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Demand for Money in India: 1953-2003*

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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 2003. 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 and Parikh (1994) for a brief survey of various empirical works on India. 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 ($M1$) for India for half a century, from 1953 to 2003. Our estimates, based on the unit roots and cointegration methodology, show that both the income and interest elasticities of the demand for $M1$ are significant and close to some earlier estimates. 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 ($M1/P$), based on the partial adjustment model (PAM):[1]

$$ln\Big(\frac{M_t}{P_t}\Big) = \lambda\alpha_0 + \lambda\alpha_1 lnY_t + \lambda\alpha_2 i_t + (1-\lambda) ln\Big(\frac{M_{t-1}}{P_{t-1}}\Big) + \epsilon_t \quad (1)$$

[1] Equation (1) is derived from an equilibrium relationship:

$$ln\Big(\frac{M_t^*}{P_t^*}\Big) = \alpha_0 + \alpha_1 lnY_t + \alpha_2 i_t$$

augmenting with a partial adjustment equation:

$$ln\Big(\frac{M_t}{P_t}\Big) - ln\Big(\frac{M_{t-1}}{P_{t-1}}\Big) = \lambda\Big[ln\Big(\frac{M_t^*}{P_t^*}\Big) - ln\Big(\frac{M_{t-1}}{P_{t-1}}\Big)\Big]$$

where λ is the speed of adjustment. See Taylor (1994) and Cuthbertson (1988).

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 empirical work on the developing countries, there seems to be some confusion about whether the interest rate variable should be a nominal or a real rate.[2] 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. Some investigators may have also mistaken that since the dependent variable is measured in real magnitudes, the explanatory variables should be also in real terms. 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.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 ($ln\Delta P_{t-1}$), the effect of the cost of money is given by $\alpha_2(i_t - ln\Delta P_{t-1})$. Since this should have a negative effect, α_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.

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., consumer durables, gold and other precious metals, real estate etc., 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

[2] 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.

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* (Mumbay: 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 is from 1952 to 2003.

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 PAM were popular. Therefore, we start with the estimation of a PAM equation. This also serves to illustrate the usefulness of the VAR modeling because compared to PAM the general to specific approach based VAR specifications allow for a more flexible dynamic lag structure that is consistent with the underlying data generating process. Equation (2) below reports the results for PAM based demand for money.

$$\begin{aligned} ln\Big(\frac{M_t}{P_t}\Big) =& -3.243 + 0.454 lnY_t - 0.007 i_t + 0.612 ln\Big(\frac{M_{t-1}}{P_{t-1}}\Big) \\ & \quad (4.50)* \quad\;\; (4.42)* \quad\;\; (1.99)* \quad\;\; (6.60)* \end{aligned} \tag{2}$$

$$\overline{R}^2 = 0.997, DW = 1.776, h = 1.052, SEE = 0.037$$

$$Period : 1953 - 2003$$

$$\chi^2_{sc1} = 0.864, \chi^2_{ff} = 0.028, \chi^2_{hs} = 3.107, \chi^2_{n} = 0.331$$

$t-$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 χ^2 statistics are for the null hypotheses that the first order serial correlation (χ^2_{sc1}), functional form misspecification (χ^2_{ff}), non-normality of the residuals (χ^2_n) and heteroscedasticity in the residuals (χ^2_{hs}) are absent. These are all less than the 95% critical value of 3.84. Therefore, the null hypotheses are accepted.

It is of interest to note that this simple equation, based on a somewhat obsolete PAM, looks impressive on the basis of conventional criteria. In particular, unlike in the US and UK demand for money functions based on PAM, there is no serial correlation in the residuals and the coefficient of the lagged money term at 0.612 is significantly less than one. In the Wald test, that this coefficient is one, the χ^2 test statistic is 17.684 with a p value of 0. Taylor (1994) examines, in detail, the serial correlation problem and the near unit and significant value for the coefficient of the lagged dependent variable in the PAM specifications. He rejected the Goodfriend (1985) hypothesis that serial correlation in the PAM specifications is due to errors in the measurement of the explanatory variables–essentially due to errors in measuring income because data on the rate of interest are relatively error free. Taylor's results imply that the PAM specification is restrictive and does not adequately capture the underlying dynamic adjustments in the demand for money. In another development to overcome the serial correlation problem in PAM specifications, Haache (1974) estimated the demand for broad money ($M3$) in the UK in the first differences of all the variables. However, this was quickly dismissed by Hendry and Mizon (1978) as *ad hoc*. It may be said that the Hendry and Mizon work has literally put an end to PAM based demand for money estimates in the UK; see also Cuthbertson (1988) for a useful survey of the demand for money based on PAM.[3]

