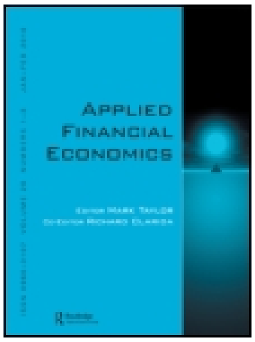

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Is the US demand for money unstable?

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The demand for money (M1) for the US is estimated with annual data from 1960 to 2008 and its stability is analysed with the extended Gregory and Hansen (1996b) test. In addition to estimating the canonical specification, alternative specifications are estimated which include a trend and additional variables to proxy the cost of holding money. Results with our extended specification showed that there has been a structural change in 1998 and the constraint that income elasticity is unity could not be rejected by subsample estimates. Short run dynamic adjustment equations are estimated with the lagged residuals from the Fully Modified Ordinary Least Squares (FMOLS) estimates of cointegrating equation and also with the General to Specific (GETS) approach.

Keywords: demand for M1; US; structural breaks; income elasticity; cost of holding money

JEL Classification: E41

I. Introduction

The US demand for money function and its stability have been analysed by many studies. Some often cited works are Goldfeld (1976), Judd and Scadding (1982), Lucas (1988), Poole (1988), Baba *et al*. (1992), McNown and Wallace (1992), Stock and Watson (1993), Hoffman *et al*. (1995) and Yossifov (1998), Ball (2001) and Choi and Jung (2009).[1] Duca and VanHoose (2004) have surveyed important developments in monetary economics including the need to study the demand for money. They also noted that the current view of this relationship is unimportant because many central banks have abandoned targeting monetary aggregates and switched to the rate of interest as the monetary policy instrument. However, according to Duca and VanHoose, studying this relationship is not an irrelevant activity and therefore summarized the salient features of some key empirical works. Others who take a similar view are Leeper and Roush (2003) and Ireland (2004). Ireland has estimated a business cycle model for the USA within the ISLM model framework, augmented with a Phillips curve, with the post-1980 quarterly data. However, he found that money played relatively a smaller role in explaining the dynamics of inflation

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[1] Studies by Lucas (1988), Poole (1988), Stock and Watson (1993) and Hoffman *et al*. (1995) asserts that the demand for M1 in US is stable over the twentieth century. However, Goldfeld (1976) and Judd and Scadding (1982) found that the money demand is unstable during 1970s. Further, McNown and Wallace (1992) and Yossifov (1998) obtained implausible income elasticity of M1 and M2. See Sriram (1999) for a survey.

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and output. In our view this does not mean that the study of the demand for money is redundant because Ireland's results are also consistent with instability in the demand for money, which might have contributed to the poor correlation between money, inflation and output.[2] The dependent variable in the demand for money varied from the narrow definition of money (M1) to broader measure with weighted averages of monetary aggregates. Although some influential studies have found that the US demand for money (mainly M1) is stable for a long period up to the early or mid 1970s, others have found that it has become unstable since then due to financial reforms, improvements in payments technology and cash management practices, which have been significantly improved with parallel progress in computer technology. In this article, we shall examine if these efficiency effects can be captured with modified specifications to improve the stability of this relationship and if the income elasticity of the demand for money is unity as found in many earlier studies. For this purpose, we shall use annual data from 1960 to 2008 for M1.

This article is organized as follows. Section II reviews a few recent contributions on the US demand for M1. Section III discusses our modifications to its canonical specification and presents estimates of cointegrating equations. Structural break tests are also conducted in this section. Since these tests have weak power against the null of no cointegration, it is necessary to use discretion to specify and estimate stable demand for money functions. Section IV concludes.

II. Recent Studies of US Demand for Money

Several pre-1980s studies have generally used annual data and the following canonical specification for the demand for real money.

$$\ln m_t = \theta_0 + \theta_y \ln(y_t) + \theta_R R_t + \varepsilon_t \qquad (1)$$

where θ_0 = intercept, m = real money stock, y = real output, R = cost of holding money proxied with the nominal short-term interest rate and $\varepsilon \sim N(0, \sigma)$. They found that this function is stable and estimates of the income elasticity (θ_y) are about unity and the semi-interest rate elasticity (θ_R) is around -0.1. According to Friedman and Kuttner (1992), the above canonical specification for M1 is cointegrated with income and the rate of interest for the period 1960 to 1979, but became unstable if samples are extended to include data from the 1980s. However, Ball (2001), in an insightful study, noted that stability tests did not show breaks in the demand for M1 with data up to 1987, but a break is generally found if the samples include data through 1996; Duca and VanHoose (2004, p. 259). He also found that when the data are extended beyond 1987, the pre-1970s estimates of income and interest rate elasticities reduce by half so that $\theta_y = 0.5$ and $\theta_R = -0.05$.

