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Aaron Mehrotra

Demand for money in transition: Evidence from China's disinflation

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Tiivistelmä

Tutkimuksessa tarkastellaan rahan kysyntää Kiinan taloudessa. Tutkimusajanjaksolla Kiinan talous kävi läpi merkittävän disinflaation ja lyhyen deflaation, mutta kokonaistuotannon kasvu oli vahvaa. Lavean rahan (M2) kysynnälle estimoitu malli on vakaa. Mallin tasapaino saavutetaan osaksi inflaation kautta, ja impulssivasteanalyysi osoittaa rahasokkien johtavan voimakkaampaan inflaatioon. Tutkimuksessa ei löydetty viitteitä rahan kysynnän epälineaarisuudesta disinflaation oloissa. Tulosten voidaan katsoa puoltavan Kiinan keskuspankin strategiaa määritellä tavoitearvoja rahan määrän kasvulle, mutta nimellisen efektiivisen valuuttakurssin muutosten ottaminen huomioon on tärkeää strategian onnistuneen toteuttamisen kannalta.

Asiasanat: rahan kysyntä, disinflaatio, deflaatio, Kiina

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Demand for Money in Transition: Evidence from China's Disinflation*

Aaron Mehrotra

August 8, 2006

Abstract

We examine money demand in the Chinese economy during a period characterized by significant disinflation and outright deflation, coupled with strong output growth. Our study establishes a stable money demand system for broad money M2. Inflation affects the adjustment of the system towards equilibrium, and shocks to broad money are found to lead to higher inflation in the context of an impulse response analysis. No evidence of non-linearity in money demand is found for the disinflationary period. The results provide support for the PBoC's policy of specifying intermediate targets for money growth. Importantly, our results suggest that movements in the nominal effective exchange rate should be taken into account in a successful implementation of such a policy.

Keywords: Money demand; Disinflation; Deflation; China
JEL Classification: E31, E41

1 Introduction

Since the mid-1990s, the People's Bank of China (PBoC) has emphasized the roles of money growth and credit targets in formulating monetary policy. Given banking's dominance in the financial system, the emphasis on money supply and credit growth is obvious. Indeed, the well-publicized flow of foreign investment into the Chinese banking sector during 2005 and the persistent weakness of local stock markets highlighted the fact that the importance of the banking sector is unlikely to wane anytime soon. However, a prerequisite for meaningful monitoring or targeting of monetary aggregates is a stable relationship of

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the money stock to real output and/or inflation, i.e. the indicators generally accorded the greatest policy interest. Similarly, as changes in monetary stance may affect the economy with considerable lags, a formal investigation of the dynamics between money and other macroeconomic variables may help in crafting a successful monetary stabilization policy.

The paper examines the demand for domestic broad money M2 in Mainland China during the period 1994-2005, focusing on the possible existence of a stable money demand function as the economy went through a notable process of disinflation (and even deflation). Despite strong growth in real GDP (an average of over 9% during 1994-2004), consumer price inflation slid from an annual rate of over 20% in 1994 to negative values in 1998-2002.[1] The stability of a money demand system during such a time period cannot be taken for granted, given the speed of ongoing structural change from a command to a market economy, and the possibility that deflation may have caused some instability in macroeconomic relationships.

Moreover, the potential inflationary consequences of broad money shocks are of interest here in light of China's surprisingly persistent low inflation in recent years. With interest rates yet to assume a role similar to that of more developed financial systems, monetary targets defined in terms of growth rates of broad money M2 have occupied an important place in the formulation of policy. In contrast, central banks of most developed economies have diminished the policy significance of monetary aggregates (with the prominent exception of the European Central Bank). Finally, we are interested in the role of the exchange rate in the demand for Chinese broad money. In a transition economy with restrictions on capital flows and imposed ceilings on domestic residents' holdings of foreign currency, traditional currency substitution effects may be limited. However, expectations of domestic currency appreciation may still increase capital inflows (unauthorized and authorized) and thereby increase the domestic money stock.

We find a stable and an economically meaningful money demand relationship in a cointegration framework for the Chinese economy. In line with earlier research on Chinese money demand functions and results from developing financial systems, our estimated income elasticity is significantly greater than one. Notably, the nominal effective exchange rate is found to be a statistically significant determinant of money demand. No evidence of non-linearity in money demand is observed for the disinflationary period. Excess liquidity appears to lead to both higher output growth and consumer price inflation. Similarly, analysis of model dynamics in the context of a structural vector error correction (SVEC) system with contemporaneous restrictions suggests that broad money shocks have inflationary consequences. As such, our findings lend support to the policy of closely monitoring the developments of broad money M2 by the

[1] According to data prior to the GDP revision of 2005.

PBoC to pursue monetary stabilization. Further, it appears that movements in the nominal effective exchange rate should not be ignored when analyzing Chinese monetary developments.

Most of the work on Chinese money demand functions focus on the period before economic reforms or on the reform period prior to the emergence of low inflation. Chow (1987) is generally credited for pioneering work on Chinese money demand. He used a money demand specification based on the quantity theory for 1952-1983. Portes and Santorum (1987) also estimate both real and nominal money demand systems for M0 during 1954-1983, while Girardin (1996) focuses on the demand for currency during 1988-1993. Huang (1994) estimates an error correction model during the reform period of 1979-1990. Xu (1998) uses a disaggregate approach to Chinese money demand in order to study three components of M2: currency, personal deposits and institutional deposits during 1980-1996.

Surprisingly few studies examine money demand in the Chinese economy during the disinflationary and deflationary period. Yet, given that inflation and deflation are fundamentally monetary phenomena, the behavior of money demand and the dynamics between the money stock and prices are of obvious interest. Moreover, to our knowledge the possible effects of the exchange rate on money demand in the Chinese case have not been investigated. Gerlach and Kong (2005) establish a stable money demand system for M2 for 1980-2004. Using a trivariate system of real income, the consumer price index and the money stock, the authors find a long-run income elasticity comparable to previous findings on Chinese money demand functions. A measure of money overhang (defined as the difference between observed and equilibrium money stock) was found to provide leading information about future inflation. The period of disinflation is also included in the study of Cargill and Parker (2004), who examine the possible deflation-induced discontinuity in monetary policy. Comparing estimates for money demand equations for the US, Japan and China, the authors find no asymmetry in Chinese money demand that could be attributed to deflation, but add that their money demand equations did not give a good fit in the Chinese case. In contrast to our approach, which uses a systems framework with a cointegration relationship, Cargill and Parker (2004) use a single equation approach, providing estimates from both level and difference-form equations.

