Demand for Money in India: 1953-2002*

B. Bhaskara Rao and Rup Singh

University of the South Pacific, Suva (Fiji)

Abstract

The demand for money, especially in the developing countries, is an important relationship for formulating appropriate monetary policy and targeting monetary variables. In this paper we estimate the demand for narrow money in India and evaluate its robustness. It is found that there is a stable demand for money for almost half a century from 1953 to 2002. However, there was a shift from the demand towards time deposits during 1979-83 and its effects can be adequately captured by the error-correction model, instead of a shift dummy. There is no evidence for any significant effects of the 1991 financial reforms.

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**KEYWORDS:** Demand for money, Developing countries, Income and interest rate elasticities, Cointegration, Financial reforms.

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1. INTRODUCTION

The demand for money function is probably the most widely researched topic. Econometric estimates of this function abound in the developed and developing countries. Parametric estimates in myriad developed and developing countries look similar. In general, estimates of the income elasticities are close to unity and interest rate elasticities are small, negative and often insignificant in many developing countries; see Sriram (1999) for a recent survey. In this paper, we take a fresh look at the demand for money of a large developing country viz., India. Our study shows that there is a well defined and stable demand for narrow money ($M_1$) for India for half a century, from 1953 to 2002. Our estimates, based on the unit roots and cointegration methodology, show that both the income and interest elasticities of the demand for $M_1$ are significant and close to some earlier estimates of Rao and Shalabh (1995). The outline of this paper is as follows: In Section 2 specification and definitional issues are examined. Sections 3 and 4 present empirical results and investigate stability and robustness of our estimates. Finally in Section 5, conclusions and summary are given.

2. SPECIFICATION AND DATA

We start with a standard and well-trodden specification of the demand for real narrow money ($M_1/P$):

\[ \ln \left( \frac{M_t}{P_t} \right) = \alpha_0 + \alpha_1 \ln Y_t + \alpha_2 i_t + \epsilon_t \] (1)

where, $M$ is narrow money consisting of currency plus demand deposits, $P$ is the GDP deflator, $Y$ is real GDP, $i$ is
a nominal rate of interest to capture the cost of money and ε is the error term with the standard classical properties.

In the empirical works on the developing countries, there seems to be some confusion about whether the interest rate variable should be a nominal or a real rate. Generally nominal rates show less variation in the developing countries and their coefficients are usually insignificant in the money demand functions. Since real rates show more variation, mainly due to the larger variation in the inflation rate, the real rate is mistakenly thought to be a better explanatory variable. Inclusion of the real rate of interest implies the counter intuitive result that the demand for real balances increases with the expected rate of inflation. In addition, some investigators may have mistaken that since the dependent variable is measured in real magnitudes, the explanatory variables should be also in real terms. The drawback of such formulations is that when, for example in equation (1), the real rate is included and the expected inflation is proxied with the lagged inflation rate ($\Delta P_{t-1}$), $\alpha_2(i_t - \Delta P_{t-1})$ will be the coefficient of the cost of money. Since this should have a negative effect, $\alpha_2$ is negative and implies that the real demand for money increases with the inflation rate. The correct interest rate variable is, therefore, the nominal rate of interest.

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1 We draw attention to recent publications with the real rate of interest. An early study by the IMF and a recent one by Jayaraman and Ward (2000) have used the real rate of interest in the demand for money for Fiji. Ahmed (2001) has also used the real rate for Bangladesh. Perhaps there are several other empirical works in the developing countries with similar weaknesses in the demand for money. For obvious reasons such studies seldom appear in the well-known journals.
Inclusion of the real rate, along with other nominal rates and the expected rate of inflation, is perhaps justified if substitution between money and real assets, e.g., real estate, is important. If there are high inflation periods in the data, it is appropriate to include both the nominal rate of interest and the expected rate of inflation, without constraining their coefficients to be equal and opposite in sign. Expected rate of inflation also proxies the negative cost of holding money; see Friedman (1969) and Sriram (1999). It is to be expected that the coefficient of the expected rate of inflation would be negative, not positive.