[3] It is well known by now that Hendry and Mizon have popularized the London School of Economics approach to modeling dynamic equations with the general to specific approach (GETS). GETS is a pragmatic and flexible solution (compared to PAM) to reconcile a methodological conflict between the equilibrium nature of theoretical relationships and the data from the real world that is seldom in a state of equilibrium. Although the LSE-GETS approach had a mixed reception, there is a renewed interest due to the development of an automated model selection software, PcGets, by Hendry and Krolzig (2001) and the seminal contributions by Hoover and Perez (1999, 2004). See also Rao and Singh (2005)

In addition to the aforesaid reasons and the fact that (a) the CUSUM and CUSUMSQ tests for the stability of our estimated equation in (2) showed temporal instability and (b) the residuals of (2) contain a unit root made us to conjecture that (a) PAM specifications are unlikely to satisfactorily capture the underlying dynamic adjusts and (b) the estimated goodness of fit statistics that are impressive for the PAM are somewhat spurious. A similar instability in the demand for money for India was also recently found by Bahmani-Oskooee and Rehman (2005) although they have used a more flexible dynamic lag structure than PAM. However, Bahmani-Oskooee and Rehman have used quarterly data and their scale variable was an index of industrial production since quarterly data on GDP for India are not available. There seem to be two problems with their results. Firstly, their data are not seasonally adjusted and secondly, industrial production may not be a satisfactory scale variable in a developing country like India where the coverage of the industrial production index is limited. Therefore, it would be useful to reestimate the demand for money in India with a more flexible dynamic lag structure than PAM and by using GDP as the scale variable.

4. COINTEGRATION AND ECM FRAMEWORK

Before we proceed further it is necessary to test for the presence of unit roots in our variables.[4] The Augmented Dicky-Fuller tests (ADF) are now a standard procedure for testing for the order of the variables. However, there is a problem with our variables. Not only there is a trend in their levels but, unlike in many cases, the rates of growth of these variables also have trends. Therefore when applying ADF tests the relevant test statistics for both the levels and first differences of the variables are those with an intercept and trend. 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

on PcGets.

[4] Some early studies based on the unit roots and cointegration approach to demand for money in India are: Nag and Upadhyay (1993), Ghatak and Ghatak (1994), Parikh (1994) and Rao and Shalabh (1995). However, Vasudevan (1977) and Bhoi (1992) have used equations based on the PAM.

Table 1

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

Variable	Period	m	Test Statistic	95% CV
$ln(M/P)$	1952 – 2003	0	-1.201	-3.505
$\Delta ln(M/P)$	1956 – 2003	3	$-5.575*$	-3.511
lnY	1953 – 2003	1	-0.610	-3.505
ΔY	1955 – 2003	0	$-8.850*$	-3.507
i	1954 – 2003	2	-2.150	-3.505
Δi	1955 – 2003	1	$-6.174*$	-3.507

Notes: m is the lag length of the first differences of the variable in the ADF equations. Significance at 95% level is indicated with *. A time trend is included in the levels and first differences of the variables because in both sets of variables the coefficients of trend are significant.

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 modeled within the VAR framework.

Tests for the selection of the order of the VAR model clearly favoured the first order. The Akaike Information Criterion (AIC) and Schwarz Bayesian Criterion (SBC) reached a maximum of 134.646 and 120.770 for the first order. For the second order VAR, AIC and SBC are 131.842 and 109.640 respectively. Therefore we estimated an unrestricted first order VAR model of the three variables $ln(M/P), lnY$ and i with the deterministic variables viz., an intercept and a trend. This model is used to conduct the validity of a few restrictions.