The remainder of this paper is structured as follows. The next section discusses the use of monetary aggregates in the formulation of monetary policy in China, followed by a presentation of theoretical issues pertinent to our research question and information on the time series used. We estimate the empirical model in Section 4 and discuss the results. The final section concludes.

2 Some Observations on Policy

Before the start of economic and financial reforms in the late 1970s, macroeconomic policy-making was the sole responsibility of the State Council. Hafer and Kutan (1993) argue “currency in circulation” was the relevant concept at the time to Chinese money demand. The People’s Bank of China was only founded in 1983 to conduct monetary policy, with its objective later defined as maintaining the stability of the value of the currency and thereby the promotion of economic growth. Monetary aggregates were adopted as intermediate targets in 1996. The basic objective of the operation of policy in recent years has been to maintain an appropriate growth in money and credit (Yu and Ming, 2001; Dai, 2002). The PBoC has claimed that such a rate of money growth will promote "economic growth positively and contribute to preventing both inflation and deflation" (PBoC, 2005).

For our purposes, these considerations justify the use of broad money M2 (i.e. currency in circulation, demand, time and savings deposits) as the monetary aggregate of interest. Considering only M0 in an analysis of Chinese monetary developments potentially leaves out information on inflationary pressures,[2] since relatively liquid household and enterprise deposits represent an important part of Chinese financial wealth. Indeed, China’s broad money-to-GDP ratio is among the highest in the world, partly reflecting the underdevelopment of bond markets and a relatively weak performance of stock markets. The strikingly high ratio of 184% in 2003 was matched only by the financial centre of Hong Kong (265%) and Lebanon (358%). Moreover, the Chinese M2-to-GDP ratio increased for most of the estimation sample (Fig. 1).

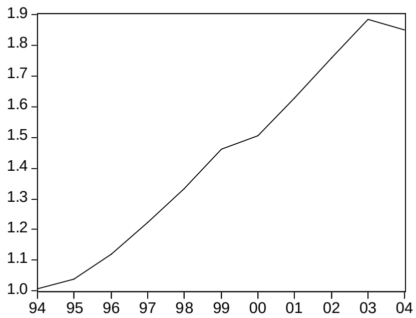

Figure 1. M2-to-GDP ratio

Annual targets for the growth rates for M2 have been explicitly spelled out by the PBoC. In Table 1, we list the annual (unrevised) growth targets set for

[2] It may also be preferable to examine an aggregate that internalizes possible portfolio shifts among components included in broad money.

M2 by the PBoC, together with actual outcomes for 1999-2005. The targets for 1999 and 2000 were specified as ranges rather than point targets.

Year	Target	Actual outcome
1999	14-15%	14.7%
2000	14-15%	12.3%
2001	14%	14.4%
2002	13%	16.8%
2003	16%	20.0%
2004	17%	14.7%
2005	15%	17.6%

Table 1. Money growth targets and actual outcomes.
Source: PBoC Annual Bulletin & Quarterly Statistical Bulletin, various issues.

While money growth was roughly on target in 1999, the deflationary climate of 2000 saw the midpoint of the target range of 14-15 percent undershot by over two percentage points. In 2001, the targeted growth in broad money supply remained in line with the PBoC's goal. The pace of money growth picked up in 2002 and 2003, but declined in 2004 as measures to prevent overheating in the Chinese economy were taken. In sum, for a developing economy, money growth has not deviated excessively from the announced targets. Moreover, policy has accommodated impressive GDP growth rates without leading to runaway inflation. Indeed, low inflation seems to have become a well entrenched feature of the Chinese economy. However, considering the fact that China went through a period a deflation, perhaps even higher money growth targets may have been appropriate in the late 1990s and early 2000s. Indeed, when deflation turned to positive inflation in 2002, actual money growth exceeded the announced target by almost four percentage points.

3 Theoretical Considerations and Data Issues

3.1 Specification of the Money Demand Relationship

We analyze the demand for broad money M2 based on a general long-run money demand specification of the following form:

$$(m-p)_t = \beta_1 y_t + \beta_2 \pi_t^e + \beta_3 neer_t + v_t. \quad (1)$$

In (1), real balances are specified as a function of real income y, the expected inflation rate π^e, and the nominal effective exchange rate $neer$. Small letters here and elsewhere in the text denote logarithms. Data sources and definitions are specified in Table 2 below, with the series depicted in Figure 2. We use quarterly data. Variables for real output, money, and consumer prices are seasonally adjusted using the TRAMO-SEATS procedure for seasonal adjustment.[3]

[3] EViews 5.1 and JMulTi 4.06 software were used in the estimation procedure.

Variable	Data Series	Source
m	Nominal money stock M2	OECD Main Economic Indicators (MEI)
p	Consumer price index	International Financial Statistics (IFS)
		China Monthly Economic Indicators
		MEI, own calculations
y	Real GDP	MEI, IFS, own calculations
$neer$	Nominal effective exchange rate	IFS

Table 2. The data.

Our estimation sample for assessing the demand for Chinese broad money commences in 1994. Not only is there little reason to justify a much earlier starting date for drawing implications for current policy due to rapid structural change, but the start date of our estimation sample coincides with a major shift in Chinese macroeconomic policy. The previous dual-exchange rate system ended with the introduction of a managed-float regime in January 1994. A single unified rate for all authorized foreign exchange transactions was set against the US dollar with a narrow band of $\pm 0.25\%$ (Huang and Wang, 2004). We now consider each system variable in turn.

As is conventional, we model money demand as the demand for real money balances $(m - p)$, with the latter obtained by deflating the series for nominal money by the consumer price index.[4] Major price liberalization measures were introduced before the start of our estimation sample, further justifying the use of the consumer price index and the chosen sample range.[5] Even if some commodity prices were liberalized in the early 1980s, price liberalization was still an important issue in 1991 and 1992, when the authorities further liberalized commodity prices and reduced administered prices in agriculture and industrial sectors (Gerlach and Kong, 2005).