The definitions of the variables in this study are as follows: Nominal (narrow) money ($M$) is currency with the non-bank public, demand deposits and other deposits with the Reserve Bank of India. $Y$ is real $GDP$ at factor cost in 1993 – 94 prices. $P$ is the implicit $GDP$ deflator and $i$ is the average rate of interest on one to three year time deposits. Data on the monetary variables are from various issues of *The Currency and Finance Report* (Mumbai: Reserve Bank of India) and also downloaded from the Home Page of the Reserve Bank of India. Data on $GDP$ and $P$ are from various issues of *The Economic Survey* (New Delhi: Government of India) and also downloaded from the Home Page of the Ministry of Finance, Government of India. Our sample period covers from 1952 to 2002.

3. ESTIMATES WITH PARTIAL ADJUSTMENT

The often used specification in equation (1) was found to be quite adequate to explain the demand for narrow money in many developing countries; see Sriram (1999). Prior to the current popularity of the VAR methodology, OLS equations based on partial adjustment were popular. Therefore,
we start with the estimation of a partial adjustment equation. This also serves to illustrate the usefulness of the VAR modeling. Equation (2) below reports the results.

\[
\ln\left(\frac{M_t}{P_t}\right) = -3.223 + 0.453\ln Y_t - 0.007i_t + 0.611\ln\left(\frac{M_{t-1}}{P_{t-1}}\right)
\]

\( (4.50) \ast (4.42) \ast (1.99) \ast (6.60) \ast \)

\[R^2 = 0.997, DW = 1.775, h = 1.052, SEE = 0.038 \]

Period : 1953 – 2002

\[\chi^2_{sc1} = 0.864, \chi^2_{ff} = 0.028, \chi^2_{hs} = 3.107, \chi^2_n = 0.331 \]

t-ratios are in the parentheses and * indicates significance at the 5% level. All coefficients in the above equation have the expected signs and are significant. The \(\chi^2\) statistics are for the null hypotheses that the first order serial correlation (\(\chi^2_{sc1}\)), functional form misspecification (\(\chi^2_{ff}\)), non-normality of the residuals (\(\chi^2_n\)) and heteroscedasticity in the residuals (\(\chi^2_{hs}\)) are absent. These are all less than the 95% critical value of 3.84. Therefore, the null hypotheses are accepted.

However, TIMVR tests for its stability indicate that while the test with the sum of recursive residuals shows stability, the sum of squares test indicates mild instability during 1978 – 1983. This indicates that there was no gradual shift but perhaps a sudden and haphazard shift in this function. This was perhaps due to a significant shift from the demand to time deposits during this period because there were significant increases in the time deposit rates, beginning from 1979, or due to errors in the measurement of demand and time deposits. Time deposits, as a proportion of
broad money ($M_3$), increased from an average of 37.5% during 1973 to 1978 to 59% in the subsequent 5 years. Interest rates on time deposits increased from an average of 8.5% to 11% between these two periods.

It is usual to explain such shifts with a dummy variable. Reestimation of equation (2) with a shift dummy (unity during 1979 to 1983 and zero for all other years) showed that the coefficient of the dummy variable is significant. Nevertheless, temptation to include a dummy variable should be postponed until the dynamic adjustments are adequately investigated, for example within the VAR framework instead of the simple partial adjustment in equation (2); see Hendry and Ericsson (1991) and Sriram (1999).

Furthermore, the residuals of this partial adjustment equation are found to contain unit root and implies that its summary statistics are biased. Therefore, we test for the presence of unit roots in the variables and use the cointegration and ECM approach for estimation.

4. COINTEGRATION AND ECM FRAMEWORK

The Dicky-Fuller tests are now a standard procedure for testing for the order of the variables. The computed test statistics for the levels and first differences of the variables are given in Table 1 below.

For the levels of the three variables the null hypothesis of unit roots cannot be rejected at the 95% level, but the null that their first differences have unit roots is clearly rejected. Therefore the level variables are $I(1)$ and can be modelled within the VAR framework.