First, the significance of the trend variable is tested. The restriction that the coefficient of the trend is insignificant is rejected. The computed $\chi^2(3)$ test statistic is 11.579 and the 95% critical value is 7.815. Therefore, the trend variable is retained for the time being although its coefficient turned out to be insignificant in the later cointegration tests.

Second, we also conducted the Granger non-causality tests; see Enders (2004) and Pesaran and Pesaran (1997). The computed $\chi^2(2)$

test statistic for the null hypothesis that $ln(M/P)$ does not Granger cause i and lnY is 4.953 and the 95% critical value is 5.991. Therefore, the null is accepted.[5] This is partly plausible because interest rates on the fixed deposits are very much set by the Reserve Bank of India. In our subsequent analysis we have assumed that i is an exogenous variable.

The test for determining the number of cointegrating vectors is conducted with the Johansen maximum likelihood procedure in $Microfit$ 4.1. First, an unrestricted intercept and a restricted trend are retained and i is treated as an exogenous variable. Since these tests have become well-known by now, we avoid tabulating the details of the results. The null hypothesis that there are no cointegration is rejected but the null that the number of cointegrating vectors is one is not rejected.[6]

The cointegrating vector, normalized on $ln(M_t/P_t)$, is given below:

$$ln\left(\frac{M_t}{P_t}\right) = \underset{(3.96)*}{1.1866\, lnY_t} - \underset{(1.44)}{0.023\, i_t} + \underset{(0.013)}{0.0002\, Trend} \qquad (3)$$

Asymptotic t-ratios are in the parentheses and * indicate significance at the 95% level. The coefficients have the expected signs, but only the coefficient of real income is significant at the 95% level. The coefficient of the trend is highly insignificant and this may be due to the highly correlated common trends in both real money and real income. Since equation (3) (also equation (4) below) is a long run equilibrium relationship and does not incorporate the short run dynamics, the validity and use of these t-ratios are doubtful. However, we have reported these asymptotic t-ratios from the $Microfit$ out-

[5] This test is essentially testing the restriction that the coefficients of the lagged values of $ln(M/P)$ in the block of equations explaining the variables lnY and i are zero.

[6] The maximal eigenvalue and trace test statistics for the null that there is no cointegration are 25.2268 and 34.5347 respectively. The 95% critical values, respectively, are 22.16 and 30.77. For the null that there is one cointegrating vector, the corresponding computed values, with the critical values in the parentheses, are: 9.3079 (15.44) and 9.3079 (15.44) respectively.

put.

Since the coefficient of the trend is negligible and also highly insignificant, we also tested for cointegration without a deterministic trend and found the following cointegrating relationship:

$$ln\left(\frac{M_t}{P_t}\right) = 1.1899\ lnY_t - 0.023\ i_t \qquad (4)$$
$$(27.93)* \qquad (3.33)*$$

It is noteworthy that removal of trend did not change the coefficient estimates and now the coefficient of the rate of interest is also significant. Nevertheless, it should be kept in mind that deleting the trend, because its coefficient is insignificant in the cointegrating equation, is somewhat an arbitrary procedure. That may lead to instability in the residuals of the error correction equation, in the model without the trend. As our subsequent results show the errors in the ECM, based on equation (4), do not have significant $ARCH$ effects.

The estimated long run income elasticity of demand for money is about 1.2% in both equations (3) and (4) and similarly the 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 PAM equation, indicated that the coefficient of the rate of interest showed more fluctuations in the pre 1979 period, reaching a maximum in 1979. Since then it slowly decreased and became stable after 1995.[7]

In estimating the error correction model (ECM) for the short run, we used both the cointegrating equations (3) and (4) and ob-

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[7] 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.

tained very similar error correction equations. 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 ΔY_t and Δi_t and the lagged error of the corresponding cointegrating equation. We have used lags up to 4 periods on money, income and rate of interest. $(ECM_3)_{t-1}$ and $(ECM_4)_{t-1}$ are the lagged residuals from the cointegrating equations (3) and (4). By using the standard variable deletion tests, we arrived at the following parsimonious ECM equations.

