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Source: *The Journal of Finance*, Vol. 39, No. 5 (Dec., 1984), pp. 1383-1396
Published by: Blackwell Publishing for the American Finance Association
Stable URL: http://www.jstor.org/stable/2327733
Accessed: 27/11/2009 07:50

THE JOURNAL OF FINANCE • VOL. XXXIX, NO. 5 • DECEMBER 1984

Stability of the U.S. Short-Run Money Demand Function, 1959–81

KUAN-PIN LIN and JOHN S. OH*

ABSTRACT

Stability tests are performed for the conventional U.S. money demand equation using switch regression techniques. This methodology provides for the identification of the shift point and the type of shift (abrupt or drift), and is conducive to hypothesis testing to determine the sources of the shift for the regression equation. Our findings do not support the contention that the 1974 change in money demand equation is a downward shift in the constant term, as suggested by many recent empirical money demand studies.

EMPIRICAL METHODOLOGIES EMPLOYED to investigate the stability of the conventional U.S. money demand function are normally selected from one of the following: the Brown-Durbin-Evans (BDE) recursive residuals approach that utilizes the "cumsum" or "cumsum of squares" test statistics (Khan [19], Hafer and Hein [11]); the varying parameter regression (VPR) which is a special case of the random coefficient method introduced by Cooley and Prescott (Laumas and Mehra [20], Garbade [6]); the Chow test; and the dummy variable approach when an a priori shift point is known. These methods are limited in that they either estimate different sets of parameters and test for the existence of more than one state—known a priori—for the parameter vector (e.g., Chow test, and dummy variables method), or provide single estimates for the entire data set and test the hypothesis that the estimated parameters of unknown states are stable (e.g., BDE and VPR).

It should be noted that in the cases of the Chow test and the dummy variable method, prior information of the shift point is necessary and that this shift is of the abrupt type. If the shift is gradual through time (i.e., drift), separate regression estimates based on an assumed a priori shift point will result in unreliable parameter estimates for the separate regimes. On the other hand, for the methods that provide single estimates of the entire time series to test for parameter stability (e.g., BDE and VPR), no estimate of the shift point can be made explicit.[1] Additionally, it should also be noted that both the BDE and VPR

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[1] For an attempt at locating the exact shift point through BDE for the Canadian money demand function, see Cameron [2]. However, the BDE cannot adequately pinpoint the time of shift (Dufour [5, p. 59]), and the power of this test is quite weak (Garbade [6, p. 54]).

methods encounter severe difficulties when the error structure of the regression changes over time. Finally, the method of post-sample forecast does not explicitly test the null hypothesis of no structural shift nor does it provide separate parameter estimates for the separate regimes.

The purpose of this paper is twofold. First, it is to report the relative success of applying the method of switch regression to the short-run U.S. money demand function (see Goldfeld and Quandt [10]). The reported results and the accompanying inferences would add to the inventory of similar studies that employ different methodologies.[2] The particular version of switch regression used in this investigation is the "Deterministic Switching Based on Time" that utilizes the "D-technique" proposed by Goldfeld and Quandt [10]. It is a departure from existing empirical money demand studies in that it is possible to make parameter estimates for separate regimes which may have different error structures and to identify the degree of contamination through examinations of serial correlations and regression variances in the separate regimes. Furthermore, this method of estimation provides for the identification of the shift point and the type of shift (abrupt or drift), and is conducive to hypothesis testing to determine the source(s) of the shift. The second purpose is to indicate that the statistical inferences made in previous studies should be interpreted with much caution.

The paper is arranged as follows. The first section describes the deterministic switch regression model and the D-technique applied to various specifications of the U.S. short-run money demand function. Section II presents the empirical results and interprets the evidence. The third section reports the results from the investigation on the possible sources of the identified shift in the function, and the final section makes some tentative conclusions from this study.

I. Switch Regression Model of the U.S. Short-Run Demand for Money

It is somewhat surprising to find that the switch regression technique has not been applied to assess the stability of the short-run money demand function. The technique was introduced by Goldfeld and Quandt [8, 9, 10], for the study of structural stability of econometric relationships and the methodological issues have been addressed extensively in Goldfeld and Quandt [10]. The following describes the switch regression model as it would apply to the U.S. short-run demand for money.

