Research Paper No. 2004/56

Output and Price Fluctuations in China's Reform Years

What Role did Money Play?

Yin Zhang* and Guanghua Wan*

August 2004

Abstract

The Chinese economy underwent cyclical fluctuations in growth and inflation in the reform period. Contrasting views exist on the role of money in such fluctuations. This paper assesses these views employing structural VEC models based on the exchange equation. It is found that in the long run money accommodates, rather than causes, changes in output and prices. In the short run, price fluctuations are mostly attributable to shocks that have permanent effects on prices and money but not on real output. These shocks also account for a large proportion of fluctuations in money, and strongly influence the movements of output.

Keywords: money, output fluctuation, price fluctuation, structural VEC model, China

JEL classification: E32, E49, O53

* UNU-WIDER, Katajanokanlaituri 6 B, FIN-00160 Helsinki, Finland

This study has been prepared within the UNU-WIDER internship programme and research on 'New Directions in Development Economics'.

UNU-WIDER acknowledges the financial contributions to the research programme by the governments of Denmark (Royal Ministry of Foreign Affairs), Finland (Ministry for Foreign Affairs), Norway (Royal Ministry of Foreign Affairs), Sweden (Swedish International Development Cooperation Agency – Sida) and the United Kingdom (Department for International Development).

ISSN 1810-2611 ISBN 92-9190-645-X (internet version)

UNU World Institute for Development Economics Research (UNU-WIDER)
Katajanokanlaituri 6 B, 00160 Helsinki, Finland

Camera-ready typescript prepared by Anne Ruohonen at UNU-WIDER
Printed at UNU-WIDER, Helsinki

1 Introduction

The role of money in initiating and transmitting macroeconomic fluctuations remains a question dividing opinions in the economics profession. The need for continuous research into this subject is also highlighted by the reliance on monetary policy as the principal tool for economic stabilisation in market economies. Inflation and output fluctuations in transition economies provide the occasion to re-examine alternative theories in the context of fast-changing economic and social institutions. As monetary management in these economies used to be the residual component of central planning, achieving a better understanding of the interactions between money and other macroeconomic variables is all the more imperative.

This paper examines the experience of China – the largest transition economy. The Chinese economy exhibited cyclical fluctuations in output growth and inflation in the reform period. Different hypotheses have been advanced about the causes and mechanisms of these fluctuations. These conjectures differ in the roles they ascribe to money and thus in their implications for monetary policy. The majority of empirical investigations consist in testing Granger causality between money, output and prices (e.g., Chen, 1989; Chan, Deaves and Wang, 1992; Li and Leung, 1994; Hasan and Taghavi, 1996; Yu, 1997) and in estimating the money demand function (e.g., Chow, 1987; Hafer and Kutan, 1994; Huang, 1994; Qin, 1994; Hasan, 1999). While the latter studies claimed some success, the causality tests returned ambiguous findings: uni-directional causality from money to output and/or prices, the other way around, and bi-directional causality. Thus, it can be said that the existing empirical studies fail to provide an evaluation of the validity of the various arguments about money.

Partly responsible for the failure are some of the limitations associated with the Granger (non-)causality test and the conventional specification of the money demand function. Apart from the well-known pitfalls of interpreting temporal precedence as theoretical causality, the Granger test is sensitive to the specification of the vector auto-regression models (Stock and Watson, 1989). More importantly, empirically-derived pairwise causality, even if undisputed, reveals little about the nature of the shocks initiating changes in the 'causing' variable, neither does it offer a measure quantifying the responses of the 'caused' variable. The conventional money demand function describes an equilibrium relationship that attains only in the long run. It tells little about the short-run feedbacks among money and the other variables. In typical specifications of the function, homogeneity is imposed between prices and nominal balances. As shown below, this restriction may well be invalid.

A more efficient way to uncover the structural relationship between money, output and prices is to directly test the assumptions underlying the different views. To minimise the interference of model specification in the results, it is desirable to conduct the testing in a unified framework. In section 2, we argue that the Fisher equation of exchange is of sufficient flexibility to accommodate the competing views. Also, we emphasise the need to distinguish between the long-run equilibrium and the short-run dynamics. This makes the structural vector error correction (VEC) approach the suitable modelling strategy. As detailed in section 3, we undertake the empirical analysis in two stages: The first stage focuses on identifying the form of the long-run relationship among money, output and prices and the responsiveness of individual variables to monetary disequilibrium. In the second stage, information from the first stage is utilised to characterise a structural model so that the transmission of different shocks can be traced and their impacts quantified. In section 4, we discuss major findings. A stable long-run relation among money, output and prices is identified. However, most of the fluctuations in output and prices are not attributable to monetary shocks. It seems that money accommodated, rather than caused, the changes in output and prices. The paper concludes with a summary of the monetary aspects of China's output and price fluctuations in section 5.

2 Alternative views about the money-output-price nexus

The different views about the role of money in output and price fluctuations in China's reform period are summarised in Table 1. The first view doubts the existence of a systematic relationship among money, output and prices in transition economies. It argues that many factors could be disruptive to money demand, for example the release of the monetary overhang caused by quantity-rationing in the planning era, households' enhanced motive for precautionary saving, monetisation and re-monetisation of some economic transactions, and increased financial sophistication. Consequently, even if money does influence output and prices, the relationship would not be sufficiently stable to be exploitable by monetary policy. This argument may explain why many transition economies opted for exchange-rate-based stabilisation instead of money-based stabilisation.[1]

The second view holds that the major difficulty in targeting monetary aggregates in transition economies arises not from the instability of money demand, but from the endogeneity of money (Peebles, 1991; Development Research Institute, 1995; Chang and Hou, 1997). Setting up quantity targets is pointless when the government lacks effective means or political will to achieve them. This does not imply, however, that monetary aggregates should be consigned to oblivion. Rather, money contains information about output, prices and other determinants of money demand. As monetary data are relatively easier to collect, changes in monetary aggregates should be monitored and fed into the formulation of fiscal, income, interest rate and exchange rate policies.

Proponents of the third view refer to the fact that China's financial system is dominated by state banks (Bennett and Dixon, 2001). Most firms, particularly state-owned enterprises (SOEs), rely on bank credits to finance investment or even working capital. Therefore, monetary policy impacts directly on aggregate demand and supply via bank lending. Its effects on prices are uncertain due to the non-monetary causes of inflation such as relative price realignment, price reform, fluctuations in grain prices, currency devaluation, and so on.

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1 For example, Estonia, Lithuania and Bulgaria adopted the currency board regime. Hungary, Latvia and the Slovak Republic announced narrow bands for their exchange rates.