$$
\begin{aligned}
\Delta ln\Big(\frac{M_t}{P_t}\Big) = &- 3.186 - 0.373(ECM_3)_{t-1} + 0.211\ \Delta ln(M_{t-1}/P_{t-1}) \\
&(4.15)*\ \ (4.20)* \qquad\qquad (1.84)** \\
&+ 0.333\Delta lnY_t - 0.010\Delta i_{t-2} - 0.011\Delta i_{t-4} \\
&(1.84)** \qquad (1.40) \qquad (1.54)
\end{aligned} \tag{5}
$$

$$\overline{R}^2 = 0.390,\ SEE = 0.033,\ LLH = 97.20$$

$$Period : 1956 - 2003$$

$$\chi^2_{sc1} = 2.02, \chi^2_{ff} = 1.49, \chi^2_{hs} = 0.005, \chi^2_{n} = 1.853$$

$$
\begin{aligned}
\Delta ln\Big(\frac{M_t}{P_t}\Big) = &- 3.201 - 0.373(ECM_4)_{t-1} + 0.210\ \Delta ln(M_{t-1}/P_{t-1}) \\
&(4.15)*\ \ (4.20)* \qquad\qquad (1.83)** \\
&+ 0.334\Delta lnY_t - 0.010\Delta i_{t-2} - 0.011\Delta i_{t-4} \\
&(2.14)* \qquad (1.40) \qquad (1.54)
\end{aligned} \tag{6}
$$

$$\overline{R}^2 = 0.390,\ SEE = 0.033,\ LLH = 97.10$$

$$Period : 1956 - 2003$$

$$\chi^2_{sc1} = 2.01, \chi^2_{ff} = 1.47, \chi^2_{hs} = 0.005, \chi^2_{n} = 1.847$$

where * and ** indicate significance at the 5% and 10% levels respectively. It is obvious that but for very small differences in the coefficient estimates and summary statistics, the results of VAR with and without trend are identical. All the estimated coefficients, but for the lagged interest rate variables, are significant. The coefficients of lagged error terms are negative, and serve as negative feedback mechanisms in both equations. This implies that if there are departures from equilibrium in the previous period, this departure is reduced by about 37% in the current period. The χ^2 statistics indicate that there is no serial correlation, functional form misspecification, non-normality and heteroscedasticity in the residuals. We also estimated these two equations without the lagged interest rates. However, these equations are found to be unstable in the TIMVAR tests and therefore the lagged interest variables are retained. Furthermore, we tested the residuals of both the equations for the presence of $ARCH$ effects and instability in the residuals. The computed $\chi^2(4)$ test statistics of 4.866 and 4.880 are less than the 5% significance value of 9.488. Thus, our ECM models are satisfactory for forecasting and policy. We have subjected equation (5) and (6) to TIMVAR stability tests and found that both the tests based on SRR and $SSRR$ indicated stability. The plots of $SSRR$ for both equations are given in Figure 1 and Figure 2 below.

Thus the two ECM formulations can be said to have captured the dynamics underlying the money demand function better than the PAM equation (2). Similar results on the relative merits of the ECM modeling over the PAM formulations were obtained for the USA and the UK; see Hendry and Ericsson (1991). It is difficult to say which of these two ECM equations is better, since their parametric estimates, summary statistics and stability are almost identical. However, since the removal of the trend variable from the VAR yielded significant estimates of the long run equilibrium parameters, we prefer equation (4) and the corresponding ECM in equation (6). Therefore, our earlier conclusion that the long run income elasticity of demand for money is about 1.2%, interest rate elasticity is -0.18 and both are significant can be said to be robust.