Consider that there are two structures or regimes with an unknown shift point t_0 for the U.S. short-run money demand function over the period 1959(II)–1981(IV).[3] The function under investigation is conventional in the sense that the demand for real balances (m_t) depends on a scale variable (real GNP_t), interest rate variables reflecting the opportunity cost of holding money (RCP_t, RCPP_t,

[2] To our knowledge, this is the first study of the U.S. money demand that utilizes the switch regression model.

[3] Although a two-regime case is assumed, the method of estimation described below can be extended to cases with more than two regimes. However, the number of estimated parameters increases rapidly when the number of regimes increases which in turn reduces the degree of freedom for regression estimation in each regime. Difficulties may arise for the convergence in maximizing Equation (5) below if the structural changes occur in the beginning or at the end of the sample period.

RGB_t, and RSD_t are commercial paper, previous peaks of commercial paper, long-term government bond rates, and maximum rates payable on savings deposits by federally insured institutions), and a lagged dependent variable (m_{t-1}) to account for incomplete adjustments in the short run.[4] A first-order autoregressive structure is assumed for the error term and all variables are specified in the logarithmic form. The two-regime money demand model can be written as:

$$m_t = X_t\beta^{(i)} + \gamma^{(i)}m_{t-1} + \varepsilon_t^{(i)}$$

and (1)

$$\varepsilon_t^{(i)} = \rho^{(i)}\varepsilon_{t-1}^{(i)} + u_t^{(i)}; \qquad i = 1, 2; \quad t = 2, \cdots, T$$

where superscript (i) denotes the regime index and the subscript t is the observation index. X_t is a vector of exogenous independent variables with unity as the first element. For example, $X_t = (1, \mathrm{GNP}_t, \mathrm{RCP}_t, \mathrm{RGB}_t)$ and $\beta^{(i)} = (\beta_0^{(i)}, \beta_1^{(i)}, \beta_2^{(i)}, \beta_3^{(i)})'$. Furthermore, the structural change is assumed to be time dependent only, that is, $i = 1$ if $t \le t_0$ and $i = 2$ otherwise, where $2 \le t_0 \le T$ is the specific time of structure shift. Given that there is no prior information on the timing of regime switch t_0, a D-variable is introduced to indicate the probability of a specific structure or regime for each observation t. If the structural change occurs abruptly at t_0, then

$$D_t = \begin{cases} 0 & \text{if } t \le t_0 \\ 1 & \text{otherwise} \end{cases}$$

On the other hand, if the structural change is gradual, the value of D_t should increase gradually from 0 to 1 as t increases from 2 to T. Since t_0 is unknown, D_t must be estimated for each observation. Goldfeld and Quandt [10, p. 10] have suggested an approximation of the dichotomous variable D_t by a continuous function such as,

$$D_t(t_0, \sigma_0^2) = \int_{-\infty}^{t} \frac{1}{\sqrt{2\pi}\sigma_0} \exp\left[\frac{-(\tau - t_0)^2}{2\sigma_0^2}\right] d\tau \tag{2}$$

$D_t(t_0, \sigma_0^2)$ is the normal integral with two parameters t_0 and σ_0^2. Therefore, $D_t(t_0, \sigma_0^2) \le 0.5$ if $t \le t_0$ and > 0.5 otherwise. The timing of the switch is therefore identified by a value of t so that $D_t(t_0, \sigma_0^2) = 0.5$, and the magnitude of the variance of the time shift σ_0^2 is used to indicate the length of the structural transition. If σ_0^2 is statistically different from 0, the null hypothesis of an abrupt structural change should be rejected—that is, a gradual change in the structure is identified around t_0.[5]

[4] The definition of money and the arguments in the money demand function are issues that have been addressed by numerous studies. We selected for this study the standard (or received version) of the quarterly money demand function as depicted by Equation (1) in Judd and Scadding [18, p. 995].

[5] Goldfeld and Quandt [10] have pointed out that $D_t(t_0, \sigma_0^2)$ can be interpreted as a device to allow for a hybrid transition to exist between two pure regimes. The magnitude of σ_0^2 provides an estimate of the success in discriminating between the two regimes. Therefore, the value of σ_0^2 departing from 0 serves to identify the type of structural change, i.e., abrupt or gradual.