[4] We construct the consumer price index here by setting 2003M1=100, using month-on-month growth rates for 2003 and year-on-year growth rates otherwise. The quarterly series is obtained as an average of monthly values. The GDP deflator is not available for China on a quarterly basis.

[5] The retail price index was used in the study by Girardin (1996). However, inflation rates defined in terms of the retail price index correspond closely to those based on the consumer price index. For example, during the period from the emergence of deflation (defined by a negative y-o-y growth rate in the consumer price index) to the end of our estimation sample, the correlation coefficient between the two series amounts to 0.96.

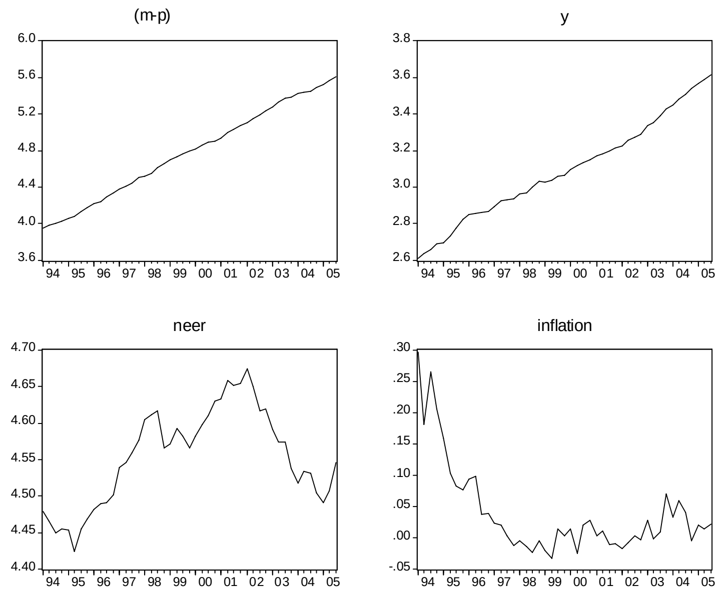

Figure 2. The series

We use real GDP as our measure of real income, with the expected sign for the coefficient $\beta_1 \succ 0$.[6] The appropriateness of this output variable could be questioned – not least due to quality concerns about Chinese GDP data (see e.g. Rawski, 2001).[7] Moreover, in December 2005 China's National Bureau of Statistics announced an upward revision of the level of GDP by 17%, mostly to correct an underestimation of the importance of the service sector in the Chinese economy. While this study relies on pre-revision GDP statistics, a comparison of the revised annual GDP growth rates with the formerly used series does not suggest significant changes in the growth dynamics of real GDP. Nevertheless, as revised quarterly statistics become available, it will be worthwhile to estimate money demand systems with the new data.

The inflation rate, defined as $\pi_t \equiv 4 * (p_t - p_{t-1})$, is used as the opportunity cost of holding money. Chinese financial markets remained relatively undeveloped during our estimation sample and there is a lack of a proper market-based interest rate. This grants the expected sign $\beta_2 \prec 0$. Of course, it is the expected

[6] Real output is obtained using the quarterly series on nominal GDP deflated by the consumer price index from 1995 onwards. As no quarterly level data is available, the observations for 1994 are obtained using annual growth rates.

[7] The importance of money stock data in providing information about the true level of output when the policymaker only observes a noisy measure of economic activity has been discussed by Masuch et al. (2003). This could also be important when the role of unrecorded activity (e.g. through the underground economy, or because statistical procedures remain underdeveloped) in the economy is large.

inflation rate, rather than the actual outcome, that properly represents the opportunity cost for holding money instead of goods. As applied by Chen (1997), if rational expectations are assumed, realized inflation differs from the expected rate only by an unpredictable error term. This error term is assumed to be stationary, so realized inflation can be included in a straightforward fashion in the long-run money demand relationship.

The sign of the coefficient β_3 on the nominal effective exchange rate is ambiguous a priori. Arango and Nadiri (1981) suggest that a domestic exchange rate appreciation (increase in *neer*) lowers the domestic currency value of foreign assets, decreasing domestic wealth and thereby lowering the demand for real money balances. This would suggest a sign $\beta_3 \prec 0$. Alternatively, Bahmani-Oskooee and Pourheydarian (1990) argue that if a domestic currency appreciation leads to expectations of further appreciation, currency substitution in favor of domestic assets could be induced. This yields an expected sign $\beta_3 \succ 0$. The same expected sign would be obtained by invoking a rational expectations hypothesis similar to the case of the inflation rate discussed above. In the Chinese case, restrictions on domestic holdings of foreign currency make it likely that the effect of the exchange rate on money demand works through expected movements in the Chinese renminbi and the consequent (perhaps unauthorized) currency inflows that may add to the domestic money stock. Prasad and Wei (2005) mention that China experienced a significant increase in non-FDI capital inflows during 2001-2004. These years also saw an intense debate about the possible undervaluation of the renminbi. Similarly, Gunter (2004) provides estimates of non-negligible capital flight for 1994-2001, especially during 1997-2000. It is also plausible that Chinese firms report and convert their foreign exchange earnings to renminbi depending on their expectations of future appreciation, thereby affecting the domestic money supply. The currency substitution phenomenon in China therefore differs from what has occurred in the transition economies of Eastern Europe, Russia and some countries in Latin America, where foreign currency has been used as a hedge against high domestic inflation. The latter has not presented a problem for the Chinese policymakers during the low inflation environment of our study. The use of the nominal effective exchange rate can be defended as the Chinese currency remained de facto pegged to the US dollar throughout our estimation sample.[8] Additionally, the effective exchange rate allows us to incorporate considerations of currency substitution between the Chinese renminbi and currencies other than the US dollar.