More formal break tests are conducted on the US demand for money by Gregory and Hansen (1996a) with annual (1901 to 1985) and quarterly (1960Q1 to 1994Q4) data. They have used the canonical specification, mainly to illustrate their method for testing for a *single* endogenous break in a cointegrating equation, and not to examine the adequacy of alternative specifications of the demand for money. More recently Choi and Jung (2009) have applied the Bai and Perron (2003) tests for testing for *multiple* endogenous breaks in the US demand for money. Unlike the Gregory and Hansen one step tests for cointegration with one break, the Bai and Perron tests are tests for multiple breaks in equations estimated with Ordinary Least Squares (OLS). Therefore, it is necessary to test for cointegration for each subsample implied by these first stage break tests. Choi and Jung have used a pragmatic option to test for multiple breaks because there is no formal test so far to test for cointegration with multiple endogenous breaks in a single step. Furthermore, the main objective of Choi and Jung seems to illustrate the use of the Bai and Perron tests for testing for multiple endogenous breaks with the canonical specification of the demand for money and not to examine the merits of alternative specifications of this relationship.

Gregory and Hansen found an intercept break in 1941 in their annual data, but in their quarterly data there is only weak evidence for both an intercept and slope shift in 1975Q2. In contrast, Choi and Jung, using quarterly data (1960Q1 to 2000Q2) have found that there are two breaks in the US demand for money in 1974Q2 and 1986Q1. Therefore, they tested for cointegration with the Johansen Maximum Likelihood (JML) method and found one cointegrating equation for each of the three subsamples implied by their break tests. However, instead of reporting JML estimates of the cointegrating equations, which should have been straightforward, they reported

[2] The pros and cons on whether monetary aggregates or the interest rate should be used as monetary policy instrument are based on whether the effects of monetary policy are through the real balance effect or through interest rate effect. Objections to the use of the rate of interest as policy instrument may also be raised because changes to the rate of interest would have important distributional effects and these are ignored in debates on the merits of using it as a policy instrument.

estimates with the four alternative methods.[3] Since we shall use later the Phillips and Hansen Fully Modified OLS (FMOLS) method to estimate the cointegrating equations, besides JML, we shall briefly summarize Choi and Jung's findings based on FMOLS. They found that in the first subsample (1959Q1 to 1974Q1) income elasticity was 0.33 and increased to 0.49 in the second subsample (1974Q2 to 1986Q1) and then declined to 0.25 in the third subsample (1986Q2 to 2000Q2).[4] Estimates of semi-interest rate elasticity was insignificant in the first subsample but increased in absolute value to −0.045 and became significant in the third subsample. They did not report estimates of the intercepts.

These aforesaid works provided valuable insights and in particular Choi and Jung are the first to test formally for structural breaks in the demand for money function of the US.[5] However, a weakness in the previous studies is that there is no trend in their specifications. In fact many other earlier studies have also ignored trend including a path breaking work by Baba *et al*. (1992). Ideally, the following specification with trend could be estimated to capture some effects of improvements in the transactions technology[6]:

$$\ln m_t = \theta_0 + \theta_T \text{Trend} + \theta_y \ln y_t + \theta_R R_t + \varepsilon_t \qquad (2)$$

There is a problem with including trend. As Ball has pointed, including trend may give unreliable estimates of the parameters because of the high colinearity between trend and income. Therefore, he has also estimated cointegrating equations with the constraint that income elasticity is unity and these estimates turned out to be informative and did not change his major finding that the demand for money has become unstable when the sample size is extended beyond the late 1980s and up to the mid-1990s.

The specification in (2) may still be inadequate because it can be extended to include other variables that could better proxy the cost of holding money. For example, in addition to the short-term interest rate, the inflation rate and exchange rate can be used by money holders to proxy the cost of holding money because inflation reduces the real value of money. As Choi and Jung have noted, it is likely that cash management practices might have significantly changed after the introduction of the flexible exchange rates in 1973. Ignoring these variables in the demand for money may also cause instability in this relationship. However, the usefulness of such extensions depends on the empirical results, but inclusion of these two additional variables is justified when the Fisher condition holds only weakly, i.e. the correlation between the rate of interest and inflation is not high. In our sample this correlation is about 0.6 and the correlation between the rate of interest and Real Effective Exchange Rate (REER) is very low.[7] Subject to this caveat our extended specification of the demand for money is

$$\begin{aligned}\ln m_t = \theta_0 + \theta_T \text{Trend} + \theta_y \ln(y_t) + \theta_R R_t \\ + \theta_p \Delta \ln P_t + \theta_x \ln(\text{REER}_t) + \varepsilon_t \qquad (3)\end{aligned}$$

where $\Delta \ln(P) =$ rate of inflation. This improved specification may improve the stability of the demand for money and or show that the structural changes are minor.[8] We shall estimate (3) in the following section. Definitions of the variables and data sources are in the Appendix.

III. Empirical Results

In this section, we shall estimate the cointegrating equations for the demand for M1 with alternative specifications and sample periods with annual data from 1960 to 2008. The Phillips and Hansen FMOLS method is used because it is simpler and quick to implement and convenient to estimate a number of cointegrating equations with different specifications and sample periods. However, there is no formal cointegration test for FMOLS estimates and the significance of the coefficients is generally used for

[3] One presumes that their JML estimates were unsatisfactory. The four alternative methods are Stock and Watson's (1993) Static OLS (SOLS) and Dynamic OLS (DOLS), Phillips's (1991) Band Spectral (PBSR) estimator and Phillips and Hansen's (1990) Fully Modified (PHFM) estimator.

