these specifications, we first estimate β₀ and recover the time-varying coefficient αₜ using Eq. (7). The estimation results for S1 and S2–S4 are reported in Figs. 3 and 4, respectively. In these figures the estimates of coefficients are plotted in solid lines, and their 90% confidence intervals (confidence bands) are represented by dashed lines.

Before we interpret the estimation results, we check whether the time-varying cointegration regression specifications are appropriate using the test statistics τ^* and τ^**. Test statistics τ^* and τ^** are reported in Table 3. When we perform the test, the polynomial terms t, t², and t³ are considered as the superfluous regressors.

For models S1–S3, the test statistic τ^* rejects the null of FC model in favor of TVC model and, moreover, the test statistic τ^** cannot reject the TVC model at 1% significance level. Both of them imply that there exist time-varying long-run relationships among the variables. For the last model, we reject the legitimacy of both FC and TVC models.

3.2.1. Time-varying income elasticity

For the income elasticity, many previous studies report different estimates. Yi (1993) uses annual data from 1952 to 1989 and reports that the estimated income elasticities are between 0.7 and 0.9. With similar sample periods and frequencies, Hafer and Kutan (1994) obtain 1.33 for nominal M2, and Chen (1997) reports 1.93 for real M2. Using quarterly data from 1983 to 2002, Bahmani-Oskooee and Wang (2007) obtain 1.69 while Mehrotra (2008) find it to be 1.73 using the quarterly data from 1994 to 2005. With slightly longer sample periods, Wu (2009) finds the estimate to be 0.74 (1994 to 2008), and Baharumshah, Mohd, and Yol (2009) obtain 0.65 (1990 to 2005). Our estimates are between 0.60 and 0.75, which are comparable with previous studies using post-1990 data on a quarterly basis.

In addition, in Fig. 3, an increasing, although not monotonic, pattern of the income elasticity is also observed. This pattern indicates that, in recent years, given 1% growth in real income, people are likely to hold more assets falling into the category of M2. This finding is consistent with the fact that the saving ratio also shows an increasing pattern in recent years. Chamon and Prasad (2008) claim that households’ precautionary saving motives to cover the housing expenditure, education and health care become stronger in recent years, and people prefer to postpone their consumption by saving a large portion of their disposable income in the banks (around 24% in 2005).

We have to note that the above increasing pattern is not affected by changing covariates. As we can see in the first subplot of Fig. 4, the shape of the time-varying elasticity remains unchanged with slightly different magnitudes with alternative choices of other covariates. We can reach the conclusion that our result is robust to alternative choices of the scale variables.

3.2.2. Time-varying interest rate elasticity

For the interest rate elasticity, lots of studies show quite inconsistent estimation results. Hafer and Kutan (1994) show that the interest rate elasticity is 0.15 but not statistically significant. Bahmani-Oskooee and Wang (2007) obtain −1.54 but not significant neither. Baharumshah, Mohd, and Yol (2009) report that the estimated interest rate elasticity is insignificant, and they exclude it from their cointegrating regression equation. Wu (2009) reaches the similar conclusion that the interest rate plays an insignificant role in the cointegrating regression.

There are many potential reasons for the insignificance of the interest rate elasticity estimates. First, Chen (1997) points out that the major interest rates of saving deposits in China are set by the central bank directly in advance and rarely change over time. Second, as mentioned in Koivu (2009), although the centrally planned credit rationing (a predetermined and direct control of commercial banks credit plan performed by the central bank) have been abandoned in 1998, the central bank still uses a window guidance policy to influence the decision of commercial banks. Furthermore, the majority of bank loans still flows to the state-owned companies (SOEs) (Prasad & Rajan, 2006; García-Herrero et al. (2009). All these facts result in the severe restriction of the market mechanism. Lastly, as mentioned before, due to the financial depression and capital flow regulation, saving deposit is the main investment channel for domestic investors. Observing that the official interest rate does not respond to market effectively, Wu (2009) argues that interest rate does not play a significant role in affecting money holdings. Qin, Quising, He, and Liu (2005).
Fig. 4. The estimates of time-varying elasticities under model S2–S4. Note: The top, middle and bottom panel depicts the estimates of time-varying coefficients in solid lines and their corresponding 10% confidence bands in dashed lines for model S2, S3 and S4, respectively.