Tests for the selection of the order of the VAR model clearly favoured the first order. The Akaike Information
Table 1

ADF Tests for unit roots:
Levels and first differences of variables
with intercepts and linear trends in levels

<table>
<thead>
<tr>
<th>Variable</th>
<th>Period</th>
<th>m</th>
<th>Test Statistic</th>
<th>95% CV</th>
</tr>
</thead>
<tbody>
<tr>
<td>ln(M/P)</td>
<td>1955–2002</td>
<td>2</td>
<td>−0.271</td>
<td>−3.505</td>
</tr>
<tr>
<td>∆ln(M/P)</td>
<td>1956–2002</td>
<td>2</td>
<td>−5.855∗</td>
<td>−3.507</td>
</tr>
<tr>
<td>lnY</td>
<td>1955–2002</td>
<td>2</td>
<td>−0.021</td>
<td>−3.505</td>
</tr>
<tr>
<td>i</td>
<td>1955–2002</td>
<td>2</td>
<td>−1.924</td>
<td>−2.920</td>
</tr>
<tr>
<td>∆i</td>
<td>1956–2002</td>
<td>2</td>
<td>−4.055∗</td>
<td>−3.503</td>
</tr>
</tbody>
</table>

Notes: Significance at 95% level is indicated with *. A time trend was not included in the first differences of the variables.

Criterion (AIC) and Schwarz Bayesian Criterion (SBC) statistics reached a maximum of 132.298 and 121.197 for the first order. For the second order VAR, AIC and SBC are 128.234 and 106.807 respectively. We tested for the number of cointegrating vectors among the variables using the Johansen maximum likelihood procedure in Microfit 4.1. This procedure has by now become the standard method, and we avoid tabulating the details of these results. The null hypothesis that there are no cointegrating vectors is rejected but the null that the number of cointegrating vectors is at least one is not rejected.3

3 The maximal eigenvalue and trace test statistics for the null that there is no cointegration are 25.334 and 35.918 respectively. The 95% critical values, respectively, are 21.120 and 31.540. For the null that there is at least one cointegrating vector, the corresponding computed
The cointegrating vector, normalised on $\ln(M_t/P_t)$, is given below:

$$ln\left(\frac{M_t}{P_t}\right) = 1.1850 \ln Y_t - 0.023 \, i_t$$

(3)

(27.21) * (3.43)*

Asymptotic $t$-ratios are in the parantheses and * indicate significance at the 95% level. The coefficients have the expected signs and are significant. The estimated long run income elasticity of demand for money is about 1.2% and interest rate elasticity, at the mean rate of interest of 7.65, is $-0.18$. These are comparable to but different from the earlier estimates by Rao and Shalabh (1995) for the period 1952 – 1992. In that study income elasticity was 1.5 and the interest rate elasticity was $-0.420$. Thus, both these elasticities seem to have decreased during the 1990s. The decline in the interest rate elasticity might be due to more flexible and market oriented interest rates on time deposits which might have induced the shift from demand to time deposits in the 1980s. Therefore the interest rate sensitivity of the balance of the hard core demand deposits might have reached a bottom. A rolling regression, with the partial adjustment equation, indicated that the coefficient of the rate of interest showed more fluctuations in the pre1979 period, values, with the critical values in the parentheses, are: 10.559 (14.880) and 10.5849 (17.860) respectively. In conducting these tests, we allowed for an unrestricted intercept but no trend. When a restricted trend was also included, there was no difference in the results based on the trace statistic but the asymptotic standard error of the trend was insignificant in the cointegrating equation. Therefore the trend variable is ignored.
reaching a maximum in 1979. Since then it slowly decreased and became stable after 1995.\(^4\)

To determine whether causation runs from income and the rate of interest to money, we tested for cointegration, using the option in *Microfit* by interchanging the dependent and independent variables; see Pesaran and Pesaran (1997). The hypothesis that \(r = 1\) is accepted only when money was treated as the dependent variable. In the other two versions the null that \(r = 0\) could not be rejected. We also estimated two bi-variate *VAR* models with the *SURE* method. In the first one, real money and output are endogenous variables and in the second real money and interest rate are endogenous. When these are estimated with the constraint that the coefficient of the lagged residual of the cointegrating equation is zero in the equation for output (model 1) and the rate of interest (model 2), the null hypotheses are easily accepted. Therefore it can be said that both output and interest rate Granger cause money.