[4] Unfortunately, Choi and Jung's dating the break dates gives the impression that they are using monthly and not quarterly data. Therefore, our notation of the dates for the subsamples based on what they might have intended to.

[5] Gregory and Hansen used the demand for money to illustrate their techniques and *per se* they are not interested in a wider sense in the issues on the demand for money. Rao and Kumar (2007, 2009) have used their tests for testing structural breaks in the demand for money of Fiji and Bangladesh.

[6] A non linear trend or a linear trend with dummy variables can also be included.

[7] See Baba *et al*. (1992, p. 29) for a similar reasoning for the inclusion of the rate of inflation. In our sample the correlation between interest rate and inflation rate is about 0.7 implying that 49% of interest changes are explained by inflation.

[8] Baba *et al*. (1992) analysed the US demand for money from this perspective and although they did not conduct formally tests for structural breaks, they showed that their improved specification adequately explains the missing money episode (1974 to 1976), great velocity decline (1982 to 1983) and M1 Explosion (1985 to 1986); see Baba *et al*. (1992, fn. 1) for the details of these episodes.

Table 1. Estimates of cointegrating equations

Method	Period	Equation	θ_0	θ_T	θ_y	θ_R	θ_p	θ_x
FMOLS	1962 to 2008	1	1.412		0.185	−0.009		
			(9.89)*		(11.60)*	(3.51)*		
FMOLS	1962 to 2008	2	−1.033	−0.010	0.496	−0.011		
			(0.53)	(1.28)	(2.00)*	(4.24)*		
FMOLS	1962 to 2008	3	−3.900	−0.035	1.018	−0.013	−1.177	−0.431
			(3.05)*	(6.31)*	(6.19)*	(5.52)*	(4.29)*	(5.57)*
JML	1962 to 2008	3	A	−0.033	0.928	−0.014	−2.282	−0.347
				(2.88)*	(2.83)*	(3.42)*	(4.02)*	(3.31)*
JML	1962 to 2008	1	4.457		−0.071	−0.058		
			(0.97)		(0.18)	(0.78)		

Notes: 'A' denotes values estimated with unrestricted intercepts and restricted trends in the VAR. Absolute *t*-ratios are in the parentheses.
* Denotes significance at 5% level.

the validity of the estimated cointegrating equations. Therefore, we shall use, on a selected basis, the JML test for cointegration, as in Choi and Jung, and report estimates of the JML cointegrating equations if they are plausible. Finally, we shall estimate the short-run dynamic adjustment equations with the lagged residuals from the FMOLS equations and also with the General to Specific (GETS) approach of the London School of Economics, of which David Hendry is the most ardent exponent.

In Table 1, the estimated cointegrating equations with FMOLS for specifications in Equations 1–3 are reported in the first three rows. All the estimated coefficients are significant at the 5% level but estimates of income elasticities change significantly with the specifications. While the estimate of income elasticity (θ_y) is low at about 0.2 in the canonical specification, it has increased to about 0.5 when trend is added. Only in our extended specification in Equation 3 $\theta_y = 1.0178$ and equal to its stylized value of unity.

The cointegrating equations with JML are estimated for Equations 3 and 1 and reported, respectively, in rows 4 and 5 of Table 1 to highlight the improvements with our extended specification. Both the trace and maximal eigenvalue tests could not reject the null of a single cointegrating equation.[9] The estimated income and interest rate elasticities in the extended Equation 3 are close in both the FMOLS and JML estimates with some differences in the coefficients of inflation and real effective exchange rate. JML estimates of these two parameters are more in absolute magnitudes than in FMOLS. Note that, compared to this closeness, there are significant differences in the FMOLS and JML estimates of the canonical specification in (1).

With the canonical specification estimates of income and interest rate elasticities are lower and significant in FMOLS, but neither is significant in the JML estimates. Since our extended specification seems more robust, we select this to test for structural breaks. The estimate of the trend coefficient of this equation implies that demand for money has been declining at the rate of about 3.5% per year due to financial reforms and improvements in payments technology. However, as noted by Ball, the estimate of the coefficients of trend and income are unlikely to be accurate due to the high colinearity between these two variables. We shall discuss this problem later.

		Equation 1		Equation 3	
$H0$	$H1$	Test statistic	95% CV	Test statistic	95% CV
$r = 0$	$r = 1$	49.377	37.860	81.231	22.040
$r \leq 1$	$r = 2$	30.884	31.790	5.963	15.870
$r \leq 2$	$r = 3$	13.189	25.420	3.754	9.160

We shall use the Gregory and Hansen (1996b) extended break test to test for cointegration with a single structural break with the trend variable in the specification. When this test is implemented we found that there is no cointegration with a significant break in the extended Equation 3. The test results are in row 1 of Table 2. This result may partly be due to colinearity between trend and income. To avoid this problem we followed a suggestion by Ball and assumed that $\theta_y = 1$ in (3). This is a reasonable assumption because in both the FMOLS and JML estimates, the constraint that $\theta_y = 1$ could not be

[9] To conserve space we report below only the Maximal Eigenvalue test for Equations 1 and 3.