The use of a continuous approximation $D_t(t_0, \sigma_0^2)$ for D_t presents some complications in the error structure which may be different for different regimes (see Goldfeld and Quandt [10, pp. 15–18]). When a structural change occurs at t_0, it is likely that the assumed first-order autoregressive error structure will also change. Therefore, the transition of the error structure has to be built into the model using the continuous variable $D_t(t_0, \sigma_0^2)$. If we are certain that t and $t-1$ are in the same regime i, then $\varepsilon_t^{(i)} = \rho^{(i)}\varepsilon_{t-1}^{(i)} + u_t^{(i)}$; otherwise, $\varepsilon_t^{(i)} = u_t^{(i)}/(1-\rho^{(i)2})^{1/2}$. The latter transformation is due to a reinitialization of the error term when the structure shifts (see Goldfeld and Quandt [10, p. 31]).[6]

With the approximation of the D_t variable shown in Equation (2) for structure discrimination, and using the convenient notations of $\theta = (\beta^{(1)}, \gamma^{(1)}, \rho^{(1)}, \sigma^{2(1)}, \beta^{(2)}, \gamma^{(2)}, \rho^{(2)}, \sigma^{2(2)}, t_0, \sigma_0^2)$ and $Z_t = (X_t, m_{t-1}, m_{t-2})$, the two-regime money demand model can be combined as follows:

$$
\begin{aligned}
m_t &= g_t(Z_t; \theta) + w_t, \qquad \text{where for} \quad t = 3, \cdots, T \\
g_t(Z_t; \theta) &= (1 - D_t)[X_t\beta^{(1)} + \gamma^{(1)}m_{t-1} + (1 - D_{t-1})\rho^{(1)}(m_{t-1} \\
&\qquad - X_{t-1}\beta^{(1)} - \gamma^{(1)}m_{t-2})] \\
&\quad + D_t[X_t\beta^{(2)} + \gamma^{(2)}m_{t-1} + D_{t-1}\rho^{(2)}(m_{t-1} - X_{t-1}\beta^{(2)} - \gamma^{(2)}m_{t-2})]
\end{aligned}
$$

and

$$
\begin{aligned}
w_t &= (1 - D_t)[(1 - D_{t-1}) + D_{t-1}/(1 - \rho^{(1)2})^{1/2}]u_t^{(1)} \\
&\quad + D_t[D_{t-1} + (1 - D_{t-1})/(1 - \rho^{(2)2})^{1/2}]u_t^{(2)} \qquad (3)
\end{aligned}
$$

Furthermore, the usual initialization for $t = 2$ gives

$$g_2(Z_2; \theta) = (1 - D_2)(X_2\beta^{(1)} + \gamma^{(1)}m_1) + D_2(X_2\beta^{(2)} + \gamma^{(2)}m_1)$$

and

$$w_2 = (1 - D_2)u_2^{(1)}/(1 - \rho^{(1)2})^{1/2} + D_2u_2^{(2)}/(1 - \rho^{(2)2})^{1/2}$$

Note that $Z_2 = (X_2, m_1, m_0)$, and m_0 is a null entry.

For $i = 1, 2$, and $t = 2, \cdots, T$, we assume that $u_t^{(i)}$ are independently and normally distributed with mean zero and variance $\sigma^{2(i)}$. The variance of the combined error w_t is obviously heteroscedastic as follows:

$$
\begin{aligned}
\sigma_{w_t}^2 &= (1 - D_t)^2[(1 - D_{t-1}) + D_{t-1}/(1 - \rho^{(1)2})^{1/2}]^2\sigma^{2(1)} \\
&\quad + D_t^2[D_{t-1} + (1 - D_{t-1})/(1 - \rho^{(2)2})^{1/2}]^2\sigma^{2(2)} \qquad (4)
\end{aligned}
$$

for $t = 3, \cdots, T$, and

$$\sigma_{w_2}^2 = (1 - D_2)^2\sigma^{2(1)}/(1 - \rho^{(1)2}) + D_2^2\sigma^{2(2)}/(1 - \rho^{(2)2})$$

Accordingly, for $t = 2, \cdots, T$, the conditional density of m_t given Z_t and θ is

[6] This formulation also conforms to the maximum likelihood estimation of Beach and MacKinnon [1] for single equation models with autocorrelation.