Table 3
Model specification tests.

<table>
<thead>
<tr>
<th>Model</th>
<th>Test statistics</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$\tau^{*}$</td>
</tr>
<tr>
<td>S1: The scale variable is real GDP</td>
<td>10.8405</td>
</tr>
<tr>
<td>S2: The scale variable is real IVA</td>
<td>3.0660</td>
</tr>
<tr>
<td>S3: The opportunity variable is expected inflation</td>
<td>5.1538</td>
</tr>
<tr>
<td>S4: The real stock price is excluded</td>
<td>21.0237</td>
</tr>
<tr>
<td>1% critical value</td>
<td>13.2767</td>
</tr>
</tbody>
</table>

Note: $\tau^{*}$ and $\tau_{1}^{*}$ are the test statistics for the null hypothesis that the variables are fixed coefficient cointegrating regression and time-varying coefficient cointegrating regression, respectively. The additional superfluous regressors are time polynomial terms, $t$, $t^2$, $t^3$ and $t^4$. If the null hypothesis is true, the corresponding statistics converges to $\chi^2_4$ in distribution. Otherwise, it will diverge as the sample size increases.

Although, traditionally, interest rate is a rather weak monetary policy instrument in China, its effect becomes stronger with the progress of the interest rate liberalization. Koivu (2009) uses more recent data and argues that the interest rate policy begins to influence the economy to a larger extent than before. More specifically, he splits the whole sample into two different sub-samples and estimates models with each sub-sample to show the changing role of the interest rate policy. Our results of the time-varying interest rate elasticity estimates support previous findings without splitting the sample. The estimated time-varying interest rate elasticities are between $-0.01$ and $0.04$ and they are significant over the recent time horizons. The interest rate elasticity estimates turn to be positive after 2004, which implies a relatively stronger role of saving deposits in M2 demand.

There are some evidences for the recent stronger effect of the interest rate policy. First, as mentioned before, the commercial banks are not restricted by the ceiling of the lending rates or the floor of the deposit rates since 2004. Second, as mentioned in Koivu (2009), the ownership structure of the commercial bank has been changing from the original state-ownership to more market oriented one. This provides more freedom to the commercial banks in the sense of issuing loans through the market mechanism. Third, with the development of the financial and housing markets, more investment alternatives are available for domestic investors in more recent years. This is necessary for the interest rate to have impacts on household saving decision. Otherwise, even if households want to reallocate their assets in response to the change of the interest rate, they may be constrained by investment opportunities.

We also consider another proxy of the opportunity cost, the quarterly expected inflation rate. As shown in Fig. 2, the fact that nominal interest rate stays low level and roughly the same over time since 2004 suggests a large portion of the variation of real interest rate comes from the changes in expected inflation rate. That is the reason why we choose the expected inflation rate as another proxy, and the similar choices have been made by Chen (1997), Mehrotra (2008), Deng and Liu (1999), among many others. Furthermore, according to economic theory, households would prefer to transfer saving deposits to real assets during high inflation periods if the nominal interest rate is unattractive. Thus, the corresponding coefficient should be negative.

The estimates of the time-varying elasticities have been depicted in the second panel of Fig. 4. We could draw at least two conclusions under this model specification. First, the shape of the income and stock price elasticities remain unchanged but the values of the income and stock price elasticities are larger and smaller (in absolute value) compared with those of $S1$. Second, similar to the case of real saving rate, the elasticity of expected inflation remains insignificant over a large portion of the time horizon. The negative sign after 2004 is not surprising as we mentioned the above.

3.2.3. Time-varying real stock price elasticity

Traditionally, stock market prices have not been considered as a determinant of the money demand until Friedman (1988). According to Friedman (1988), the stock prices may exert a positive wealth effect and a negative substitution effect. There are mainly three channels contributing to a positive wealth effect: (i) an increase in the stock prices would yield an increase in nominal wealth. This results in a positive effect on the money demand to facilitate consumption; (ii) the better the condition of the stock market is, the more money is needed in order to facilitate transactions; (iii) since higher stock prices imply higher future expected returns and, in turn, higher risks (assuming investors’ preference of risk to be constant), investors are willing to shift the large proportion of their wealth to risk free ones such as cashes or saving deposits. The substitution effect works exactly in the opposite direction. As the stock prices rise, equity would become more desirable, and therefore, the demand for money decreases. The whole effect depends on the relative magnitude of both effects. Choudhry (1996) studies the relationship between the stock prices and money demand (real M2) in the U.S. and Canada from 1955 to 1989 using Johansen’s cointegration method and finds that the stock price elasticities are negative and positive in Canada and U.S., respectively. He further argues that there is no strong