Table 2. Tests for cointegration with a structural break

	Specification	Test-statistic	5% CV	10% CV	Break date
1	$\ln m_t = \text{Intercept} + \theta_T\text{Trend} + \theta_y \ln(y_t) + \theta_R R_t + \theta_p \Delta \ln P_t + \theta_x \ln(\text{REER}_t)$	−6.20	−6.84	−6.58	1990
2	$\ln m_t - \ln y_t = \text{Intercept} + \theta_T\text{Trend} + \theta_R R_t + \theta_p \Delta \ln P_t + \theta_x \ln(\text{REER}_t)$	−5.64	−6.32	−6.16	1998
3	$\ln m_t - \ln y_t = \text{Intercept} + \theta_T\text{Trend} + \theta_{pc} PC_t$	−5.22	−5.50	−5.24	1998

Table 3. Estimates of cointegrating equations for subsamples

Method	Period	Equation	θ_0	θ_T	θ_y	θ_{PC}
FMOLS	1962 to 1997	4 unconstrained	−2.579 (1.83)**	−0.015 (2.72)*	0.684 (3.84)*	−0.035 (8.41)*
FMOLS	1998 to 2007[a]	4 unconstrained	−2.621 (0.83)	−0.014 (1.44)	0.676 (1.74)**	−0.043 (5.36)*
FMOLS	1962 to 1997	4 constrained	−5.074 (381.63)*	−0.025 (44.99)*	1.00	−0.036 (7.84)*
FMOLS	1998 to 2008	4 constrained	−5.293 (146.72)*	−0.022 (26.49)*	1.00	−0.053 (8.91)*

Notes: [a]Inclusion of data for 2008 caused convergence problems and we estimated with data to 2007. θ_{pc} = Coefficient of the principal component of R, DLP and LREER. Absolute t-ratios are in parentheses.
* and ** denote significance at 5 and 10% levels, respectively.

rejected at the 5% level by the Wald test. Test results, in row 2 of Table 2, with this modification also failed to detect a significant break. However, the identified break date is 1998 but it should be noted that this is not significant even at the 10% level. To increase the degrees of freedom and efficiency of the test, we have proxied the cost of holding money with the Principal Component (PC) of R, ΔDLP and LREER and the specification with this modification is as follows:

$$\ln m_t - \ln y_t = \text{Intercept} + \theta_T\text{Trend} + \theta_{pc}\text{PC}_t \qquad (4)$$

where the dependent variable can also be interpreted as the inverse of velocity. When Equation 4 is tested for a break, the test statistic (absolute value) is only marginally less than the CV (absolute value) at the 10% level, also indicating a break in 1998 and the results are in row 3 of Table 2.

These break test results should be taken with some caution for a few reasons. First, the Gregory and Hansen tests are joint tests for cointegration with a structural break and there may actually be cointegration without a structural break. Second, there may be more than one break and both the intercept and slope coefficients may change. Third, they have low power against the null of no cointegration and discretion is necessary to interpret them.

With a somewhat weak but not a totally unsatisfactory test result for a structural break, we proceeded further as follows. We have estimated the cointegrating equations for the subsamples (implied by a break in 1998) for the specification in Equation 4 in both an unconstrained and constrained form on θ_y. The dependent variable in the former is $\ln m$ and in the latter $(\ln m - \ln y)$ and these formulations help also to test if the income elasticity is unity in both subsamples viz., 1962 to 1997 and 1998 to 2008. Both FMOLS and JML methods are used for estimating the cointegrating equations but JML did not yield any meaningful results for the second subsample perhaps because there are only 11 observations.[10] FMOLS estimates are good and given in Table 3. While estimates in the unconstrained equations the intercept and the coefficients of T and $\ln y$ are almost equal in both subsamples, the coefficient of PC is higher in the second subsample. However, a Wald test rejected the null that all the coefficients are equal

[10] Perhaps this may be the reason why Choi and Jung did not report JML estimates of the cointegrating equations for their subsamples even though JML procedure has been used to test for cointegration. In our JML estimates for the subsample all the estimated coefficients are insignificant but are correctly signed. In the second subsample the coefficients of income and trend are wrongly signed and all are insignificant.

in both subsamples. The computed test-statistic, with p-value in square brackets, is $\chi^2_{(4)}$ =420.845[0.000]. Although point estimates of income elasticity are about 0.7, a Wald test could not reject the null that income elasticity in both subsamples is one at the 5% level. The computed test statistics for the first and second subsamples are $\chi^2_{(1)}$ =3.143[0.076] and $\chi^2_{(1)}$ =0.697[0.404]. Next, we have re-estimated with FMOLS the cointegrating equations for both subsamples with the constraint that income elasticity is unity and these are in the third and fourth rows of Table 3.

Estimates of trend have increased in both subsamples compared to the unconstrained estimates in rows 1 and 2. Although the coefficients of PC have remained the same in the first subsample in both the unconstrained and constrained estimates, it has increased in absolute value in the constrained estimates of the second subsample. A Wald test that all the coefficients in both subsample equations in rows 3 and 4 are equal has rejected the null. We then tested that each individual coefficient is the same in both subsamples in rows 3 and 4. The computed χ^2 test-statistics for the null that intercept and the coefficients of trend and PC, with p-values in the square brackets, are respectively: 270.271 [0.000], 31.309 [0.000] and 13.744 [0.000] indicating they differ significantly even at the 1% level. Therefore, we may conclude that there has been a structural change in the US demand for money in 1998 and in particular the intercept and the response to the cost of holding money have increased in absolute magnitude after 1997.