the following:

$$f(m_t \mid Z_t; \theta) = \frac{1}{\sqrt{2\pi}\sigma_{w_t}} \exp\left[- \frac{(m_t - g_t(Z_t; \theta))^2}{2\sigma_{w_t}^2} \right]$$

Therefore, the log-likelihood function is

$$\log L(\theta; Z_2, \cdots, Z_t) = \textstyle\sum_{t=2}^{T} \log f(m_t \mid Z_t; \theta)$$
$$= -\frac{T-1}{2}(\log 2\pi) - \tfrac{1}{2} \textstyle\sum_{t=2}^{T} \left[\log \sigma_{w_t}^2 + \frac{(m_t - g_t(Z_t; \theta))^2}{\sigma_{w_t}^2} \right] \quad (5)$$

Maximizing Equation (5) with respect to the parameter vector θ gives the regression coefficient estimates of the two structures along with their asymptotic standard errors. The latter are obtained as square roots of the diagonal elements of the negative inverse Hessian of the log-likelihood function evaluated at the maximum likelihood estimator of θ (see Goldfeld and Quandt [8, Chap. 2] and Malinvaud [22, Chap. 9]).[7]

From the parameter estimates obtained by the maximum likelihood procedure, the point of the time shift t_0 and its variance σ_0^2 provide the information on the discrimination between the two structures of the money demand equation. If t_0 takes a value either at the beginning or at the end of the sample period, $D_t(t_0, \sigma_0^2) = 0$ or 1 for almost all observations. The log-likelihood function (5) of the two-regime model then collapses to that of a single-regime case (see the Appendix).

Various hypothesis testings can be performed on the values of t_0 and σ_0^2 to determine the existence and the type of instability in the money demand equation. Moreover, by further restricting pairs of coefficients to be the same for the two structures, a nest of hypothesis testings can successfully identify the sources of the structural change.

II. Empirical Results

The direct maximization of the log-likelihood function (5) is implemented by a version of the numerical optimization routine DFP (Davidson-Fletcher-Powell algorithm) in the GQOPT package written by Goldfeld, Quandt, and Ertel.[8] To ensure that the solution obtained is indeed a global maximum, several different initial estimates are tried. In addition, the convergence criterion for the log-

[7] Alternatively, the explicit estimated standard errors can be calculated according to the square roots of the diagonal elements of the negative *expected* value of the inverse Hessian of the log-likelihood function. However, as pointed out by Cooper [3], Dhrymes [4], and Beach and Mackinnon [1], adjustments will be required for explicit standard errors when the lagged dependent variable is included as one of the explanatory variables in the regression equation. This adjustment process becomes intractable if there are two regimes of regression parameters linked by the continuous probability distribution, e.g., $D_t(t_0, \sigma_0^2)$.

[8] U.S. quarterly data for 1959(I) to 1981(IV) are used for this study. Seasonally adjusted data for the narrow definition of money, M_1, were provided by John Scadding of the Federal Reserve Bank at San Francisco; interest rates are from the Federal Reserve Bulletin; both GNP (seasonally adjusted) and CPI are obtained from the Survey of Current Business. All computations were performed on a Honeywell 60/20 computer system at Portland State University.

Table I

Two-Regime U.S. Short-Run Money Demand: 1959(II)–1981(IV)