Figure 5: Stock price and total number of investors. Note: This figure plots the Shanghai composite stock price (left scale) and number of investor accounts in China’s stock market (in 10 thousands, right scale).

evidence supporting a long-run stable money demand function without considering real stock prices. Caruso (2006) shows that
the wealth effect dominates from 1913 to 1980 while the substitution effect dominates in the last two decades in Italy. Baharumshah, Mohd, and Yol (2009) estimate a positive stock price elasticity of 0.287 in China.

In order to get some general idea, the stock prices and the number of investors' accounts in China are plotted in Fig. 5. In Fig. 5, we can observe that more investors entered the stock market during the bullish periods of the stock market while the number of investors' accounts was non-increasing during the bearish market periods. Thus we can say that the third channel of the wealth effect may not be dominant in China. The relative strength of the wealth effect may mainly depend on the other two channels.

As we mentioned the above, since a large amount of saving deposits enjoys quite low real interest rate, and financial markets are not well developed in China, the stock market can be the natural alternative choice. As pointed out by Wu (2009), the proportion of equity assets in the total financial assets increased from 5% in 2005 to above 20% in 2007. The shift of the saving deposit and equity shares can be inferred from Fig. 6. We can observe that during two bullish market periods, 1996–2001 and 2006–2008, the growth of the saving deposits increases moderately or even decreases sharply around 2007. On the contrary, during the bearish periods, the growth of the saving deposits increases tremendously. Since the money in the stock accounts is not classified as the broad money (M2), a massive amount of money transferring from the banks to the stock market constitutes a strong substitution effect for the demand for M2. This effect is particularly strong during the bullish market. However, for the

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6 Year 2007 is a year of exceptionally high stock returns. See Monetary Policy Report, 2007 and other issues published by PBC.

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bearish market periods, we cannot observe some specific trends. In Fig. 7, we find that when the stock prices reach its peak, the stock price elasticities are more likely to be negative and, moreover, reach its bottom. For the bearish market, the stock price elasticities are positive or insignificant.

To further investigate the role of the stock prices, we estimate both the TVC and FC models excluding real stock prices in the regression equation. The result of TVC model is shown in the last subplot of Fig. 4. This graph shows rather wriggled estimates of elasticities. Unfortunately, model specification tests reject both TVC and FC models which implies that the exclusion of the stock prices from the money demand function may yield model misspecification. This implies that the long-run equilibrium between money demand and its determinants cannot be captured by either the fixed coefficient approach or the time-varying one without including the stock prices in the regression equation.

4. Concluding remarks

Many previous studies investigate the long-run equilibrium of the money demand using traditional cointegrating regression approach. However, the fixed coefficient approach may fail to capture the long-run relationship when economic condition and policy regimes are changing over time. This is especially the case in a transition economy like China, where smooth structural changes are present. As a result, the usage of traditional (fixed coefficient) parameter cointegrating regression approach may not be appropriate. In order to study the long-run relationship of the money demand function in China, we analyze the money demand in China using the smooth time-varying cointegrating approach and find the existence of long-run time-varying stable relationship.

Using recent data set, we find (i) the estimates of income elasticities are around 0.6–0.75, which are comparable with existing studies; (ii) our interest rate elasticity estimates are between −0.01 and 0.04. This finding is consistent with the fact that the role of the interest rate policy is weak in China and households' insensitivity to monetary policy changes, although there are some mild evidences that its role has been strengthened in recent years; (iii) considering the stock prices as an additional covariate in the money demand equation helps to explain the demand of real money balances. We observe that the substitutional effect dominates the wealth effect, especially, during the bullish market period.

We identify and highlight the role of the stock prices in the money demand. The strong substitution effect of equity assets is the result of underdeveloped financial market, unattractive real interest rate and high saving ratio. Even in a transition economy with immature financial market like China, asset price is one of important determinants in money demand analysis. More efforts should be devoted to explore how this phenomenon would affect the monetary policy in China.

References