Table 1 Alternative views about the money-output-price nexus

Views	Testable implications			Policy implications	Stylized representation	Implementation in VEC model
	Long-run relation	Money	Others			
1	unstable	–	–	money cannot be the nominal anchor	–	no existence of cointegration
2	stable	endogenous		(a) quantity targeting is inappropriate, (b) money can be the information variable for other types of monetary policy regimes and demand management policies	$m^s = m^d = a_1 y + a_2 p - a_3 i$	(a) existence of cointegration, (b) rejection of weak exogeneity of money
3	stable	exogenous	changes in money affect real output	quantity targeting for output stabilisation	$y = \frac{1}{a_1} m^s + \frac{a_2}{a_1} p + \frac{a_3}{a_1} i$	(a) existence of cointegration, (b) weak exogeneity of money, (c) real output is more responsive to the error-correction term than are prices
4	stable	exogenous	changes in money affect prices	quantity targeting for price stability	$p = \frac{1}{a_2} m^s - \frac{a_1}{a_2} y + \frac{a_3}{a_2} i$	(a) existence of cointegration, (b) weak exogeneity of money, (c) prices are more responsive to the error-correction term than is real output
5	stable	exogenous	prices and money vary proportionally in the long run	(a) quantity targeting for long-run price stability, (b) uncertainty in short-run effects of monetary policy	$m^s - p = a_1 y - a_3 i$	(a) existence of cointegration, (b) weak exogeneity of money, (c) homogeneity of money and prices

By contrast, the fourth view disputes the link between money and real activity. It claims that certain institutional features of the Chinese economy, such as the government commitment to supporting the SOEs, make it possible for firms to circumvent financial constraints (Wang, 1991; Brandt and Zhu, 2000, 2001). Hence, changes in money supply will largely affect prices rather than real activity.

The studies attempting to estimate the demand function of *real* balances embody the fifth view. Implicit in this argument is the monetarist view that changes in money lead, on a one-to-one basis, to changes in prices in the same direction. Money may affect real output temporarily. In the long run, real output tends towards its 'natural' level which is determined by the amount of labour, capital, natural resources, technology and institutions, yet independent of the stock of money.

To assess the extent to which the competing views are borne out by reality, it is desirable to use a framework of sufficient generality to encompass all the alternatives. This would ensure the comparability of the testing results. A suitable vehicle is the equation of exchange, $MV = PT$, where M is the stock of nominal balances, V is the velocity of circulation, T is the amount of total transactions, and P a price index of T. The logarithmic form of the exchange equation is

$$m + v = p + t\,, \qquad (1)$$

where the lower case letters represent the logarithms of the variables denoted by the corresponding upper case letters. Although an *ex post* identity, equation (1) can be turned into stylised representations of views 2 to 5 when supplemented with causal assumptions. Before the transformation can be done, however, three modifications to the equation are necessary: (i) Total transactions need to be replaced by a measure of final output, y, since it is the latter that is of more interest to policy decisions. As there is no obvious reason that the ratio of total output to total transaction will stay constant, the coefficient on y may not be unitary. (ii) As a store of value, money competes with other forms of asset. The rate of return on alternative assets may bear significantly on velocity if money can be easily substituted for by the other assets. Therefore, a nominal interest rate that proxies for the opportunity cost of holding money should be included in the monetary equilibrium relationship.[2] (iii) Price-money homogeneity is routinely pre-imposed in previous studies about money demand. Since the factors affecting prices did not coincide exactly with those affecting money, the validity of this claim needs to be verified. Thus, the coefficients on money and prices should not be restricted to be equal as in equation (1). With these modifications, equation (1) becomes

$$m + a_3 i = a_1 y + a_2 p\,, \qquad (2)$$

where i is the nominal interest rate. Equation (2) can now be used as the framework for turning the alternative views into structural models.

The necessary condition for the first view to hold is the absence of a stable relationship among the four variables in equation (2). Should such a relationship exist, equation (2)

2 The Chinese currency was not convertible on current account until 1996 and is still not convertible on capital account. The substitution between domestic and foreign assets is expected to be negligible.

adequately describes the monetary equilibrium. Adding different behavioural assumptions to the equation yields four stylised monetary models, each corresponding to one of the remaining views. As shown in the second last column of Table 1, the view of endogenous money implies a model where the supply of nominal balances adjusts to meet the demand for nominal balances. The amount of money in existence is determined by the level of real output, prices and velocity (which in turn is influenced by the nominal interest rate). The model representing the third view is similar to the naïve LM model: prices are rigid; the stock of money is exogenously determined by money supply; changes in money supply affect real output. According to the fourth view, the effects of changes in money supply fall primarily on prices. It is thus essentially a hypothesis about the determination of prices. A representation of this view can be obtained from equation (2) by making prices the endogenous variable. The fifth view rests upon three assumptions – exogenous money, endogenous prices and the proportionality between money and prices. Restricting the coefficient on price, a_2, to unity and making real balances the dependent variable change equation (2) into a model for the fifth view.

Clearly, the empirical relevance of these models would vary with the time horizon of investigation. For example, endogenous variables may anticipate as well as respond to changes in exogenous variables. Such interactions complicate the short-run temporal ordering of changes in exogenous and endogenous variables. Investigations focusing on the short run may end up with supporting evidence for several views. Furthermore, some of the implications of the five views are more tenable in the long run. Because the adaptation of expectations and adjustment of prices take time to complete, short-run investigations may fail to detect any stable link among the variables even though an equilibrium relationship does exist. By the same token, the price-money homogeneity postulated by view 5 may only be discernible in the long run. A proper empirical evaluation of what the observed relation between money, prices and output means causally, therefore, necessitates differentiating between the short run and the long run. The structural VEC modelling approach lends itself easily to such a distinction and is adopted for this study.

3 Econometric implementation

Introductions to structural VEC modelling can be found in a number of econometrics textbooks (e.g., Hamilton, 1994; Enders, 2004). A brief description of the approach suffices here. Suppose that the $n \times 1$ vector $\mathbf{x}$ is $I(1)$. If there exists a non-singular $n \times r$ matrix $\boldsymbol{\beta}$ such that $\boldsymbol{\beta}'\mathbf{x}$ is $I(0)$, then the variables in $\mathbf{x}$ are said to be cointegrated. According to the Granger representation theorem, the reduced-form vector autoregression (VAR) of the first difference of $\mathbf{x}$ can be re-parameterised to

$$\Delta \mathbf{x}_t = \sum_{k=1}^{p-1} \Pi_k \Delta \mathbf{x}_{t-k} - \Pi \mathbf{x}_{t-1} + \mathbf{e}_t \;, \qquad (3)$$

where Δ is the first difference operator, Π_k's and Π are $n \times n$ matrices, $\mathbf{e}$ is a vector of white noises with zero means and variance-covariance matrix Σ, and the deterministic components of $\Delta\mathbf{x}$ have been suppressed to simplify the notation. The matrix $\prod$ can be decomposed as $\Pi = \boldsymbol{\alpha\beta}'$, where $\boldsymbol{\alpha}$ has the same dimension as $\boldsymbol{\beta}$. If the parameters in each column of $\boldsymbol{\beta}$ are considered as depicting an equilibrium relationship among the variables of $\mathbf{x}$, then $\boldsymbol{\beta}'\mathbf{x}_{t-1}$ represents the deviations from the r equilibrium relationships in the previous period, and the speed of reversion to equilibrium is captured by the parameters in $\boldsymbol{\alpha}$. Johansen (1988) and Johansen and Juselius (1990) show that the number of cointegration vectors, r, can be determined by maximum likelihood estimation and likelihood ratio (LR) testing. Restrictions on $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ can be tested with the same techniques.