It would be interesting and informative to proceed further and estimate the short-run dynamic equations based on the cointegrating equations with a structural change. We shall use two procedures for this purpose. First, we shall use the lagged error terms i.e. Error Correction Models (ECMs) implied by the cointegrating equations, based on FMOLS, for the two subsamples in Table 3. Next we shall use the GETS approach. In contrast to various cointegration methods, where the dynamic short-run equation is estimated in two steps, the dynamic equation with the cointegrating equation can be estimated in one step with GETS. Recent studies seem to have neglected the short-run dynamic equations of the demand for money and are satisfied with estimating the cointegrating equations and it is not known if estimating long-run relationships without short-run dynamics would give robust and reliable estimates of the cointegrating equations.[11] In this respect GETS approach has an advantage over conventional cointegration methods. We have used PcGets to select the optimal lag structure for the dynamic equations in both methods. The search procedures in PcGets minimize the path dependent biases; see Hendry and Krolzig (2001) and Rao and Singh (2006). PcGets has selected, for the whole sample period, a parsimonious lag structure for the short-run equation with the 2 lagged ECM terms, implied by the FMOLS estimates for the constrained specification in Table 3. The estimate of the parsimonious Equation 5 is as follows:

$$\begin{aligned}\Delta(\ln m_t - \ln y_t) = & -\underset{(4.64)*}{0.624}\text{ECM97}_{t-1} - \underset{(2.87)*}{1.600}\text{ECM08}_{t-1} \\ & -\underset{(6.01)*}{0.031}\Delta\text{PC}_t \quad \underset{(5.59)*}{0.030}\Delta PC_{t-1} \\ & +\underset{(9.41)*}{0.797}\Delta(\ln m_{t-1} - \ln y_{t-1}) \\ & \overline{R}^2 = 0.494; \text{SER} = 0.019 \\ & \chi^2_{sc} = 1.664(0.197); \chi^2_{ff} = 5.328(0.021); \\ & \chi^2_{nn} = 1.785(0.410); \chi^2_{hs} = 0.835(0.361)\end{aligned} \tag{5}$$

All the coefficients in Equation 5 are significant at the 5% level and the adjustment coefficients are correctly signed. The $\overline{R}^2$ at about 0.5 is satisfactory and the χ^2 tests on the residuals indicate no serial correlation and non normality in the residuals. However, the functional form misspecification test is only insignificant at the 1% level but becomes significant at 2.1%. It may be noted that search for a dynamic structure is an empirical issue and it is hard to discover the correct dynamic structure. The estimate of the adjustment coefficient for the second subsample period at −1.6 is more than twice for the first subsample of −0.6, implying that the speed of adjustment towards equilibrium has substantially increased after 1997. Since the absolute value of the adjustment coefficient for the second subsample exceeds unity, there would be fluctuations in the adjustment path around the equilibrium value.[12] The coefficient of ΔPC_t is negative and its one

[11] Perhaps the exception is Baba *et al.* (1992) who estimated with quarterly data (1960Q3 to 1988Q3) demand for money in the US. Baba *et al.* (1992, p. 26) observe that 'We infer that the reason for the shifts in alternative models is their omission of appropriate dynamic structure and of important variables'.

[12] Generally it is mistaken that if the adjustment coefficient exceeds unity there is no convergence. However this is valid only if the absolute value of the coefficient exceeds 2 but there would be fluctuations in the adjustment path if the estimate is unity or more than unity but below 2.