In developing an appropriate error correction model (*ECM*) for the short run, we started with a very general specification in which \(\Delta(M_t/P_t)\) is regressed on its lagged values, the current and lagged values of \(\Delta Y_t, \Delta i_t\) and \(ECM_{t-1}\) where the last term is the lagged residuals from the cointegrating equation (3). This is a tedious procedure because in our initial estimation we have used lags up to 4 periods. By using the standard variable deletion tests,

\(^4\) We have also tried to estimate the cointegrating equations for the pre and post 1979 periods, but found that it is not possible to obtain a satisfactory cointegrating relationship for the first period.
we arrived at the following parsimonious ECM version.

\[
\ln\left(\frac{\Delta M_t}{P_t}\right) = -3.060 - 0.360 ECM_{t-1} + 0.370 \Delta \ln Y_t
\]

\[
(3.84) \ast \quad (3.89) \ast \quad (2.28) \ast
\]

\[-0.012 \Delta i_{t-2} - 0.129 \Delta i_{t-4}
\]

\[
(1.66) \ast\ast \quad (1.74) \ast\ast
\]

\[R^2 = 0.356, \ SEE = 0.048, \ LLH = 95.270\]

Period: 1956 – 2002

\[\chi^2_{sc1} = 0.927, \chi^2_{ff} = 0.123, \chi^2_{hs} = 0.196, \chi^2_{n} = 0.143\]

where \(*\) and \(*\ast\) indicate significance at the 5% and 10% levels respectively. All the estimated coefficients are significant and the ECM variable, since its coefficient is negative, serves as a negative feedback mechanism. This implies that if there are departures from equilibrium in the previous period, this departure is reduced by about 36% in the current period. The \(\chi^2\) statistics indicate that there is no serial correlation, functional form misspecification, non-normality and heteroscedasticity in the residuals. Therefore, our ECM model is satisfactory.

We subjected equation (4) to TIMVAR stability tests and found that both the tests based on the recursive residuals indicated stability. Thus the ECM formulation can be said to have captured the dynamics underlying the money demand function better than the partial adjustment equation (2). Similar results on the relative merits of the ECM modeling over the partial adjustment formulations were obtained for the USA and the UK; see Hendry and Ericsson (1991).

Economic reforms in India started from 1991 and to examine their impact we introduced a reforms dummy (unity
from 1991 to 2002 and zero otherwise) and reestimated equation (4). A similar procedure was used for the Philippines by Haper and Kutan (2003) and the coefficient of their dummy variable, in the narrow money equation, is positive and significant at the 10% level. This indicates that reforms in the Philippines have mildly increased perhaps bank deposits and their use in transactions. For India the coefficient of this reforms dummy while negative, is almost zero (−0.0015) and insignificant. While reforms did not have any significant effects on narrow money in India, it is likely that they might have increased saving and time deposits and therefore such effects are worth investigation in the demand for broader money. In most transactions in India cheques drawn on demand deposits are not popular. In fact even to withdraw cash from a time deposit account is not convenient because of long delays in processing the cheques. Therefore, our finding that reforms had an insignificant effect on narrow money is not unexpected.

5. CONCLUSIONS

In this paper we have shown that the variables in the demand for money in India are non-stationary in their levels but stationary in their first differences. Therefore, standard partial adjustment based specifications in the levels of variables would be unsatisfactory and the demand for money should be modeled within the VAR framework. Our estimates using the VAR methodology imply that there is a well determined and stable demand for money for half a century. Therefore, in comparison to partial adjustment equations, it should have better properties and predictive performance.

Our estimates imply that both the income and interest rate elasticities are significant, have the expected signs
and are consistent with their expected magnitudes. Income elasticity is about 1.2 and interest elasticity is about $-0.2$. Nevertheless, it should be noted that income and interest rate elasticities implied by the partial adjustment equation are close to those yielded by the VAR model. The latter model is better in the sense that it captured the dynamic adjustment process far better than the partial adjustment equation. Therefore, estimates based on the VAR model are more appropriate for policy formulation.

We believe that our study is perhaps the first attempt to estimate and test the demand for money of a developing country for such a long period. Our finding that the demand for money, based on VAR estimates, is stable is in contrast to some recent findings that the demand for money in several countries has become unstable due to financial innovations and reforms. This has lead many central banks to switch from targeting money supply to interest rate, since it is well known that targeting interest rate is more appropriate when demand for money is unstable; see Poole (1970). However, there does not seem to be any need for shifting from targeting narrow money to interest rate targeting in India, unless financial reforms are found to destabilize the demand for money.

With the benefit of our results it would be useful to investigate the effects of targeting narrow money on nominal income and/or real income and price level as well as the implications of targeting money supply for seignorage revenue to finance the budget deficits. Furthermore, it would be also valuable to investigate the nature and stability of the demand for broader money ($M3$) within the VAR framework. These issues fall outside the scope of our current paper.
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