The vector moving average (VMA) form of equation (3) can be written as $\Delta\mathbf{x}_t = C(L)\mathbf{e}_t$, where $C(L)$ is a $n \times n$ matrix of polynomials in the lag operator L. This VMA equation cannot be directly used for computing the impulse response functions and forecast error variance decompositions because the Wold innovations in $\mathbf{e}$ are not mutually independent. Intuitively, each element of $\mathbf{e}$ is a combination of the fundamental shocks driving the system, and as such does not admit of an economic interpretation. Assume that $\mathbf{e}$ is a linear transformation of the fundamental shocks $\boldsymbol{\varepsilon}$, that is $\mathbf{e} = A\boldsymbol{\varepsilon}$, where A is an $n \times n$ matrix, $A \neq I_n$, and the elements of $\boldsymbol{\varepsilon}$ have zero means and are not serially or cross-correlated. The structural VMA of equation (3) can then be expressed as

$$\Delta\mathbf{x}_t = C(L)A\boldsymbol{\varepsilon}_t = A(L)\boldsymbol{\varepsilon}_t . \qquad (4)$$

Exact identification of the parameters in A requires $n(n-1)/2$ additional restrictions apart from normalising the variance of $\boldsymbol{\varepsilon}$ to unity and making use of the estimated Σ.

King et al. (1991) demonstrate that the information about the number of cointegration vectors can be combined with restrictions on the long-run impact matrix $A(1)$ to partially identify A. More specifically, because $\boldsymbol{\beta}$ has a rank of r, the n variables comprising $\mathbf{x}$ share $n - r$ stochastic trends, each driven by an independent shock. These $n - r$ shocks leave permanent imprints on $\mathbf{x}$ whereas the effects of the remaining r shocks vanish in time. This entails that the matrix $A(1)$ must have a reduced rank of $n - r$, thereby separating the permanent shocks from the transitory shocks. King *et al.* suggest that the Blanchard-Quah approach (Blanchard and Quah, 1989) be used to identify the permanent shocks, that is the block in $A(1)$ corresponding to the permanent shocks be made triangular (upon proper re-ordering).[3]

To apply the structural VEC approach to testing the views discussed in section 2, the first thing to ascertain is whether money, output, prices and the nominal interest rate or any selection from them are cointegrated. Evidence against cointegration is not equivalent to conclusive proof of the first view. Non-cointegration may result from missing variable bias, that is the four variables do not fully summarise the monetary equilibrium. Evidence in favour of cointegration, however, disproves the first view and establishes that the four variables are systematically linked together. In addition, it is expected that there should be only one cointegration relationship (or, $\boldsymbol{\beta}$ has only one column).

Once cointegration is confirmed, the price-money homogeneity hypothesis implied by the fifth view can be tested by checking if the coefficient estimates on money and prices

3 This would still leave the r transitory shocks unidentified unless $r = 1$.

in the cointegration vector $\boldsymbol{\beta}$ are of equal magnitude but opposite signs. Also of interest is whether the nominal interest rate belongs in the cointegration vector. This question is not merely of econometric concern (because the inclusion of superfluous variables would bias coefficient estimates), it is actually associated with whether households' portfolio decisions impinge upon monetary equilibrium. The nominal interest rate can be excluded from the equilibrium relationship if its coefficient in $\boldsymbol{\beta}$ is not significantly different from zero.

To compare the validity of views 2 through 4, two types of evidence will be examined. The first type of evidence is the responsiveness of money, output and prices to monetary disequilibrium, which is indicated by the coefficient estimates in $\boldsymbol{\alpha}$ in the respective equations of the three variables. If a variable has an insignificant $\boldsymbol{\alpha}$ coefficient in its equation, that variable is said to be weakly exogenous. Weakly exogenous variables barely respond to monetary disequilibrium. Thus, if money is found to be weakly exogenous, view 2 becomes untenable. Conversely, rejection of the weak exogeneity of money would cast doubt on views 3 and 4. With exogenous money, the relative merits of views 3 and 4 can be assessed by testing and comparing the significance and magnitudes of the $\boldsymbol{\alpha}$ coefficients in the price and output equations.

Weakly exogenous variables can be considered as the sources of the permanent shocks. This knowledge, together with the number of cointegration vectors, would help with determining the form of $A(1)$, and hence with the identification of the structural model. The innovation accounting tools – impulse responses functions and forecast error variance decompositions – applied to the identified model will provide insights into the transmission processes of different shocks and their historical contributions to price and output fluctuations. These constitute the second type of evidence.

Several empirical issues arise. The first issue concerns the practical definition of money. Apparently, whether money is defined narrowly or broadly would have a bearing on the testing results. A diversity of measures of money have been used in previous studies, with no clear indication which one is superior to the others. It is thus decided that three candidate measures, for which quarterly data are available, be subjected to the cointegration test. The three monetary aggregates are: M_0 (currency in circulation), M_1 (the sum of M_0 and demand deposits), and M_2 (the sum of M_1, time deposits, saving deposits and other deposits).

The second issue is the choice of the nominal interest rate variable. Among the interest rate series relevant to this study, the rate on one-year savings deposits is the only one that is consistently defined and available for a sufficient length of time. Unfortunately, this interest rate is not a market rate, but is administratively determined. What this rate denotes differs for the three monetary aggregates. Saving deposits are a major component of M_2, yet a substitute asset for M_0 and M_1. Thus, the interest rate on savings deposits is expected to be related negatively to M_0 and M_1, and positively to M_2.

The selection of the real output variable is also constrained by the availability of data. Quarterly data of real GDP start from 1994 only. Industrial production is thus used as a proxy. For prices, the RPI and the CPI are the only price indices of adequate length. They behaved quite similarly during the sample period 1985–2000. Because the CPI includes the prices of transport, housing, medical services, and so on, which had been under strict government control until recently, the RPI is chosen to reduce bias.

However, even the RPI, particularly the pre-1991 data, may still suffer from distortions caused by dual-track pricing.