	Coefficient		A	B	C(i)	C(ii)	D	E
			Specification					
Regime I	Constant	$\beta^{(1)}$	0.428	0.434	0.181	0.401	0.380	0.291
			(1.336)	(1.200)	(0.641)	(1.235)	(1.127)	(0.895)
	GNP		0.111**	0.086**	0.108**	0.108**	0.084**	0.102**
			(3.120)	(2.874)	(3.416)	(3.083)	(2.872)	(3.347)
	RCP		−0.017**	−0.018**	−0.018**	−0.017**	−0.017**	−0.015**
			(−3.168)	(−3.032)	(−4.784)	(−3.783)	(−3.928)	(−3.605)
	RSD		−0.020	—	−0.023	−0.019	—	—
			(−1.236)		(−1.425)	(−1.140)		
	RGB		0.0002	0.002	—	—	—	—
			(0.0126)	(0.133)				
	RCPP		—	—	—	—	—	−0.015*
								(−2.209)
	Lag m_1	$\gamma^{(1)}$	0.790**	0.816**	0.841**	0.799**	0.828**	0.827**
			(8.205)	(7.977)	(9.930)	(8.156)	(8.431)	(8.657)
	Serial Corr.	$\rho^{(1)}$	0.354**	0.364**	0.258	0.348**	0.353**	0.286
			(2.451)	(2.426)	(1.789)	(2.514)	(2.479)	(1.907)
	Reg. Var. (10^{-5})	$\sigma^{2(1)}$	1.795	1.835	1.909	1.784	1.819	1.677
Regime II	Constant	$\beta^{(2)}$	1.582*	1.565*	−5.595	0.101	−0.018	1.946**
			(2.183)	(2.234)	(−1.013)	(0.136)	(−0.040)	(2.530)
	GNP		0.165**	0.170**	0.109	0.120	0.117	0.131**
			(2.896)	(2.628)	(1.145)	(1.788)	(1.849)	(2.912)
	RCP		0.003	0.002	−0.084	−0.026**	−0.027**	−0.010
			(0.226)	(0.155)	(−1.325)	(−2.489)	(3.200)	(−1.208)
	RSD		−0.0003	—	0.182	−0.029	—	—
			(−0.0025)		(0.571)	(−0.233)		
	RGB		−0.075**	−0.075**	—	—	—	—
			(−2.754)	(−2.660)				
	RCPP		—	—	—	—	—	−0.084**
								(−2.887)

Lag m_1	$\gamma^{(2)}$	0.517**	0.514**	0.546	0.840**	0.858**	0.510**
		(3.296)	(3.160)	(1.421)	(7.935)	(9.638)	(3.376)
Serial Corr.	$\rho^{(2)}$	−0.313	−0.314	−0.731	−0.261	−0.268	−0.264
		(−1.530)	(−1.583)	(−1.648)	(−1.294)	(−1.504)	(−1.176)
Reg. Var (10^{-5})	$\sigma^{2(2)}$	4.174	4.242	17.988	5.654	5.691	4.120
Time Shift	t_0	62.92**	63.01**	86.16**	63.13**	63.02**	63.43**
		(48.940)	(44.161)	(8.331)	(31.039)	(29.885)	(50.115)
Shift Variance	σ_0^2	3.118	4.225	205.928*	9.294	10.479	7.358
		(0.506)	(0.467)	(2.080)	(0.587)	(0.830)	(0.893)
Max. Log-Likelihood		356.65	355.94	355.11	353.14	352.47	359.76
Dynamic Simulation	R^2	0.983	0.982	0.788	0.766	0.622	0.984

Note: Numbers in parentheses are asymptotic t-ratios.
* Denotes significant coefficient at 5% level.
** Denotes significant coefficient at 1% level.

likelihood function value is set at a very small value, 10^{-8}, with vanishing gradient at the maximum.

Table I reports the estimation results for the two-regime case for alternative specifications (A–E) of the U.S. short-run money demand equation over the period 1959(II)–1981(IV). In specification A, three interest rates, RCP, RSD, and RGB (commercial paper, savings deposits, and long-term government bond rates) enter as arguments in addition to real GNP and the lag dependent variable. The coefficient estimates of specification A are almost identical to specification B where RSD is excluded due to its insignificance in the two regimes of the former. Specifications A and B provide an interesting finding regarding the role of the short-term and long-term interest rates.[9] The commercial paper rate (RCP) dominates the first regime where the long-term government bond rate is insignificant. However, in the second regime, the long-term government bond rate becomes the only significant interest rate variable reflecting the opportunity cost of holding money. It should also be noted that the adjustment speed (one minus the coefficient on lagged m_1) is much faster in the second regime, and that the estimated shift point t_0 is the 63rd observation, or 1974(III).

In specification C, only two interest rates, RCP and RSD are included to reflect the opportunity cost of holding money (see Hafer and Hein [14] and Laumas and Spencer [21]). Two inconsistent solutions are obtained for this specification and they are reported as C(i) and C(ii). C(i) indicates that all coefficients are insignificant in the second regime and the shift point is the 86th observation, or 1980(II). C(ii) is comparable to specifications A and B for the first regime but differs substantially for the second regime where only the RCP and lagged dependent variables are significant. The shift point for C(ii) is 1974(III)—similar to that obtained for specifications A and B—and the lag coefficients are not substantially different in each of the two regimes.