A final observation relates to the usefulness of focusing on the money stock in the conduct of policy in an environment of falling prices. When the zero floor on interest rates is hit, the interest in the money stock in the conduct of

[8] The effective exchange rate has been used in money demand studies covering the US, Spain, Canada and Japan by McNown and Wallace (1992); Bahmani-Oskooee et al. (1998); Bahmani-Oskooee and Pourheydarian (1990), among others. As in our study, all these studies utilize the level of the exchange rate. Buch (1998) uses the level of the dollar/ruble exchange rate for the transition economy of Russia.

policy is revived and the Keynesian model with the short-term interest rate as a comprehensive indicator of the monetary policy stance gives way to a more monetarist approach. Burdekin and Siklos (2005) provide reaction function estimates for the PBoC utilizing a McCallum rule (see McCallum, 1988), where the central bank uses the monetary base as its policy instrument instead of the short-term interest rate. Admittedly, interest rates in the Chinese economy have remained far from the zero floor and at a relatively high level considering the substantial disinflation during our estimation sample. During April-May 1999, for example, when deflation was strongest, the administratively set one-year lending rate still amounted to 6.39! However, given that little inflationary pressure in the Chinese economy exists and the average nominal interest rate should decline with the average rate of inflation, a binding zero bound in the Chinese economy is not inconceivable.

3.2 Time Series Properties of the Series

Augmented Dickey-Fuller (ADF) tests are used to investigate the order of integration of the series in our system. We follow the Pantula principle in the testing procedure, whereby the series is first differenced as many times as necessary to make it stationary, assuming a maximum order of integration of 2. In Table 3, we display the results for the series in levels and first differences. Akaike information criteria are used to determine the lag length, allowing a maximum lag length of p_{max} =8 in the test.

Series	Det. term	Lagged differences	Test stat.
$\Delta(m-p)$	constant	0	-5.95***
$(m-p)$	constant, trend	0	-2.18
$\Delta\pi$	constant	4	-3.63***
π	constant, trend	1	-2.17
Δy	constant	7	-2.59*
y	constant, trend	8	-0.19
$\Delta neer$	constant	6	-2.81*
$neer$	constant, trend	7	-0.44

Table 3. Augmented Dickey-Fuller test for unit roots.

* indicates significance at 10% level, ** at 5% and *** at 1% level.

For the rate of inflation and the real money stock, a unit root cannot be rejected for the series in levels, while tests on the first-differenced series strongly suggest stationarity. In the case of real GDP and nominal effective exchange rate in first differences, the rejection of a unit root at a 10% level only may indicate a power problem due to a high number of lags included in the test. Indeed, using the lag length of zero as suggested by the Schwarz criterion causes a rejection of a unit root even at 1% level for both series. The lag length used in the test for these series in levels is high as well, but here the Schwarz information criterion does not lead to a rejection of a unit root. We therefore continue with the assumption that all the series in levels are integrated of order one. This order of integration makes it possible that the series share common stochastic trends, i.e. cointegrating relationships. We now move on to examine the cointegrating rank of our system.

The examination of the cointegrating rank is achieved with the popular Johansen trace and maximum eigenvalue tests (see e.g. Johansen, 1991, 1995). As the series display trending behavior, it is preferable to include both constant and trend as the deterministic terms in the testing procedure. Results from the Johansen trace and maximum eigenvalue tests for our full system are displayed below in Table 4.

Test	Lagged levels	Null hypothesis	Test value
Johansen (trace)	1	$r=0$	71.11***
		$r=1$	34.60
		$r=2$	16.52
		$r=3$	3.72
Johansen (maximum eigenvalue)	1	$r=0$	36.51**
		$r=1$	18.08
		$r=2$	12.80
		$r=3$	3.72

Table 4. Cointegration tests.

* indicates significance at 10% level, ** at 5% and *** at 1% level.

Constant and trend used as deterministic terms in all cases.

Both Johansen trace and maximum eigenvalue tests suggest a cointegrating rank of one, as a cointegrating rank of zero can be rejected, and a cointegrating rank of one cannot be rejected at conventional levels of significance. We use the lag length indicated by the Schwarz criterion for the trace test, as both Akaike and Hannan-Quinn criteria suggest unreasonably high lag lengths for our sample (8 lags in all cases). Due to limitations posed by the sample size, trace test results using the nonstandard Osterwald-Lenum (1992) critical values are examined. They confirm our previous findings; the trace test again suggests a rejection of a cointegration rank of zero at 1% level, while a cointegration rank of one cannot be rejected at any conventional significance level.

Given the relatively low power of cointegration tests in our system of four variables, subset tests may also be of interest. Moreover, a normalization that leads to the interpretation of the long-run relationship as a plausible money demand relation is problematic if money does not actually enter the cointegration relation. Three out of four tests on trivariate subsystems in the Johansen framework indicate the existence of a cointegrating rank of one, while one subsystem suggests the possibility of even a second cointegration relation. In conclusion, we proceed to the model estimation assuming a cointegrating rank of one between the variables.

4 Empirical Evidence

4.1 Estimation of the VEC Model

A reduced form representation of a vector error correction model, omitting the deterministic terms, can be written as:

$$\Delta x_t = \Pi x_{t-1} + \Gamma_1 \Delta x_{t-1} + ... + \Gamma_{p-1} \Delta x_{t-p+1} + u_t, \qquad (2)$$

where p denotes the order of the VAR-model. K being the number of variables, $x_t = (x_{1t}, ..., x_{Kt})'$ is a $(K \times 1)$ random vector and Γ_1 are fixed $(K \times K)$ coefficient matrices. The $u_t = (u_{1t}, ..., u_{Kt})'$ is a K-dimensional white noise process with $E(u_t) = 0$. When the variables are cointegrated, Π has reduced rank $r = rk(\Pi) < K$. It can be written as $\Pi = \alpha\beta'$, where α and β are $(K \times r)$ matrices that contain the loading coefficients and the cointegration vectors, respectively.