The output and price data are from *China Monthly Statistics* published by the National Bureau of Statistics. The original monthly series are converted to quarterly series and then adjusted by the X-11 procedure. Monetary and interest rate data are quarterly series from the IMF (2002) *International Financial Statistics*. As the unit-root tests results in Table 2 suggest, all series can be characterised as $I(1)$ processes.[4]

Table 2 Results of unit-root tests

Variable (x_t)	ADF[a]		PP (l=3)[b]		KPSS (l=3)[c]	
	τ_μ	τ_τ	Z_μ	Z_τ	η_μ	η_τ
Industrial production	-0.19 (l=1)	-2.05 (l=2)	-0.17	-2.13	1.69	0.15
RPI	-1.47 (l=1)	-1.27 (l=1)	-1.90	-0.10	1.61	0.24
Currency in circulation (M_0)	-1.70 (l=0)	-0.60 (l=0)	-1.57	-0.86	1.68	0.28
Narrow money (M_1)	-0.51 (l=0)	-1.54 (l=0)	-0.49	-1.80	1.70	0.15
Broad money (M_2)	-2.15 (l=0)	0.12 (l=0)	-2.13	0.14	1.70	0.22
Interest rate	-0.95 (l=0)	-0.12 (l=0)	-0.63	-1.25	0.54	0.30

Notes: The sample period is Q1/1985–Q4/2000. Reported here are results concerning the levels of the series. The same tests have also been carried out on their first-differences. The results, available upon request, show that these series are stationary in first-differences.

[a] Augmented Dickey-Fuller test. Figures in the parentheses are the number of lags included in the regressions. The lag lengths are chosen by adding lags until the Ljung-Box Q-statistics of order eight fails to reject no serial correlation at the 5% significance level. The critical values for τ_μ at the 5% and 10% significance levels are –2.89 and –2.58. Those for τ_τ are –3.45 and –3.15.

[b] Phillips-Perron test. The bandwidth of the Bartlett kernel is chosen according to Newey and West (1987). The critical values for Z_μ and Z_τ correspond to those for τ_μ and τ_τ respectively.

[c] This test is proposed by Kwiatkowski et al. (1992). The critical values for η_μ at the 5% and 10% significance levels are 0.463 and 0.347. Those for η_τ are 0.146 and 0.119.

4 Perron (1989) contends that breaks in the deterministic components of a time series can bias the results of unit-root tests towards non-stationarity. Following Perron (1997), the series are subjected to a procedure which determines the date of structural shifts endogenously. It turns out that the unit root hypothesis is supported in all cases. The first differences of these series are stationary. Unit-root test results for the first differences are not shown in Table 2.

4 Empirical findings and implications

4.1 The long-run relation between money, output and prices

The models to be examined are three VAR systems of the vector, $\mathbf{x} = [y, m, p, i]'$, where m denotes, respectively, the logarithms of M_0, M_1 and M_2 in different specifications, y and p are the logarithms of industrial production and the RPI, and i is the nominal interest rate on one-year saving deposits which enters the models in percentage points. All equations in the VARs include a constant term.

Estimation of the long-run relation. The information criteria BIC and HQ are used to set the lag orders of the VARs. Where the two criteria do not agree, the LR test is invoked as a supplementary test. Based on the results in Table 3, a lag order of two is chosen for each model. The cointegration analysis adopts the Johansen procedure. Tables 4–6 each present the results for one model. Found in panel A of each table are the results of the λ_{trace} test. The statistics have been adjusted by a small sample correction factor proposed in Reinsel and Ahn (1992).

For the M_0 model, the null of no cointegration (r = 0) can be accepted at the 5 per cent significance level, but is rejected at the 10 per cent level. There is no evidence for more than one cointegration relations. Setting r = 1 produces the estimates of $\boldsymbol{\beta}$ and $\boldsymbol{\alpha}$ in panel B of Table 4. According to the estimated $\boldsymbol{\beta}$**,** M_0 is related negatively to prices and positively to the interest rate. This is apparently nonsensical. Moreover, the coefficient estimate for M_0 in $\hat{\boldsymbol{\alpha}}$ is positive, indicating instability of the system. Taken together, the results in Table 4 indicate that no equilibrium relationship exists among M_0, industrial production, prices and the interest rate.

Results for the M_1 model are reported in Table 5, where the rank test suggests the existence of one cointegration relation. The estimates in $\hat{\boldsymbol{\alpha}}$ and $\hat{\boldsymbol{\beta}}$ are all of the expected signs. In particular, the coefficient on the interest rate in $\hat{\boldsymbol{\beta}}$ is now positive, conforming to the earlier discussion that the stock of narrow money should have an inverse relation with the opportunity costs of holding it – the nominal returns on saving deposits. The coefficients in the estimated loading vector $\hat{\boldsymbol{\alpha}}$ demonstrate that the estimated cointegration vector is an equilibrium relation towards which the system gravitates.

Table 3 Lag length of the models

Models	BIC	HQ	LR
M_0	2	2	
M_1	1	2	H_0: p=1, H_1: p=2, χ^2 (16)=58.17 (0.00)
M_2	1	2	H_0: p=1, H_1: p=2, χ^2 (16)=56.81 (0.00)

Notes: The values in the 'BIC' and 'HQ' columns indicate the optimal lag orders according to the respective criteria. H_0 in the 'LR' column denotes the null hypothesis and H_1 the alternative hypothesis. The LR test statistics have a chi-square distribution with a degree of freedom of 16. Figures in parentheses are the marginal significance of the LR statistics.

Table 4 Cointegration analysis of the M_0 model

		Panel A: Rank test			
	H_0	$r = 0$	$r \leq 1$	$r \leq 2$	$r \leq 3$
λ_{trace} tests	H_1	$r > 0$	$r > 1$	$r > 2$	$r > 3$
	adjusted statistics[a]	46.75*	21.95	7.34	0.10
		Panel B: Estimated error-correction parameters			
		y	m	p	i
cointegration vector ($\hat{\boldsymbol{\beta}}$)		-2.885	1.000	1.790	-0.128
loading vector ($\hat{\boldsymbol{\alpha}}$)		0.018	0.014	-0.031	1.083

Notes:

[a] The small sample correction factor is $(T - np)/T$, where T is the number of effective observations, n denotes the dimension of the model and p is the lag length of the model.

* The null is rejected at the significance level of 10%. Critical values are taken from Osterwald-Lenum (1992).

The rank test results in Table 6 provide strong evidence for the presence of one cointegration vector in the M_2 model. The estimated feedbacks of the four variables to short-run deviations from equilibrium, as displayed by $\hat{\boldsymbol{\alpha}}$, demonstrate the stability of the system. The unexpected result is that the interest rate takes on a negative coefficient in $\hat{\boldsymbol{\beta}}$. Nonetheless, one is cautioned against classifying this finding as an anomaly before verifying that the interest rate does belong in the long-run relation.