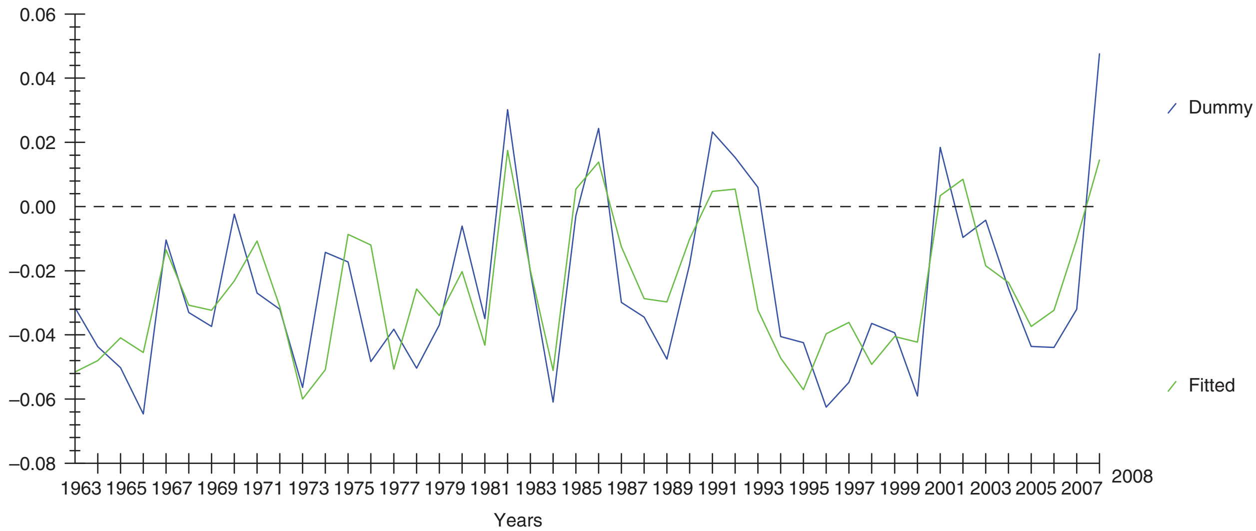

Fig. 1. Plot of actual and fitted values

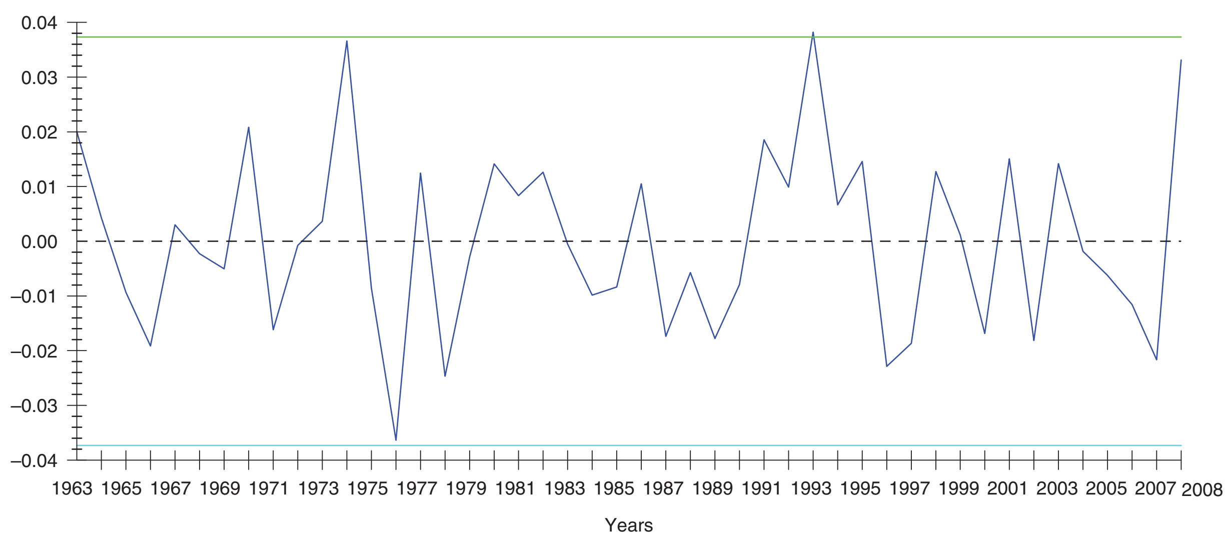

Fig. 2. Plot of residuals and two SE bands

period lagged value is positive implying that there would be an immediate decrease in the demand for money when the cost of holding money increases but then this is offset by an increase in the next period offsetting the previous effect. The coefficient of the lagged dependent variable is large at 0.8 implying that there is considerable persistence in the changes of the demand for money. Although the plots of predicted and actual values in Fig. 1 and residuals in Fig. 2 seem satisfactory, there are some large positive and negative errors that exceed 2% in absolute magnitude. There are 17 such errors and 12 are in the first subsample.[13]

GETS estimate of the demand for money with a dummy variable *D*98 (zero up to 1997 and one afterwards) to allow for a structural change in the cointegrating equation and with the constraint that income elasticity is unity are in given in the

[13] These errors are in 1963, 1966, 1971, 1974 to 1978, 1987, 1989, 1993, 1996, 1998 to 1999, 2002, 2004 and 2008.