Economic reforms in India started from 1991 and to examine their impact we introduced a reforms dummy (unity from 1991 to 2003 and zero otherwise) and reestimated equations (5) and (6). A similar procedure was used for the Philippines by Haper and Kutan (2003) to measure the effects of financial reforms. 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 is negative, almost zero (-0.0015) and insignificant. While reforms do not seem to have any significant effects on narrow money in India, FIGURE-1

CUSUMSQ Test for Equation 5

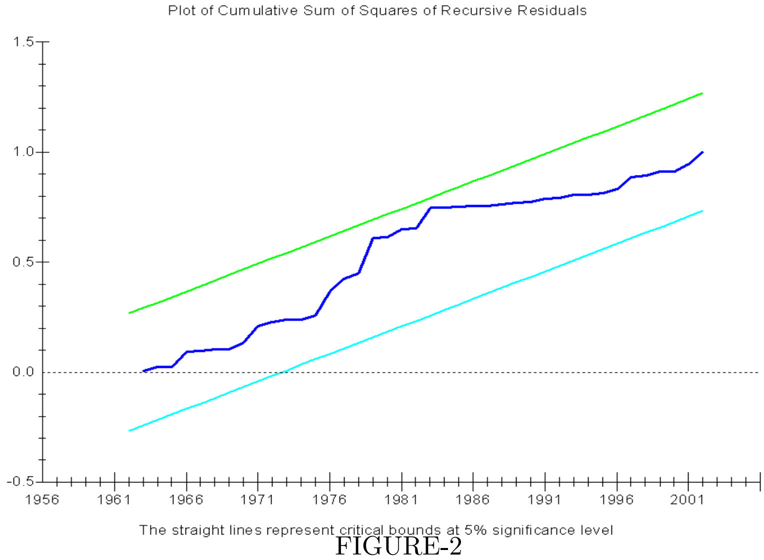
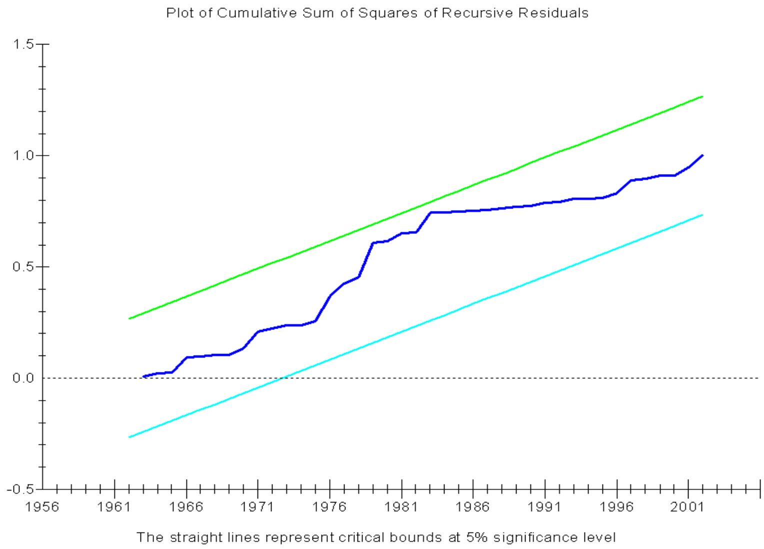


FIGURE-2

CUSUMSQ Test for Equation 6



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 demand or 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. However, it is also doubtful if the effects of financial reforms can be adequately captured with only a reforms dummy variable because, it does not capture the effects of increased competition that may reduce the margins between the interest rates on deposits and the lending rates.

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 PAM based specifications in the levels of variables are unsatisfactory and the demand for money should be modeled within the VAR framework. Our estimates with VAR methodology imply that there is a well determined and stable demand for money for half a century from 1953 to 2003. 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 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 -0.18. It should be noted that income and interest rate elasticities implied by the PAM equation are also close to the VAR model estimates. The latter model is better in the sense that it captures the dynamic adjustment process far better than the PAM equation and yielded a temporally stable demand for money function. Therefore, estimates based on the VAR model are appropriate for policy formulation and money supply targeting.

Our finding that the demand for money 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 a need for shifting from targeting narrow money to interest rate targeting in India, unless financial reforms are found to destabilize the demand for money in the future.

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 the rate of inflation as well as the implications of targeting money supply for seignorage revenue to finance the budget deficits. Furthermore, it would be valuable to investigate the nature and stability of the demand for broader money ($M3$) by extending some earlier works, e.g., Parikh (1994). These are outside the scope of our current paper.

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