Our vector of endogenous variables is written as $x_t = ((m-p)_t, y_t, \pi_t, neer_t)'$, with an unrestricted constant included as the deterministic term.[9] The estimation sample runs from 1994Q1 to 2005Q3. We estimate the model with three

[9] We do not restrict a trend in the cointegration relation, following the argument by Brüggemann and Lütkepohl (2005a) that it is difficult to theoretically motivate the inclusion of such a deterministic variable in a money demand relation. This approach is also consistent with previous studies on Chinese money demand.

lags, corresponding to four lags in a levels-form VAR (one year of quarterly data). Misspecification tests provide support for our choice of the autoregressive order (see discussion below). Since the observations 1994Q1-1994Q4 are used as presample values, the actual sample size amounts to T=43. The Johansen maximum likelihood (ML) procedure yields the following estimate for the long-run relationship (standard errors in parenthesis):

$$\begin{aligned}(m-p)_t \quad &= \quad 1.728y_t + 0.967\pi_t + 1.181neer_t + ec_t. \\ &\quad\;\; (0.026) \quad\; (0.137) \qquad (0.098)\end{aligned} \tag{3}$$

The normalization of the coefficient on real money balances to one leads to the representation of (3) as a possible money demand relationship. Note that income elasticity is significantly higher than one (1.728). This is consistent with earlier findings for China (e.g. Chen, 1997; Chow, 1987; Gerlach and Kong, 2005). The finding of money demand increasing faster than income is a reasonable result in developing economies and could arise as a result of monetization, whereby money is used more intensively in the settlement of transactions, or through the development of the commercial banking system. Both seem plausible explanations in the context of the Chinese transition, and are in line with the observed increase in broad money to GDP ratio depicted earlier in Figure 1.[10] The opportunity cost variable actually obtains the wrong sign, which is in line with some of the previous results for China (see Gerlach and Kong, 2005). This could reflect the lack of alternative investment opportunities and a huge desire to save by the Chinese money-holding sector, in which case broadly defined money stock is demanded regardless of higher inflation expectations. However, it could also follow from the successful disinflationary process by the Chinese policymakers, whereby inflation no longer enters the money demand function of the economic agents in a way predicted by conventional theoretical assumptions. Assuming that actual movements reflect expectations of the future direction of the exchange rate, the expectation of renminbi appreciation increases the demand for the Chinese currency, possibly through increased currency inflows that add to the domestic money supply. Alternatively, exporting firms decide on the timing of the conversion of their export earnings (from foreign currency to renminbi) based on their expectations of currency appreciation. Our result here is in line with the currency substitution hypothesis. It suggests that movements in the exchange rate ought to be closely monitored in the conduct of policy that uses money supply as an intermediate target.[11]

Brüggemann and Lütkepohl (2005b) point out that ML estimation occasionally produces implausible cointegration parameters due to its potentially

[10] For discussion about the institutional reasons for a downward trend in velocity, see Bordo et al. (1997).

[11] The result of an actual exchange rate depreciation leading to lower money demand seems to fit poorly with the recent empirical evidence. The *neer* fell together with the US dollar from early 2002 to 2005, and there are estimates of quite significant capital inflows to China at that time (see Prasad and Wei, 2005). However, Gunter (2004) provides high estimates of capital flight during some earlier years in our estimation sample when the *neer* fell. These include 1994, 1995, 1998 and 1999, which is consistent with our result.

poor small sample properties. They suggest that the two-step generalized least squares estimator proposed by Ahn and Reinsel (1990) may not suffer from this problem. In our case, the two-step estimator produced very similar estimates to those presented in (3); the income elasticity now amounts to 1.713 and the exchange rate elasticity to 1.190. The only material difference is the somewhat lower coefficient on the inflation rate, which the two-step procedure estimates at 0.665.

We conduct a model reduction procedure, whereby the parameter with the lowest t-value is checked and possibly eliminated from the system, until a threshold t-value of 1.00 has been reached. This grants a more parsimonious specification, which may be preferable in a reduced form VEC system with many statistically insignificant coefficients. The individual equations of the VEC system are estimated with GLS as follows (standard errors in parenthesis):

$$\begin{aligned}\Delta(m-p)_t = &-\underset{(0.032)}{0.088}\, ec_{t-1} + \underset{(0.137)}{0.145}\,\Delta(m-p)_{t-1} - \underset{(0.056)}{0.059}\,\Delta\pi_{t-1} \\ &- \underset{(0.084)}{0.218}\,\Delta neer_{t-1} + \underset{(0.128)}{0.170}\,\Delta(m-p)_{t-3} - \underset{(0.118)}{0.234}\,\Delta y_{t-3} - \underset{(0.195)}{0.497} + u_{1t}\end{aligned} \tag{4}$$

$$\begin{aligned}\Delta y_t = &\underset{(0.032)}{0.162}\, ec_{t-1} + \underset{(0.081)}{0.190}\,\Delta neer_{t-1} - \underset{(0.123)}{0.349}\,\Delta(m-p)_{t-2} \\ &+ \underset{(0.110)}{0.150}\,\Delta y_{t-2} - \underset{(0.057)}{0.142}\,\Delta\pi_{t-2} - \underset{(0.122)}{0.152}\,\Delta(m-p)_{t-3} + \underset{(0.111)}{0.178}\,\Delta y_{t-3} \\ &- \underset{(0.049)}{0.130}\,\Delta\pi_{t-3} - \underset{(0.077)}{0.090}\,\Delta neer_{t-3} + \underset{(0.191)}{0.998} + u_{2t}\end{aligned} \tag{5}$$

$$\begin{aligned}\Delta\pi_t = &\underset{(0.071)}{0.519}\, ec_{t-1} + \underset{(0.229)}{0.383}\,\Delta(m-p)_{t-1} + \underset{(0.211)}{0.777}\,\Delta y_{t-1} - \underset{(0.100)}{0.178}\,\Delta\pi_{t-1} \\ &+ \underset{(0.151)}{0.658}\,\Delta neer_{t-1} + \underset{(0.230)}{0.284}\,\Delta(m-p)_{t-2} + \underset{(0.204)}{0.240}\,\Delta y_{t-2} + \underset{(0.152)}{0.613}\,\Delta neer_{t-2} \\ &+ \underset{(0.197)}{0.710}\,\Delta y_{t-3} + \underset{(0.153)}{0.351}\,\Delta neer_{t-3} + \underset{(0.425)}{3.033} + u_{3t}\end{aligned} \tag{6}$$

$$\begin{aligned}\Delta neer_t = &-\underset{(0.250)}{0.287}\,\Delta(m-p)_{t-1} - \underset{(0.104)}{0.140}\,\Delta\pi_{t-1} + \underset{(0.147)}{0.183}\,\Delta neer_{t-1} - \underset{(0.108)}{0.134}\,\Delta\pi_{t-2} \\ &- \underset{(0.146)}{0.199}\,\Delta neer_{t-2} - \underset{(0.090)}{0.139}\,\Delta\pi_{t-3} + \underset{(0.149)}{0.323}\,\Delta neer_{t-3} + \underset{(0.009)}{0.010} + u_{4t}.\end{aligned} \tag{7}$$