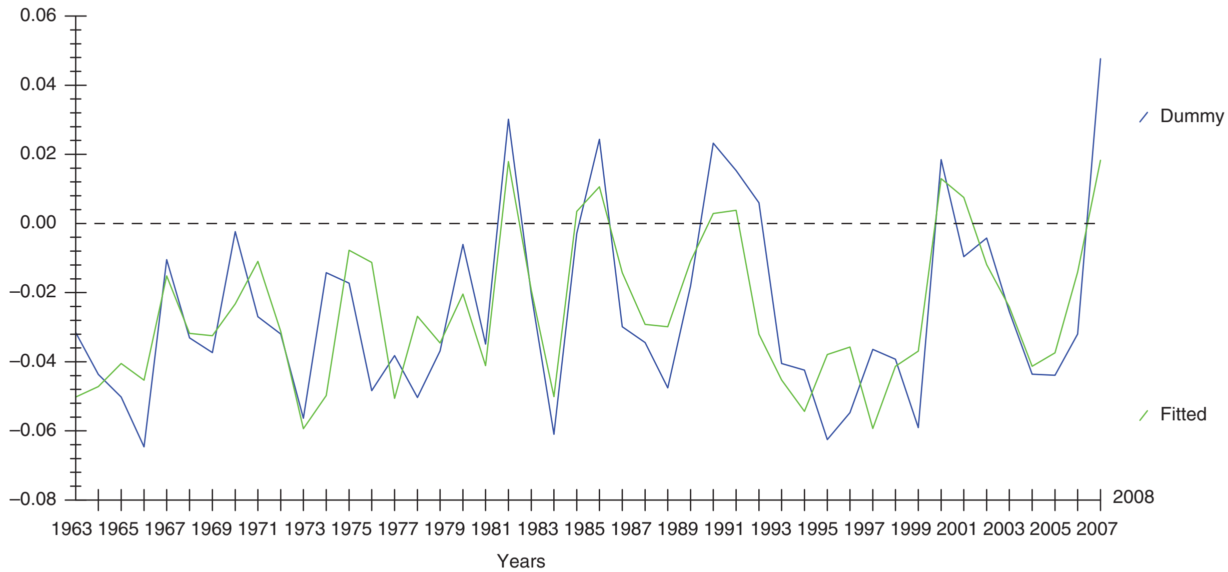

Fig. 3. Plot of actual and fitted values: GETS equation

following equation:

$$
\begin{aligned}
&(\Delta \ln m_t - \Delta \ln y_t) \\
&\quad = -\underset{(3.52)*}{0.516}(1 - D98)((\ln m_{t-1} - \ln y_{t-1}) \\
&\qquad -(-\underset{(215.36)*}{5.084} - \underset{(38.56)*}{0.025}T - \underset{(6.24)*}{0.036}\mathrm{PC}_{t-1})) \\
&\qquad -\underset{(2.36)*}{1.003}D98((\ln m_{t-1} - \ln y_{t-1}) \\
&\qquad -(-\underset{(59.28)*}{5.245} - \underset{(11.56)*}{0.022}T - \underset{(2.46)*}{0.038}\mathrm{PC}_{t-1})) \\
&\qquad -\underset{(5.28)*}{0.031}\Delta\mathrm{PC}_t + \underset{(3.03)*}{0.020}\Delta\mathrm{PC}_{t-1} + \underset{(2.56)*}{0.408}\Delta(\ln m_{t-1} - \ln y_{t-1}) \\
&\overline{R}^2 = 0.519;\ \mathrm{SER} = 0.019 \\
&\chi^2_{sc} = 0.958(0.328);\ \ \chi^2_{ff} = 11.245(0.001); \\
&\chi^2_{nm} = 0.464(0.793);\ \ \chi^2_{hs} = 0.003(0.982) \qquad (6)
\end{aligned}
$$

All the estimated coefficients are significant in Equation 6 at the 5% level and the χ^2 tests on the residuals, except the functional form test as for Equation 5, are insignificant. The $\overline{R}^2$ at about 0.52 is satisfactory and a trifle more than for Equation 5. A Wald joint test with the null that all the coefficients in both subperiods are equal is rejected at the 5% level. However, Wald tests on estimates of individual coefficients of the intercepts, adjustment coefficients and the slopes are equal produced mixed results. While the test did not rejected the null that the intercepts are equal at the 5% level, this null is rejected at the 10% level. The null that the adjustment and slope coefficients are equal is not rejected at the 5% level. Since the joint test that all the combined coefficients are equal has rejected the null, we may conclude that there has been a structural change in the US demand for money in 1998 and ignore the tests on the equality of individual coefficients. However, unlike in Equation 5, in Equation 6 it is possible to test which coefficients in the cointegrating equation have also changed because in GETS the parameters in the cointegrating equation and the short-run dynamics are estimated in one step. Based on the point estimates of the individual coefficients, there has been significant improvement in the speed of adjustment to equilibrium after 1997. While the intercept shifted down and there is a marginal improvement in the long-run response to changes in the cost of holding money, the change in the coefficient of trend is very small. In the short-run dynamics part of the equation (6), the effect of changes in the cost of holding money is similar to (5), but the change in its lagged value has a smaller effect. The coefficient of the lagged dependent variable is smaller indicating decreased persistence. The plots of the actual and predicted values are in Fig. 3 and the errors in Fig. 4. In contrast to Equation 5 the number of large errors, exceeding 2%, in (6) are less. There are only 10 such errors and seven are in the first subsample. The years in which the errors exceed 2% are 1970, 1974, 1976, 1978, 1991, 1993, 1996, 1998, 2000 and 2008. While our GETS estimates could not explain the missing money episode of 1974 to 1976, the error in 1975 is less than 1%. However, errors in