Since the variable RSD is insignificant in specifications A and C for both regimes, this variable is excluded in specification D leaving RCP as the only variable reflecting the opportunity cost of holding money (see Laumas and Spencer [21]). Although the results in the first regime are quite similar to those of specifications A and B, they are quite different in the second regime where the income coefficient is insignificant and there does not appear to be a substantial difference in the adjustment speed. The shift point, however, remains at 1974(III).

Improved cash management techniques by money holders through financial innovations have been argued by some researchers to be an important determinant in the money demand function (see Porter, Simpson, and Mauskopf [23], Simpson and Porter [24], and Judd and Scadding [18]). Specification E presents the results where the previous peak of RCP, denoted as RCPP, is included as a proxy variable for the improved cash management technique (see Goldfeld [7] and Judd and Scadding [18]).[10] Results for this specification reveal characteristics

[9] For recent reviews on the issue of incorporating the short- and/or long-term rates, see Judd and Scadding [18] and, especially, Hafer and Hein [11].

[10] RGB and RSD were excluded from this specification because of multicollinearity with RCPP. Alternative specifications of the cash management variable such as the one used by Porter and Simpson—a ratchet on the 5-year treasury notes—has been tried, however, the variable did not perform any better than the previous peak of RCP.

Table II

Hypothesis Testings on the Existence of Structural Shift

Model Specification*	Likelihood Ratio Test $-2\log(L_1^*/L_2^*)$	Degree of Freedom	Critical Value $\chi^2_{0.01}$ (d.f.)
A	52.74	10	23.21
B	53.82	9	21.67
C(ii)	46.54	9	21.67
D	47.42	8	20.09
E	47.30	9	21.67

* Specification C(i) is not included in this test.

that are similar to specifications A and B. In addition to RCP, the previous peak of RCP is significant at the 5% level in the first regime but it becomes the only significant interest rate variable at the 1% level in the second regime. Further, the shift point is 1974(III) and the adjustment speed is much faster in the second regime.

In all of the above specifications except for C(i), the common point of time shift is the 63rd observation or 1974(III). Whether or not this shift point is legitimate can be further confirmed by the likelihood ratio test (Goldfeld and Quandt [10]) and the asymptotic t-ratio test. The likelihood ratio test statistic, $-2\ln(L_1^*/L_2^*)$, where L_1^* and L_2^* are the maximum likelihood values obtained from the one-regime and the two-regime models, respectively, can be used to test the null hypothesis of a one-regime demand function.[11] The likelihood ratio test statistic follows a chi-square probability distribution with the appropriate degrees of freedom. These test results are presented in Table II where it is immediately clear that the null hypothesis of a one-regime money demand equation has to be rejected for all specifications at the 1% significance level. Additionally, it should be noted that if t_0 takes a value of 2 (the first observation of the time series) or 92 (the last observation), the log-likelihood function for the two-regime case collapses into that of a one-regime case because $D_t(t_0, \sigma_0^2) = 0$ or 1 for all t. A simple t-ratio test can be applied to the estimator of shift point t_0 against its extreme values of 2 or 92. Based on the reported results in Table I, it is easy to show that t_0 for all specifications is significantly different from 2 or 92, thus, confirming the two-regime case.[12] Furthermore, the shift variance, σ_0^2 is not significantly different from zero at the 1% level. This indicates that the shift in 1974(III) is of a fairly abrupt type.

We can conclude at this point that both the likelihood ratio test and the asymptotic t-ratio test support the findings of instability of the money demand equation by other studies. However, the use of the switch regression technique

[11] Results of the one-regime case are available from the authors.

[12] Model specification C(i) is the exception. The results indicate a significant shift at the 86th observation; however, it should be noted that the time shift at the 86th observation is not significantly different from the 92nd observation. Further, a model with three regimes was also constructed and estimated, but no sensible results were obtained due to the lack of degrees of freedom for the possible third regime.

provides information on the complete regression estimates for two regimes including the estimated point of shift and the type of structural change.