The loading coefficients, which represent the weights of the cointegration relation (denoted by ec_{t-1}) in the respective equations, provide information about the possible adjustment to a long-run equilibrium. The adjustment coefficient for Eq. (4) for real money balances, amounting to -0.088, is statistically significant and negative, suggesting a stable model. Notably, excess liquidity in our system leads to both higher real GDP and inflation, with adjustment coefficients 0.162 and 0.519 in Eqs. (5) and (6), respectively. The significance of the adjustment coefficient for the inflation equation justifies a close monitoring of the broad money stock in order to obtain information about inflationary pressures. As the nominal effective exchange rate was entirely determined outside of China during the period of the dollar peg, it is not surprising that excess real money does not lead to an adjustment through the effective exchange rate towards equilibrium. Hence, the exchange rate in our system is weakly exogenous. However, note that the nominal effective exchange rate enters significantly the equations for money, inflation, and output.

The adequacy of our model, represented by Eqs. (4)-(7) above, was tested using various misspecification tests. These included the Portmanteau and Breusch-Godfrey test to detect residual autocorrelation, the Jarque-Bera test for nonnormality in the four individual equations, and the ARCH-LM test to examine possible ARCH-effects in the residuals. None of the misspecification tests raise concerns about the estimated model.[12] The rejection of nonnormality is interesting, as possible nonlinearity in the system is sometimes detected in the form of nonnormal residuals. This could be a relevant issue for money demand in an environment where inflation rates have turned negative.

Testing for the stability of our estimated system is crucial, especially as the PBoC specifies intermediate money growth targets in the conduct of policy. Such targets are meaningful only if the money demand relationship displays stability through time. Naturally, our short estimation sample imposes strict limits on the tests that can be conducted as well as on the actual testing sample. We conduct the Chow forecast test, utilizing bootstrapped p-values proposed by Candelon and Lütkepohl (2001), who show that conventional Chow test statistics reject the null hypothesis far too often in samples of a realistic size. We test for model stability in each datapoint commencing in 2000Q3 and ending in 2005Q2, which is the longest feasible sample in our case. Using 1,000 bootstrapping replications, the Chow forecast test gives no indication of instability in our system. The lowest obtained p-values for the test statistic are 0.49 and 0.53 for 2004Q4 and 2003Q1, respectively. Indeed, most p-values are of a magnitude higher than 0.70. Detailed results are available from the author upon request.

We next consider recursive eigenvalue tests for our VEC system with cointegrated variables. Under these tests, proposed by Hansen and Johansen (1999),

[12] Specifically, the adjusted Portmanteau test statistic to detect autocorrelation in the model residuals, utilizing 16 lags, amounts to 223.12 (with a p-value of 0.50). The Breusch-Godfrey test for autocorrelation in lower lag orders yields an LM-test statistic of 89.52 (p-value 0.22) at five lags, while at four lags the statistic obtains a value of 70.31 (p-value 0.27) and at one lag 17.20 (p-value 0.37). The Jarque-Bera test for nonnormality in the four individual equations of the model yield test statistics of 0.75, 1.44, 3.02, 1.63, respectively (with the corresponding p-values 0.69, 0.49, 0.22 and 0.44). Finally, ARCH-LM tests for possible ARCH-effects in the model residuals, conducted at 16 lags for each of the four equations, yield test statistics of 15.64, 7.40, 8.00 and 12.79 (with the corresponding p-values 0.48, 0.96, 0.95 and 0.69).

a concentrated likelihood function is chosen. The short-term parameters are concentrated out based on the full sample so that only the long-run part is estimated recursively. For the confidence intervals of the recursive eigenvalues, the standard errors are estimated on the basis of the full sample. The tests provide no evidence of model instability as the recursive eigenvalue remains stable during the testing period. Moreover, the τ-statistic is much lower than the 5% critical value. These results are displayed in Figure 3.

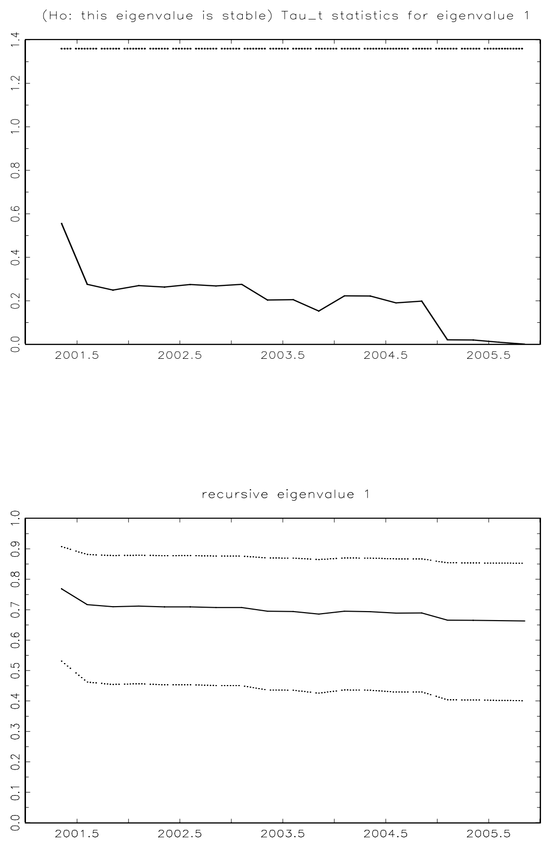

Figure 3. Recursive eigenvalue tests

4.2 Deflation and Possible Non-Linearity of Money Demand

The previous analysis, including tests for misspecification and model stability, did not suggest a rejection of the estimated system. However, a linear specification for money demand may be debatable for an economy that has moved from an environment of significantly positive inflation rates to near-zero inflation and even deflation. Cargill and Parker (2004) argue that deflation may cause discontinuity in the monetary policy process and lead to asymmetries in money demand. If nominal rates are relatively rigid (as has been the case in China) or have already hit the zero bound, higher real interest rates decrease investment spending. Similarly, expectations of future deflation may reduce consumption spending due to consumers' anticipation of even lower future prices. Lower inflation also reduces the costs associated with sub-optimal money holdings that are due to time spent on conducting real transactions (see e.g. McCallum and Goodfriend, 1987). But even more generally, the behavior of the money-holding sector may change in a non-linear fashion simply due to a significant change in the economic environment. Downward movements in the Chinese price level were entirely absent from the mid-1970s until 1997 (using annual data for retail prices where longer time series are available).