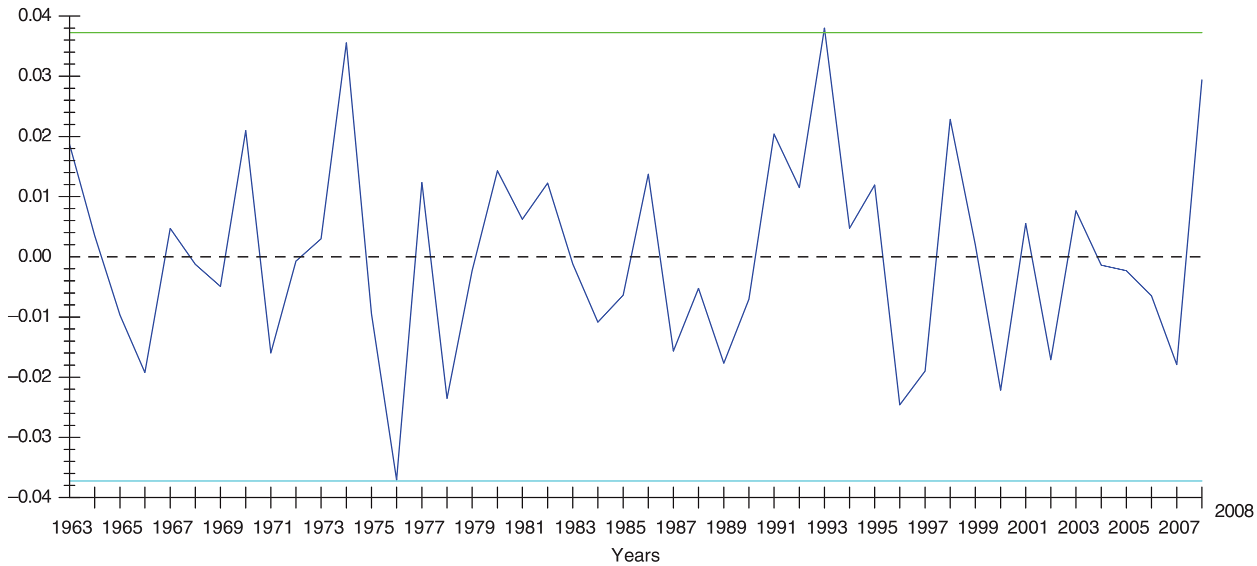

Fig. 4. Plot of residuals: GETS equation and two SE bands

1974 and 1976 are about 3.5%. Our equation has adequately explained the great velocity decline of 1982 to 1983 and the explosion in M1 during 1985 to 1986. Errors in these two episodes are less than 1% but the error in 1986 is 1.4% (see Baba *et al.* (1992) for an explanation of these episodes). Errors in the 1990s and 2000s are marginally higher than 2%. On this basis we may say that GETS estimates are as good and perhaps better than the standard approaches based on the 2-step methods of estimating the short-run dynamic adjustment equations.

Thus our extended specification, estimates of the short-run dynamic equations and the constraint that income elasticity is unity, by and large, seem to have reduced major instabilities found in several studies which have estimated only the cointegrating equations of the canonical specification. Based on the estimates of Equations 5 and 6 we may conclude that the structure of the US demand for money has changed, perhaps marginally with small changes in the intercepts and other coefficients, after 1997 mainly due to improvements in the speed of adjustment of the money market towards its equilibrium because of financial liberalization.

IV. Conclusions

This article has estimated alternative specifications of the demand for money of the US from 1960 to 2008 and examined its stability using a formal test for a single structural break. We found that inclusion of trend and additional variables, besides the rate of interest to capture the effects of cost of holding money, are useful and improved the stability of this relationship. We have tested for cointegration and presented the estimates of the cointegrating equations with FMOLS and JML and also estimated the short-run dynamic equations. Structural break tests indicated that there is no strong evidence that our extended specification is unstable. However, there is some weak evidence for a break in 1998 and this break date is different from those reported by Ball and found by Choi and Jung. When the subsample estimates are made with the constraint that income elasticity is unity, to overcome multicolinearity between income and trend, a joint Wald test showed that FMOLS estimates for the subsample periods differ significantly but point estimates showed only minor changes in the parameters. On the basis of our tests we concluded that the demand for M1 in the US has been, by and large, stable but for a small change after 1997. Financial reforms seem to have reduced the demand for M1 on average by about 2% to 2.5% annually and the response to the cost of holding liquidity has remained the same at about −0.36 in both subsamples. Finally, we estimated the short-run dynamic equations with two alternative methods and both yielded similar results. Estimates with GETS are more satisfactory because the number of large errors are relatively few. In the subsample of 1998 to 2008 there are only three errors that exceeded 2% and one is towards the end of the sample in 2008. Furthermore, GETS estimate could explain the errors as the decline in the velocity and the great explosion of M1 but not the missing money episode of the mid-1970s.

Nevertheless, this article has some limitations. We have assumed that income elasticity is unity to avoid multicolinearity. Alternative assumptions about this parameter are possible to search for improved estimates. The much coveted and superior JML method did not yield meaningful estimates of the parameters of the cointegrating equations for the subperiods although both FMOLS and JML gave virtually identical cointegrating equations for the whole sample period. We hope that our methodology and results will interest other investigators to analyse the stability of money demand function in the US with alternative data sets and also in other countries. This is timely at a time when quantitative targets have attracted many central banks to stimulate the economy from the current unprecedented worldwide depression.

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Appendix: Data

m = real currency in circulation plus demand deposits (seasonally adjusted). Data are from (IFS-2008).

y = real Gross Domestic Product (GDP) at factor cost. Data are from (IFS-2008).

R = Short term treasury bill rate (6 months). Data are from (IFS-2008).

P = GDP Deflator (2000 = 100). Data are from (IFS-2008).

REER = real effective exchange rate based on normalized unit labour costs. Data are from (IFS-2008).