Since the cointegration rank of the M_1 and M_2 models is one, no restrictions are needed for identifying $\boldsymbol{\beta}$ and $\boldsymbol{\alpha}$.[5] Any restriction on either of them is over-identifying. Thus, the significance of the interest rate coefficients in $\hat{\boldsymbol{\beta}}$ and $\hat{\boldsymbol{\alpha}}$ can be tested. In the last row of panel B in Table 5, the interest rate is found to be part of the cointegration relation in the M_1 model – the null of a zero coefficient can be rejected at the 8 per cent significance level. By contrast, the same hypothesis cannot be rejected for the M_2 model even at the 35 per cent significance level. The long-run exclusion of the interest rate might be an explanation for the wrong sign of its coefficient in $\hat{\boldsymbol{\beta}}$. Also, the interest rate is weakly exogenous in both models. The joint hypothesis of the long-run exclusion and weak exogeneity of the interest rate is accepted in the M_2 model.

Based on the above test results, the preliminary models are modified and re-estimated. The new M_1 model is conditioned on the interest rate variable, that is there is no equation for the interest rate variable in the VEC system, but the interest rate still stays in the cointegration vector. In the new M_2 model, the interest rate variable is excluded from the cointegration vector. The new estimates of $\boldsymbol{\beta}$ and $\boldsymbol{\alpha}$ are reported in panel C of

5 In general, a cointegration rank of r requires $r(r-1)$ restrictions to identify $\boldsymbol{\beta}$ and $\boldsymbol{\alpha}$.

Tables 5 and 6. In what follows, the testing of price-money homogeneity and weak exogeneity is conducted on the modified M_1 and M_2 models only.

Table 5 Cointegration analysis of the M_1 model

Panel A: Rank test

	H_0	$r = 0$	$r \leq 1$	$r \leq 2$	$r \leq 3$
λ_{trace} tests	H_1	$r > 0$	$r > 1$	$r > 2$	$r > 3$
	adjusted statistics[a]	48.22**	23.27	7.50	0.03

Panel B: The full model

	y	*m*	*p*	*i*
cointegration vector ($\hat{\boldsymbol{\beta}}$)	-1.346	1.000	-0.373	0.019
loading vector ($\hat{\boldsymbol{\alpha}}$)	0.338	-0.240	0.052	-0.848

H_0: zero semi-interest-rate elasticity[b]	$\chi^2(1)$=3.15 (0.08)
H_0: weak exogeneity of interest rate[b]	$\chi^2(1)$=0.10 (0.76)

Panel C: The partial system conditioned on the interest rate

	y	*m*	*P*	*i*
cointegration vector ($\hat{\boldsymbol{\beta}}$)	-1.354	1.000	-0.362	0.018
loading vector ($\hat{\boldsymbol{\alpha}}$)	0.341	-0.242	0.052	-

H_0: price homogeneity[b]	$\chi^2(1)$=5.25 (0.02)
H_0: weak exogeneity of money[b]	$\chi^2(1)$=3.55 (0.06)
H_0: weak exogeneity of output[b]	$\chi^2(1)$=10.30 (0.0)
H_0: weak exogeneity of prices[b]	$\chi^2(1)$=0.76 (0.38)

	L-B(15): $\chi^2(123)$=126.55 (0.39)		
Autocorrelation[c]	LM(1): $\chi^2(9)$=12.85 (0.17)	Normality[c]	$\chi^2(9)$=31.09 (0.00)
	LM(4): $\chi^2(9)$=12.83 (0.17)		

Notes:

[a] See notes to Table 3.

** The null is rejected at the 5% significance level.

[b] These are LR tests. The test statistics are χ^2 distributed. Values in parentheses indicate the marginal significance of rejecting the null.

[c] These are multivariate autocorrelation and normality testes of the residuals. L-B(15): Ljung-Box test based on the auto- and cross-correlations of the first 15 lags; LM(1): Lagrange multiplier test for first order autocorrelation; LM(4): Lagrange multiplier test for fourth order autocorrelation.

Table 6 Cointegration analysis of the M_2 model

Panel A: Rank test

	H_0	$r = 0$	$r \leq 1$	$r \leq 2$	$r \leq 3$
λ_{trace} tests	H_1	$r > 0$	$r > 1$	$r > 2$	$r > 3$
	adjusted statistics[a]	60.66***	25.83	9.75	1.78

Panel B: The full model

	y	m	p	i
cointegration vector ($\hat{\boldsymbol{\beta}}$)	-1.357	1.000	-0.817	0.010
loading vector ($\hat{\boldsymbol{\alpha}}$)	0.256	-0.186	0.039	-0.412

H_0: zero semi-interest-rate elasticity[b]	$\chi^2(1)$=0.74 (0.39)
H_0: weak exogeneity of interest rate[b]	$\chi^2(1)$=0.08 (0.78)
H_0: zero semi-elasticity and weak exogeneity of interest rate[b]	$\chi^2(1)$=0.74 (0.69)

Panel C: The system with long-run exclusion and weak exogeneity of the interest rate

	y	m	p	i
cointegration vector ($\hat{\boldsymbol{\beta}}$)	-1.498	1.000	-0.605	-
loading vector ($\hat{\boldsymbol{\alpha}}$)	0.227	-0.189	0.026	-

H_0: price homogeneity[b]	$\chi^2(1)$=5.98 (0.01)
H_0: weak exogeneity of money[b]	$\chi^2(1)$=9.87 (0.00)
H_0: weak exogeneity of output[b]	$\chi^2(1)$=17.52 (0.0)
H_0: weak exogeneity of prices[b]	$\chi^2(1)$=1.39 (0.24)

	L-B(15): $\chi^2(123)$=125.72 (0.42)		
Autocorrelation[c]	LM(1): $\chi^2(9)$=8.05 (0.53)	Normality[c]	$\chi^2(9)$=147.65 (0.00)
	LM(4): $\chi^2(9)$=10.72 (0.30)		

Notes: [a b c] See notes to Table 5. *** The null is rejected at the significance level of 1%.

Testing assumptions. The first postulate examined is that prices and money vary by the same proportion in the long run. It is rejected for both M_1 and M_2. As stated in section 2, the conventional money demand function specifies the determination of real balances. The estimated cointegration vectors here do not admit of such an interpretation. The demand for real balances is not a meaningful concept in this circumstance, as shocks to the nominal stock of money or to prices will have a smaller effect on money than on prices, thereby changing the quantity of real balances.