We investigate non-linearity in the system using the tests proposed by Teräsvirta (1994), whereby the null hypothesis of linearity is tested against a smooth transition regression (STR) model. The STR model is a nonlinear regression that nests a linear model. Logistic STR models include descriptions of processes where the dynamic properties differ from one regime to the other (perhaps from inflation to deflation) and the transition from one regime to the other is modeled to be smooth. The tests were carried out on the real money balances equation (4) of the VEC model, and the inflation rate was used as the primary transition variable from one regime to another. Accordingly, we assume that money demand may display non-linear behavior in a transition from an inflationary to a deflationary environment. We also conducted the linearity tests using other system variables as the transition variables. The reader is referred to Teräsvirta (1994, 2004) for details on the testing sequence and the exact hypotheses underlying the tests. We report the p-values from the testing procedure in Table 5.

	Transition variable		
Hypothesis	$\Delta\pi_{t-1}$	$\Delta(m-p)_{t-1}$	$\Delta(m-p)_{t-3}$
H_0	0.62		
H_{04}	0.95		
H_{03}	0.32	0.38	0.61
H_{02}	0.18	0.18	0.99
	Δy_{t-3}	$\Delta neer_{t-1}$	ec_{t-1}
H_0			
H_{04}			
H_{03}	0.57	0.71	0.23
H_{02}	0.83	0.37	0.58

Table 5. p-values of linearity tests of Equation (4)

* indicates significance at 10% level, ** at 5% and *** at 1% level. Unreported p-value indicates a matrix inversion problem in the testing procedure.

We find no evidence against our linear specification; the null hypothesis of linearity is not rejected even at 10% level of significance using any of the possible transition variables in the system. Note that the entire testing sequence was feasible only in the case of our primary transition variable, the inflation rate, due to matrix invertibility problems that arise when the values of the transition variable are close to zero or one. In sum, these results suggest that a linear specification is adequate for the Chinese economy during the disinflationary period examined in our study. Further, they support the findings from the previous misspecification tests, where normality of the residuals could not be rejected.

4.3 Impulse Response Analysis

More evidence of model dynamics is obtained from impulse response analysis, whereby we trace the effects of various shocks through the estimated system. Such analysis can also be regarded as a type of causality test. Impulse responses in our case are based on a structural vector error correction (SVEC) model with contemporaneous restrictions, where a short-run timing scheme is used to impose the structure of the model. We write the SVEC form as:

$$\mathrm{A}\Delta x_t = \Pi^* x_{t-1} + \Gamma_1^* \Delta x_{t-1} + ... + \Gamma_{p-1}^* \Delta x_{t-p+1} + \mathrm{B}\varepsilon_t, \tag{8}$$

where $\varepsilon_t \sim (0, I_K)$. $\Pi^*$and $\Gamma_j^*(j = 1, ..., p-1)$ are structural form parameter matrices. The invertible matrix A allows for the modelling of the instantaneous relations between the variables. The structural shocks, ε_t, are related to the model residuals by linear relations. We assume that these structural shocks are mutually uncorrelated and therefore orthogonal. Premultiplying (8) by A^{-1} leads to the familiar reduced form (2), where $\Pi = \mathrm{A}^{-1}\Pi^*$and $\Gamma_j = \mathrm{A}^{-1}\Gamma_j^*(j = 1, ..., p-1)$. Specifying A to be an identity matrix, the reduced form disturbances are linked to the underlying structural shocks by

$$u_t = \mathrm{B}\varepsilon_t. \tag{9}$$

For a K-dimensional system, $K(K-1)/2$ restrictions are necessary for orthogonalizing the shocks. In our case of four endogenous variables, this amounts to six restrictions. With the ordering of variables in our system as $(m-p), y$, $\pi, neer$, the following restrictions on the B matrix are imposed:

$$\mathrm{B} = \begin{bmatrix} * & * & * & 0 \\ 0 & * & 0 & 0 \\ 0 & * & * & 0 \\ * & * & * & * \end{bmatrix}, \tag{10}$$

where asterisks denote unrestricted elements. In our system, a shock to real output can have an instantaneous impact on all other variables, whereas a shock to the nominal effective exchange rate cannot have an instantaneous impact on any other variables. While only shocks to real output are allowed to have an instantaneous impact on the rate of inflation, real money stock reacts contemporaneously to all other variables except the nominal effective exchange rate. Therefore, the exchange rate in our system is a forward-looking asset price that is allowed to react rapidly to developments in the real economy. This is similar to Kim and Roubini (2000). Stickiness in real output can be justified by assuming that consumption and investment plans are to a significant extent predetermined, as in the model by Rotemberg and Woodford (1999), while price rigidity can simply be a consequence of costly price adjustment (e.g. due to menu costs). Finally, it is worth noting that due to the contemporaneous nature of our restrictions, the lags of the effects are left unrestricted.

The reduced form covariance matrix, subject to restrictions imposed in the structural form, is next used to obtain estimates for the contemporaneous impact matrix (not displayed). We use maximum likelihood estimation, together with numerical optimization methods in the form of a scoring algorithm (see Amisano and Giannini, 1997; Breitung et al., 2004). Parameter uncertainty is illustrated through Hall bootstrapped percentile confidence intervals. The conventional 95% significance level was chosen, and the number of bootstrap replications was set at 5,000. The impulse responses from our system are displayed in Figure 4 below.