To examine the results further, dynamic simulations on all five specifications are performed to aid in the selection of our preferred specifications. The simulation is dynamic in the sense that the dependent variable is simulated for the entire sample by using actual values of the independent variables with the exception of the lagged dependent variable. For the first observation ($t = 2$), the actual value of the lagged dependent variable is used; thereafter ($t = 3, \cdots, T$), the previous period's simulated value for this variable is utilized. Specifications C and D are immediately ruled out as the preferred model due to their lower squared correlation between actual and simulated (see Table I). Specification A, although similar to specification B, is also ruled out because there are too many insignificant coefficients. Specifications B and E are the remaining ones that are selected for further analysis and discussion.

When the long-term elasticity and the mean and median lags are computed for the significant coefficients in specifications B and E, the differences between the elasticities and lag coefficients in the first and second regimes become noticeable. For example, income elasticity is smaller in the second regime, and the long-term interest rate and previous peak of RCP are more elastic in the second regime in their respective specifications. Most noticeable, however, is the adjustment of money demand towards its long-run equilibrium. The speed of adjustment is about three to four times faster in the second regime, i.e., after 1974(III) for both specifications. For instance, the mean and median adjustment lags are 4.435 and 3.407 quarters before 1974(III) and dramatically reduces to 1.058 and 1.041 quarters after the shift for specification B. This characteristic also holds for specification E.

It should be further pointed out that the estimated coefficients of the preferred specifications B and E in Table I for regime I are consistent with results obtained by previous studies for the sample period before 1974(III). However, after the shift, the size of the coefficients changes substantially and this change should be recognized in the literature. The rather dramatic changes in coefficient sizes cannot be disguised by respecifications of the conventional money demand equation or through the use of more sophisticated estimation techniques (e.g., Hatanaka [16]).

III. Sources of the 1974 Shift

In a recent article by Hafer and Hein [15], it was argued that the 1974 shift in the conventional money demand function was due to a change in the constant term and, thus, a once-for-all level shift is the logical conclusion. The level shift is identified through the introduction of a dummy variable in the regression equation and, providing a shift point, a priori. In general, the method of identifying the type of shift has two shortcomings. First, prior information is required regarding the point in time when the shift occurred and this shift is by definition an abrupt type. Second, the method implicitly assumes that the model is not contaminated. This second shortcoming is more serious because contamination is defined here as the existence of different error structures for the sample

Table III

Hypothesis Testings on the Sources of Structural Shift

Null Hypothesis (Restriction)		Likelihood Ratio		Critical Value	
		Specification B	Specification E	$\chi^2_{0.05}$ (1)	$\chi^2_{0.01}$ (1)
Constant	$\beta_1^{(1)} = \beta_1^{(2)}$	2.24	4.50*		
GNP	$\beta_2^{(1)} = \beta_2^{(2)}$	2.46	5.88*		
RCP	$\beta_3^{(1)} = \beta_3^{(2)}$	1.84	0.36		
RGB or	$\beta_4^{(1)} = \beta_4^{(2)}$	6.04*	—	3.84	6.63
RCPP		—	6.46*		
Lag M_1	$\gamma^{(1)} = \gamma^{(2)}$	2.72	3.56		
Serial Corr.	$\rho^{(1)} = \rho^{(2)}$	6.98**	3.98*		
Reg. Var.	$\sigma^{2(1)} = \sigma^{2(2)}$	7.04**	8.16**		

* Denotes the rejection of hypothesis at 5% significance level.

** Denotes the rejection of hypothesis at 1% significance level.

periods. Contamination is a common phenomenon in a model with more than one regime where the serial correlations and variances of the regimes are not identical. If a model is indeed contaminated, an assumption of no contamination is tantamount to a model misspecification.

The use of deterministic switch regression allows us to statistically identify and test for the source(s) of the structural shift. In general, our results indicate that the shift is not necessarily due to the "downward shifting" of the constant term. In fact, our results show that the second regime has a much larger constant term than that of the first regime, contradicting the "once-for-all downward shift" in the constant term during the mid to late 1970s noted by Hafer and Hein [12, 13, 15]. The main problems of the instability of the short-run money demand result from the inequality of the serial correlation coefficients and the estimated variances in each of the two regimes. This suggests that the problems may be due to missing variables (e.g., the inclusion of the entire term structure of interest rates, see Heller and Khan [17]), or the possible existence of another regime which we are unable to identify because of data limitations. Nevertheless, it is this problem of contamination that limits the application of the Brown-Durbin-Evans technique and other statistical methods to test the stability of the short-run money demand equation.[13]