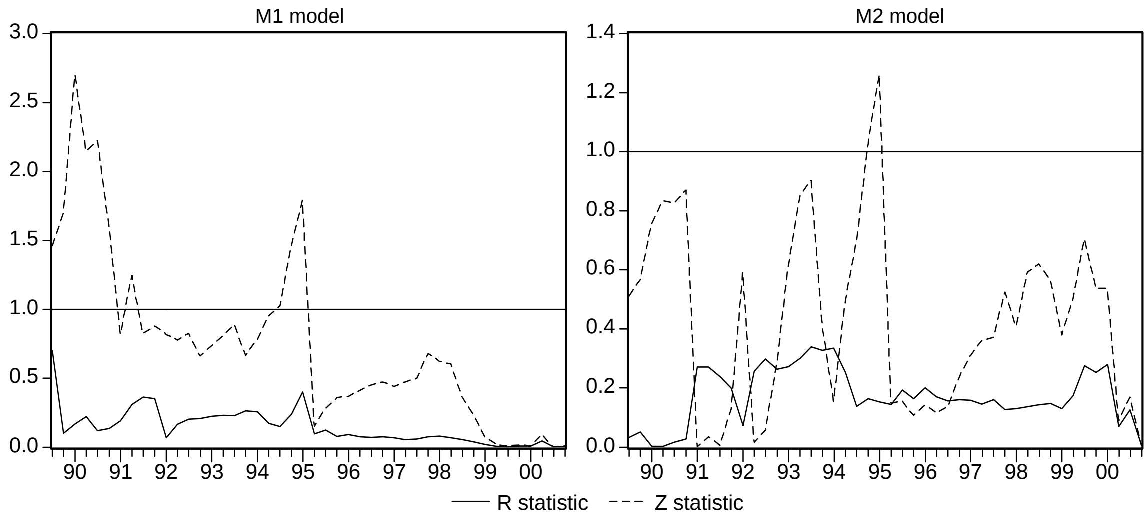
As for the weak exogeneity tests, the weak exogeneity of M_2 is rejected at a low significance level. The evidence against an exogenous M_1 is less strong, but the null can still be rejected at a significance level of 6 per cent. These results resonate with the second view that money accommodates changes in output and prices. The results also lend much support to the weak exogeneity of prices, yet little to the weak exogeneity of real output. A look at the magnitudes of the coefficient estimates also show that y adjusts the fastest among the variables whereas p is the slowest. Thus, changes in nominal demand impact on real output, but hardly on prices.

Comparing the magnitudes of the estimates across the M_1 and M_2 models reveals that the magnitude of the price coefficient in $\hat{\boldsymbol{\beta}}$ is much smaller in the M_1 model than in the M_2 model, suggesting that in the long run prices have a stronger connection with broad money M_2. Note that the coefficient on real output is also larger in the M_2 model. However, the coefficients in $\boldsymbol{\alpha}$ are larger in the M_1 models, implying that the M_1 systems adjust more quickly to disequilibrium errors.

Diagnostic tests. To ensure that the estimated models are statistically adequate descriptions of the data set, multivariate tests for autocorrelation and normality are conducted on the residuals of the modified M_1 and M_2 models. As the bottom lines of Tables 5 and 6 show, both models present little sign of autocorrelation, yet neither pass the normality test. While the normality test results call for interpreting the findings with caution, it shall not severely undermine the validity of the above analysis since the Johansen procedure is robust to non-normally distributed errors (Gonzalo, 1994).

Parameter constancy has been of particular concern to econometric modelling of transition economies. One method of testing for structural stability of the cointegration vector is the multivariate recursive procedure proposed by Hansen and Johansen (1999). The output of applying this test to the modified M_1 and M_2 models is plotted in Figure 1. Two test statistics are calculated for each model: The Z-statistic is computed by re-estimating all parameters in the VEC model for each sub-sample; the R-statistic is obtained by re-estimating the cointegration vectors only while fixing the short-run parameters. The test statistics are scaled by their 95 percentiles before being charted in the graphs so that a value greater than one means the null hypothesis can be rejected at the 5 per cent significance level. The null hypothesis of the test is the cointegration vector estimated over the entire sample lies in the space spanned by the cointegration vectors estimated for each sub-sample. Hence, rejecting the null for a specific sub-sample indicates a possible structural break in the last quarter of that sub-sample. As the two graphs show, the R-statistic stayed well below one in both models for all the sub-samples. However, the Z-statistic for the M_1 model exceeded one in the 1989–91 period and again in 1994. The Z-statistic for the M_2 model also jumped above one in the last two quarters of 1994. These results suggest that the long-run parameters were rather stable in the sample period. There appears to have been some instability in the short-run

Figure 1 Test for constancy of the cointegration vectors



parameters, especially in the M_1 model. Nonetheless, the unstable episodes did not persist long, and overall no structural break occurred.

4.2 The short-run dynamics of money, output and prices

This section attempts a structural interpretation of the conditional M_2 model. The M_1 model is not used for this exercise as the foregoing analysis suggests that it is relatively unstable during the sample period.

The analysis in section 4.1 shows that there is one cointegration relation among the three endogenous variables, *y*, *m*, *p*, of the conditional M_2 model. Thus, the model is driven by two shocks that have permanent effects on the levels of the three variables. In addition to the two 'permanent' shocks, the short-run dynamics of the system is also affected by a 'transitory' shock. As stated in section 3, exact identification of the structural shocks of the model calls for one restriction to separate the two permanent shocks since the transitory shock is already identified.

The reduced-form model is composed of one real variable and two nominal variables. A plausible characterisation of the two permanent shocks might come from the assumption of long-run real/nominal dichotomy. That is, the long-run behaviour of real output is determined solely by real factors such as technology, institutions and factor endowments. The analysis in section 4.1 also indicates that prices are weakly exogenous, which suggests that price shocks might be the source of one of the permanent shocks. Considering that the sample period saw price reforms and nominal devaluations, it is reasonable to assume that there have been independent shocks to price expectations. Hence, an illustrative representation of the structural relationship among money, output and prices may be written as

$$\begin{aligned}
&y_t^* = y_{t-1}^* + \varepsilon_t^1 \\
&E_t p_{t+1} = E_{t-1} p_t - \lambda(y_t^* - y_{t-1}^*) + \varepsilon_t^2 \\
&m_t = a_1 y_t + a_2 p_t + \varepsilon_t^3 \\
&p_t = E_{t-1} p_t + \gamma(y_t - y_t^*)
\end{aligned}, \qquad (5)$$

with the parameters, a_1, a_2, γ, $\lambda > 0$.