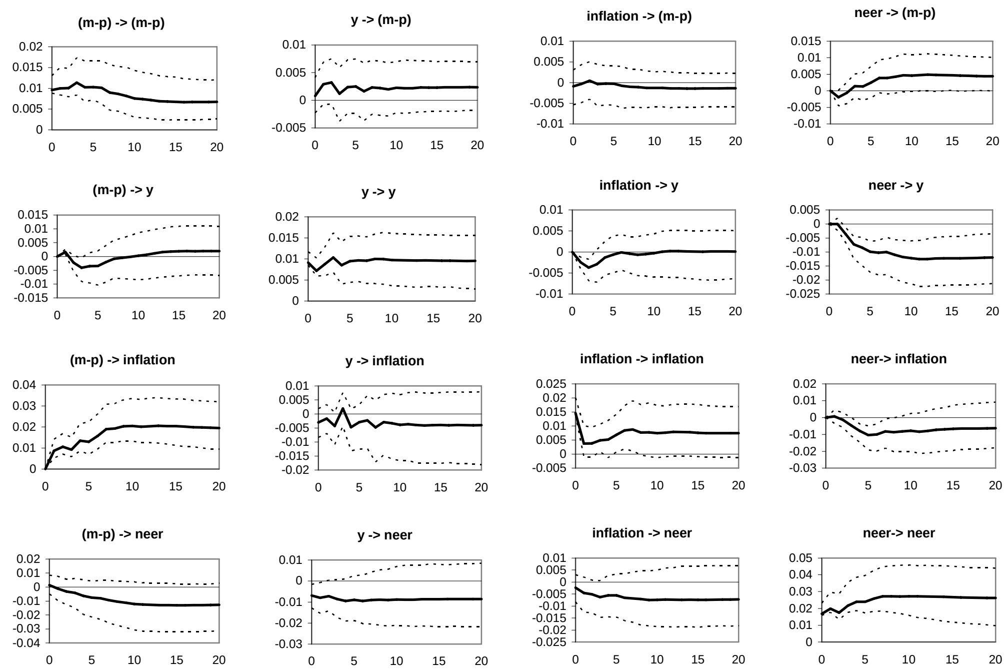

Figure 4. Impulse responses in the money demand system

A shock to real money leads to a permanent increase in the real money stock in our system. Importantly, the shock also leads to a statistically significant increase in the inflation rate, which stabilizes ten quarters after the shock. Again, our result would seem to justify the PBoC's attention on the money stock as defined by broad money M2. A shock to the inflation rate has a negligible impact on the money stock, again providing evidence that the inflation rate has not played the role of an opportunity cost variable during our estimation sample. It is notable that an appreciation shock to the nominal effective exchange rate leads to a permanent increase in the real money stock, although the impact is only borderline significant at 95% level. Such a finding is in line with the result from the cointegration analysis. It suggests that developments in the nominal effective exchange rate are important in a monetary stabilization policy that specifies intermediate money growth targets. The fact that an appreciation shock to the nominal effective exchange rate leads to a fall in the level of real GDP also deserves comment. It is an interesting finding in light of the appreciation pressure on the Chinese renminbi. The same shock also leads to a statistically significant fall in the inflation rate in the short run. Our results from the impulse response analysis can be roughly compared to those by Gerlach and Kong (2005), who provided results from a trivariate system without an exchange rate. While they also find shocks to the money stock to lead to increases in both the price level and real GDP, the statistical significance of their results cannot be judged due to the omission of confidence intervals. To our knowledge, no other studies on Chinese money demand provide estimates from impulse response analysis.

We additionally performed impulse response analysis using a system specification, where a shock to prices can have an instantaneous impact on all the other variables, and a shock to the real money stock cannot have an instantaneous impact on any of the other variables. Only shocks to the inflation rate are allowed to have an instantaneous impact on real GDP, and the nominal effective exchange rate reacts contemporaneously to all other variables except the real money stock. The B matrix for such a system can be written as:

$$\mathrm{B} = \begin{bmatrix} * & * & * & * \\ 0 & * & * & 0 \\ 0 & 0 & * & 0 \\ 0 & * & * & * \end{bmatrix}. \tag{11}$$

While the structure of the above system is more difficult to justify than our benchmark specification, some elements of the previous timing scheme are maintained (i.e. the nominal effective exchange rate and the money stock still are relatively responsive to the various shocks; prices and output less so). The specification can therefore be considered a robustness test. The literature on VAR models suggests that even minor changes to the contemporaneous identifying restrictions can produce different responses to shocks. The impulse responses obtained from (11) bear close resemblance to the benchmark system, however. The only material difference is that shocks to broad money now have a borderline statistically significant impact on the nominal effective exchange rate. The statistical significance of the exchange rate shock on inflation is lower than that obtained in our benchmark specification.

In sum, shocks to broad money in our system lead to statistically significant increases in the rate of inflation. Coupled with the observed stability of the estimated system, our findings provide support for a strategy to target the rate of money growth in order to conduct macroeconomic stabilization. Interestingly, as shocks to the nominal effective exchange rate lead to increases in the real money stock, monetary targeting by the PBoC should take movements in the yuan effective exchange rate into account.

5 Conclusions

We set out to examine the demand for broad money M2 in the Chinese economy during a period characterized by a significant process of disinflation and even deflation. Our paper established a stable money demand relationship in a vector error correction framework. The finding of system stability, together with the observation of broad money shocks leading to increased inflation, supports the PBoC's current policy of specifying intermediate targets defined in terms of the growth rate of broad money M2. Interestingly, our results suggest that movements in the nominal effective exchange rate should be taken into account in a successful conduct of such a policy.

The focus on broad money and bank lending growth has been an obvious policy choice in the Chinese financial system that is strongly dominated by the banking sector. However, as interest rates assume a more prominent role in the transmission mechanism, it may be preferable to move to a policy that specifies an inflation target for the monetary authority without an intermediate target spelled out in terms of growth rates of the money stock. Indeed, as argued by Svensson (1999), monetary targeting corresponds to optimal policy only if the central bank's instrument and the state of the economy first affect money growth and this corresponding money growth is the only determinant of inflation. Our results, while providing some evidence that the necessary conditions for monetary targeting in terms of model stability have been fulfilled, do not provide evidence of this special type of recursive transmission mechanism. Note also that explicit inflation targeting need only take into account movements in the exchange rate to the extent that they affect the preferred price index and not the developments of some intermediate target variable. Nevertheless, the specification of targets for money and credit growth can be expected to serve the Chinese economy well as its transition continues, provided that the monetary authorities pay sufficient attention to movements in the exchange rate and the resulting (authorized and unauthorized) capital flows.

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