A finding of contamination implies that the sources of instability of the conventional money demand can be due to any form of misspecification. However, because a recent study (Hafer and Hein [15]) has determined that the shift in the mid-1970s was an intercept rather than a slope shift, we confined our search for possible sources to only the variables we have used. In other words, tests of possible sources of the structural shift are confined to examining the impact of restricting the estimated individual coefficients for the two regimes to the same value. To test for the possibility of changes for the individual coefficients, a likelihood ratio test statistic $-2\ \ln(L_R^*/L_u^*)$ is constructed. L_u^* is the maximum

[13] BDE technique tests for changes in the coefficient estimates of different structures, provided that the error structure and variances for different regimes of the relationship are identical. The problem of model contamination is not addressed in the BDE method.

likelihood value of the unrestricted model and L_R^* is the corresponding maximum likelihood of the restricted model. The likelihood function (5) is then remaximized, subject to the parameter equality constraints. It is implied that the test statistic has an asymptotic chi-square distribution of one degree of freedom.

The results of these tests for the various coefficients in specifications B and E are reported in Table III. For both specifications, the null hypothesis that the coefficients are identical in the two regimes are clearly rejected for the coefficients of serial correlation and the regression variances—confirming our previous observation about model contamination in Table I. The long-term government bond rate in specification B is significantly different at the 5% level for the two regimes. Similarly, for specification E, the previous peak of RCP rate is significantly different at the 5% level for the two regimes. However, we further note that in specification E the constant term and the scale variable are also significantly different at the 5% level. It is interesting to point out that the short-term interest rate is not significantly different in two regimes for both specifications B and E.

Finally, we note that there is no evidence to indicate that the lag coefficient or the adjustment speed is significantly different for the two regimes even though the speed of adjustment towards equilibrium money holdings is considerably faster in the period after 1974(III).

IV. Conclusion

From this exercise, we conclude that the U.S. short-run money demand function is contaminated. Because contamination can be any form of misspecification, a complete search for the sources of the 1974(III) shift would be beyond the scope of this study. We have, instead, limited the search for possible sources of shift to the investigation of coefficient changes in the two regimes of our preferred specifications. Our findings do not support the contention that the 1974 shift is a downward shift in the constant term. Rather, on the basis of our limited investigation, the shift is due to the long-term interest rate for one preferred specification, and the constant term, scale variable, and the proxy variable for improved cash management technique for the other preferred specification. We cannot, however, minimize the important fact that both specifications are contaminated.

Caution must be exercised when interpreting the results of other studies that use the methods of BDE, VPR, Chow test, and the dummy variable method since these studies are not free of possible misspecifications. Our findings of contamination in the short-run money demand function highlight the need for additional studies to determine level and/or marginal shifts and the reasons underlying the shift.

Appendix

The single-regime money demand model can be written as:

$$m_t = X_t\beta + \gamma m_{t-1} + \varepsilon_t$$

and

$$\varepsilon_t = \rho\varepsilon_{t-1} + u_t, \qquad t = 2, \cdots, T$$

The corresponding log-likelihood function via Beach and MacKinnon [1] is the following:

$$\log L(\theta; Z_2, \cdots, Z_T) = -\frac{T-1}{2}(\log 2\pi + \log \sigma^2) + ½ \log(1-\rho^2)$$

$$-\frac{1}{2\sigma^2}(1-\rho^2)[m_2 - h_2(Z_2; \theta)]^2 + \textstyle\sum_{t=3}^{T} [m_t - h_t(Z_t; \theta)]^2$$

where $\theta = (\beta, \gamma, \rho, \sigma^2)$, $h_2(Z_2; \theta) = X_2\beta + \gamma m_1$, and $h_t(Z_t; \theta) = X_t\beta + \gamma m_{t-1} + \rho(m_{t-1} - \beta X_{t-1} - \gamma m_{t-2})$, for $t = 3, \cdots, T$. It is easy to show that the two-regime log-likelihood function (5) collapses to the above single-regime case when the regime discrimination variable $D_t(t_0, \sigma_0^2) = 0$ or 1 (see (2)) for each observation t. Regression estimates of single-regime money demand equation for various specifications can be obtained from the authors. In general, all specifications suffer from problems of "missing money" (see Goldfeld [7] and Judd and Scadding [18] for discussions).

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