The first two equations in (5) describe the evolution of the permanent component of output, y^*, and price expectations. Permanent output is driven by the real shock, ε^1. Price expectations are affected negatively by ε^1, and also by a shock, ε^2, which is independent of the real shock. The third equation can be regarded as the aggregate demand equation, where nominal demand is affected by the transitory shock ε^3. The last equation depicts the aggregate supply schedule. It is easy to obtain that in equilibrium, when $y_t = y_t^*$ and $p_t = E_{t-1} p_t$, the real shock, ε^1, affects all three variables. The price shock, ε^2, has permanent effects on money and prices, but not on output. The third shock has no permanent effects on any of the variables. Full identification of the structural shocks is thus achieved. Applying the identification scheme to the reduced-form M_2 model estimated in 4.1 yields the following results.

Impulse responses. The dynamic responses of real output, prices and M_2 to the three structural shocks appear in Figure 2. The solid lines trace the effects of one standard-deviation shock on the levels of the variables. The one-standard-error confidence intervals of the responses are represented by the dashed lines.[6]

As the graphs in the first column of Figure 2 show, real output completes adjustment to equilibrium in four to five years after a real permanent shock. Nominal balances react quickly to accommodate output expansion, accomplishing more than 80 per cent of its long-run increase on impact. The responses of prices are inaccurately estimated. It seems, though, prices barely move in the wake of a permanent real shock. The near non-response result is in line with the evidence obtained earlier of weakly exogenous prices.[7] For if prices do not adjust when disequilibrium occurs, as implied by weak exogeneity, they will be primarily driven by one of the shocks constituting the common stochastic trends.

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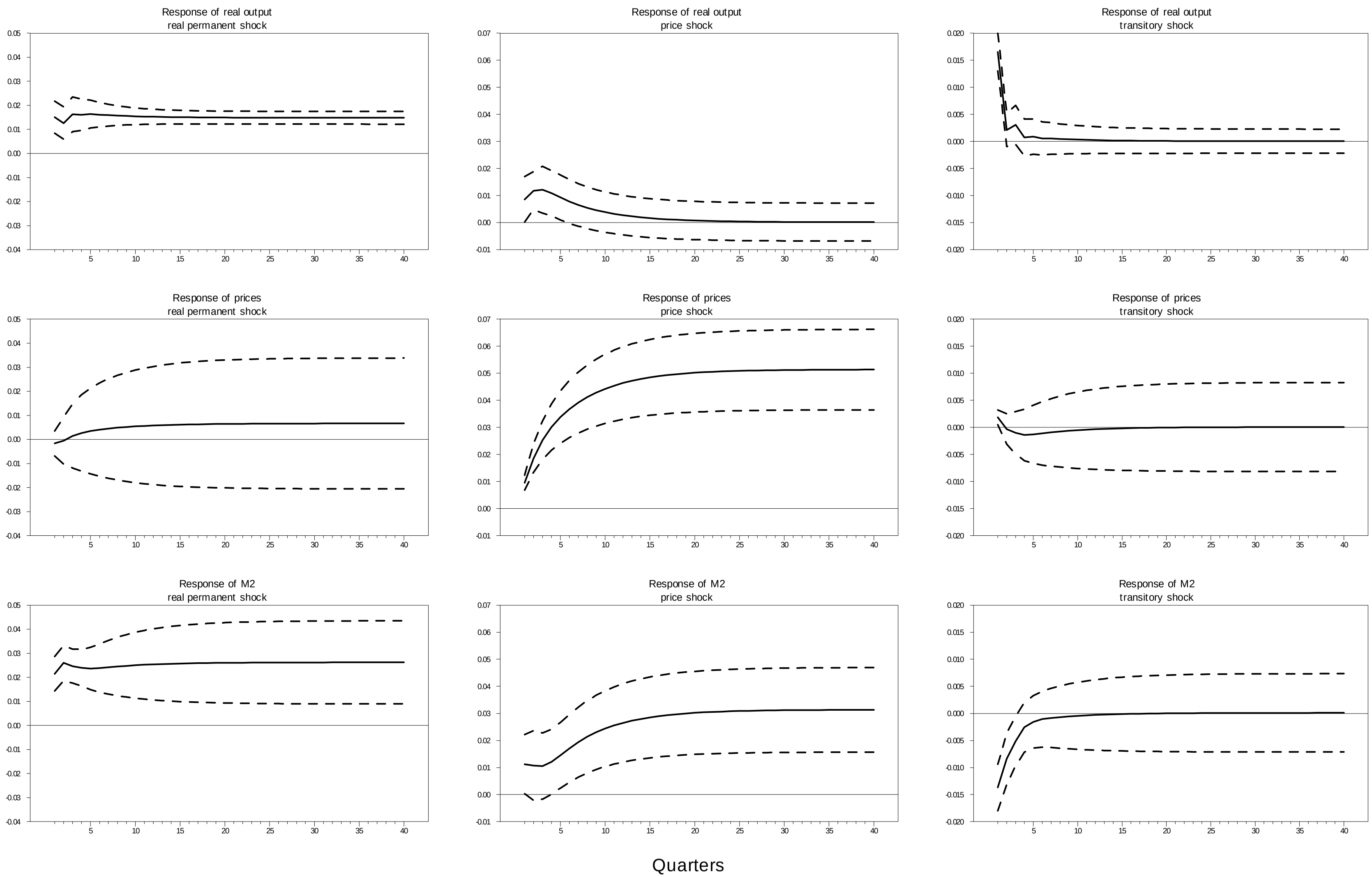
6 The standard errors are based on Monte Carlo simulations of the model using 1,000 repetitions.

7 The impulse response functions are obtained without imposing weak exogeneity restrictions. Doing so allows the use of impulse responses as a check on the robustness of the weak exogeneity results.

Figure 2 Impulse responses to structural shocks



In the second column of Figure 2, the second permanent shock is seen to have a significant positive effect on output for a year or so. The impact dampens out in about five years. Both prices and the stock of money increase in the long run. Yet the response of prices is stronger and swifter than that of money. Nearly 60 per cent of the long-run rise in prices is realised in the first year, compared with less than 40 per cent in the case of money. Therefore, money accommodates changes in prices and output.

The third column displays the responses to the transitory shock. In terms of magnitude, its initial impact on output almost matches that of the permanent real shock. But the effect is short-lived and dissipates quickly in the next two quarters. The changes in the level of prices are negligible, the impact response being a statistically insignificant increase of less than 0.2 per cent.[8] The shock elicits a relatively large temporary reduction in the stock of money. These results are consistent with a scenario where the monetary authorities react quickly to head off inflation pressure arising from positive shocks to aggregate demand. They might also be attributed to a negative shock to money demand, especially if money is endogenous. In the absence of alternative means to hold wealth, a decrease in the demand for money leads to an increase in the demand for real assets, fuelling aggregate demand and thus output.

Variance decomposition. Shocks embody new information. They cause the realised values of the variables in the current period to deviate from the forecasts based on information available up to the previous period. Forecast error variance decomposition provides a measure of the relative contributions of individual shocks to the forecast errors of a particular variable over different forecast horizons. The decomposition results of the conditional M_2 model are presented in Table 7.

As can be seen there, the real permanent shock is the most important factor in determining the evolution of output. Its influence is balanced by the combined effects of the other two shocks at short horizons, particularly within the one-year horizon. The effect of the price shock on prices overwhelmingly dominates those of the other two shocks, explaining more than 95 per cent of the forecast error variance over all horizons. This result corroborates the finding from the impulse response analysis: prices are primarily driven by the shocks comprising the second stochastic trend. By contrast, money is more susceptible to the influence of the real permanent shock through to the five-year horizon, though the price shock eventually becomes the greatest influence. The transitory shock has sizable effects on money in the short run.

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8 It is worth noting that this result is again consistent with prices being weakly exogenous, as transitory shocks do not have a contemporaneous effect on weakly exogenous variables (Fisher and Huh, 1999).

Table 7 Fraction of forecast error variances explained by structural shocks

Horizon	Industrial production			Prices			M_2		
(Quarters)	real shock	price shock	transitory shock	real shock	price shock	transitory shock	real shock	price shock	transitory shock
1	43.77	15.03	41.20	2.16	95.11	2.74	61.74	16.76	21.50
	(22.50)	(18.63)	(14.97)	(24.41)	(24.23)	(4.95)	(23.72)	(23.78)	(12.04)
2	46.80	26.52	26.68	0.49	98.89	0.62	71.27	14.95	13.78
	(24.66)	(22.24)	(10.66)	(23.74)	(23.87)	(2.45)	(24.39)	(24.58)	(9.58)
3	52.12	29.55	18.33	0.46	99.21	0.33	74.74	14.92	10.35
	(26.39)	(24.67)	(8.44)	(23.77)	(24.01)	(2.35)	(24.91)	(25.16)	(8.18)
4	55.79	30.06	14.15	0.67	99.07	0.26	75.74	16.20	8.06
	(27.07)	(25.77)	(6.91)	(23.96)	(24.24)	(2.42)	(25.49)	(25.78)	(6.98)
8	67.10	24.79	8.11	1.24	98.65	0.11	68.63	27.68	3.69
	(26.69)	(25.94)	(4.77)	(24.47)	(24.75)	(2.44)	(27.42)	(27.74)	(4.28)
20	82.02	13.83	4.14	1.63	98.34	0.03	51.83	47.14	1.03
	(22.84)	(22.41)	(3.43)	(24.82)	(25.06)	(2.40)	(29.74)	(29.95)	(2.78)
40	89.78	7.87	2.35	1.75	98.24	0.01	45.47	54.11	0.43
	(21.42)	(21.09)	(3.03)	(24.91)	(25.14)	(2.40)	(30.63)	(30.78)	(2.61)
60	92.86	5.50	1.65	1.78	98.21	0.01	43.65	56.08	0.27
	(21.59)	(21.28)	(2.95)	(24.94)	(25.16)	(2.40)	(30.92)	(31.05)	(2.59)

Note: Entries in parentheses are sample standard errors computed by Monte Carlo simulation using 1,000 repetitions.

4.3 Discussion

Although the People's Bank of China (PBC) – China's central bank – did not formally abandon the annual credit plan until 1998, it actually started targeting monetary aggregates in the early 1990s. Crucial to the success of a quantity-targeting regime are two empirical regularities. The first is the controllability of the targeted monetary aggregates. The results of this study suggest that prices hardly respond to deviations from equilibrium, while money (and real output) adjusts to eliminate disequilibrium.Therefore, nominal balances are endogenously determined by factors affecting the demand for money. The monetary authorities are either unable or unwilling to impose restrictions on the availability of liquidity. A quantity-targeting monetary regime would not be practical.

The second empirical regularity required for quantity targeting is a predictable empirical relationship between money and the other variables of interest, in particular prices and real output. The cointegration analysis shows that two of the monetary aggregates, M_1 and M_2, are cointegrated with real output, prices, and, in the case of M_1, the interest rate. Moreover, the estimated long-run relations are reasonably stable over the sample period. Hence, despite its endogeneity, the stock of money still provides valuable information about the changes in nominal demand. In the long run, price stability and stable money growth go hand in hand.

As for which monetary aggregate should be given more emphasis as an information variable, M_2 appears preferable to M_1. The parameter constancy tests indicate greater stability in the relationship of M_2 with output and prices. Judging from the coefficient estimates in the cointegration vectors, M_2 also has a stronger connection with the other variables. In view of the current rapid expansion of the stock and bonds markets and likely future reforms such as opening up of the capital market to foreign investors, however, additional caution must be taken in interpreting future movements in M_2. The presence of the interest rate variable in the cointegration vector for M_1 suggests that the long-run relation between money and other variables is indeed under the influence of agents' portfolio decisions. That the M_2 model does not include a rate of return variable in its long-run relation shall be attributed to the limited availability of substitutes for cash and bank deposits. With greater supply of domestic bonds and shares and easier access to foreign currency and assets, the stability of the M_2 model may start to unravel.

The findings of this study also help to clarify a few issues in the empirical modelling of the relationship between money and other variables. To start with, price homogeneity, as commonly presumed in the studies estimating the money demand function, is a questionable assumption. It does not hold in either the M_1 model or the M_2 model. Consequently, a model of the demand for real balances seems empirically irrelevant for the sample period. Next, since prices are found to be weakly exogenous whereas money endogenous, it is generally not valid to take prices as a function of money as is practised in Chow (1987) and Hasan (1999).[9] In a related manner, it is also inappropriate to use

[9] However, data and sample periods in these studies are different from those in the current study.

the estimated long-run relations to forecast the short-run demand for nominal balances because of the endogeneity of real output.

5 Conclusion

This paper examines the alternative views about the role of money in the movements of output and prices in China's reform period. The structural VEC modelling approach is adopted. Based on the equation of exchange, the long-run relationships relating real output and prices to three monetary aggregates are estimated using the Johansen cointegration procedure. Extensive tests are conducted to arrive at an empirically sound reduced-form VEC system. The short-run structure of the model is identified by implementing the division between permanent and transitory shocks as implied by the cointegration analysis, and by imposing a recursive structure on the permanent shocks. The findings bring out several characteristics of the monetary aspects of China's output and price fluctuations in the sample period.

First, money accommodates changes in output and prices, rather than causing them. Second, money and prices do not vary in proportion even in the long run. Third, changes in nominal interest rates induce portfolio shifts, but not alterations in expenditure decisions. Finally, shocks to price expectations are the predominant causal factor in price fluctuations. They also strongly influence real output, and account for a large proportion of fluctuations in money.

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