WP/00/188

IMF Working Paper

Endogenous Money Supply and Money Demand

Woon Gyu Choi and Seonghwan Oh

INTERNATIONAL MONETARY FUND

 WP/00/188

IMF Working Paper

IMF Institute

Endogenous Money Supply and Money Demand

Prepared by Woon Gyu Choi and Seonghwan Oh[1]

Authorized for distribution by Reza Vaez-Zadeh

November 2000

Abstract

This paper explores the behavior of money demand by explicitly accounting for the money supply endogeneity arising from endogenous monetary policy and financial innovations. Our theoretical analysis indicates that money supply factors matter in the money demand function when the money supply partially responds to money demand. Our empirical results with U.S. data provide strong evidence for the relevance of the policy stance to the demand for M1 under a regime in which monetary policy is substantially endogenous. Specifically, we find that tighter monetary policy has substantial positive impacts on money demand under the recent Federal funds rate targeting.

JEL Classification Numbers: C22; E41; E52

Keywords: Money Demand; Money Supply Endogeneity; Financial Innovation; Monetary Policy Stance

Author's E-Mail Address: wchoi@imf.org; shoh@plaza.snu.ac.kr

[1] The authors thank seminar participants at the Seoul Macro Workshop, the Hong Kong University of Science & Technology, Seoul National University, SungKyunKwan University, and the 1999 summer meeting of the Econometric Society in Singapore for helpful discussions and comments. The authors are also grateful to Stanley W. Black, Leonard Cheng, Jang-Ok Cho, David Cook, Michael Devereux, Dmitriy Gershenson, Joon Y. Park, Changyong Rhee, Keunkwan Ryu, Reza Vaez-Zadeh, and Danyang Xie for valuable comments and suggestions. The first draft of this paper was written when the first author was at the Hong Kong University of Science & Technology.

Contents | Page

I. Introduction

The conventional money demand literature has treated the money supply as exogenously given. It is well known, however, that some endogeneity of the money supply arises due to financial innovations and endogenous monetary policy. Dotsey (1984) shows empirically and Ireland (1992) shows theoretically that financial innovations affect money demand. On the other hand, Laidler (1993, p. 187) insightfully suggests that considerable turbulence in the conduct of monetary policy may have affected the stability of money demand since the 1970s. Endogeneity of the money supply renders the conventional money demand function misspecified (e.g., Cogley, 1993) and, indeed, may have been the primary cause of money demand instability during the past two decades as shown in Goldfeld and Sichel (1990).

In fact, the Federal Reserve ("Fed") has targeted the Federal funds rate since the mid-1980s (see Sellon, 1994; Rudebusch, 1995; Meulendyke, 1998). This interest rate targeting makes the money supply respond directly to interest rate changes, as shown by Poole (1970). Empirical studies on policy rules suggest that the Fed has been anti-inflationary and counter-cyclical in setting the Federal funds rate and discount rate (e.g., Taylor, 1993; Choi, 1999). Given the importance of interest rate targeting and of the (counter-cyclical) policy stance, it is natural to presume that agents' perceptions of endogenous monetary policy affect their money demand.

Christiano et al. (1996) and Choi and Kim (2000) support this presumption empirically. Christiano et al. find from a flow of funds data analysis that firms initially raise their net funds upon a tight monetary shock. This finding is corroborated by Choi and Kim, who prove with panel data that tighter monetary policy causes corporate firms initially to increase liquid asset holdings. Choi and Kim suggest that loan commitments and sluggish loan rate adjustments may motivate firms to preempt funds at low costs.

In this paper, we examine an explicit channel through which monetary policy affects money demand, in addition to endogenous financial innovations. The paper sheds light on the interactions between investors' money holdings and the monetary authority's policy rule. Investors are assumed to know that the Fed's money supply is contingent upon demand pressures on funds to a substantial extent (as implied by the real bills doctrine). Then we show that investors, utilizing the Fed's money supply response to money demand pressures, initially raise money demand upon tighter policy with a preemptive motive to mitigate the policy impact on money balances.

Specifically, in line with the general equilibrium literature that incorporates money in utility function (MIUF, e.g., Danthine and Donaldson, 1986; Boyle, 1990; Lucas, 2000), we attribute the utility gains from money to liquidity services that are provided by the real balance and financial service acquired before transactions. We account for endogenous financial innovations that facilitate transactions (e.g., through the use of credit cards in a wider range of transactions) by introducing in the model the accumulation of the financial capital stock. The financial service, which can be viewed as a substitute for money, is determined by the financial capital stock that is operational during the period. Also, we explicitly incorporate the endogeneity of monetary policy by introducing a policy rule that accounts for the Fed's response to money demand. We impose a timing structure as follows. A representative investor chooses, based on the expectation of the monetary authority's action, money holdings and all other financial assets in the first part of the period. In the second half-period, the authority sets the money supply based on the perceived investor's money demand and its policy stance. Upon observing the money supply shock, the investor decides on consumption and the financial capital expenditure, and adjusts money holdings. The investor can reallocate money holdings between the first and second half-periods at an adjustment cost if their first-half-period money holdings differ from the available funds set by the authority.

We show that the derived money demand function contains a money supply factor in addition to the variables of the conventional money demand function. In particular, tighter monetary policy increases an investor's money demand if the money supply substantially responds to money demand and if the adjustment cost of the money holdings is not prohibitively high. We also show that parameters of the money demand function depend on the degree of the money supply endogeneity, implying that they vary as the policy regime shifts (e.g., from monetary aggregate targeting to interest rate targeting).

We investigate the derived money demand using quarterly and monthly U.S. data. Using cointegration methods, we first examine a long-run relationship among level variables. Then the effect of a money supply factor is evaluated by measuring how much the monetary policy stance affects short-run deviations from the long-run relationship. The policy stance is measured either by the change in the Federal funds rate or by the Boschen-Mills (1995) index. The results provide strong evidence for the direct impact of the policy stance on money demand. Specifically, we find that tighter monetary policy increases the demand for M1 most significantly during the recent Federal funds rate targeting period—a percentage point rise in the funds rate change increases money demand by about 7 percent in the same month. This implies that money supply endogeneity, which can be attributed to the Fed's accommodation of the pressures on money demand, motivates agents to hold money with a preemptive motive to hedge against policy shocks. Furthermore, the initial rise in money balances due to tighter policy accompanies an increase in interest rates. This comovement can show up as a positive contemporaneous correlation between (broad) money and interest rates—the so-called 'liquidity puzzle' (e.g., Gordon and Leeper, 1994). An important source of that correlation may lie in the money supply endogeneity, through which tighter policy exerts a positive impact on money holdings.

This paper proceeds as follows. In section 2, we set up a general equilibrium model and show how the demand for money depends on money supply factors. In section 3, using U.S. data, we estimate the postulated money demand and evaluate the monetary policy effect on money demand. We also provide robustness checks. In section 4, we provide the concluding remarks. The appendix contains the derivation of the equations, data descriptions, and other details.

II. THEORETICAL BACKGROUND: THE MODEL

We consider a simple monetary economy in which the only source of uncertainty is the monetary policy stance, and the real output process is exogenously given. There are a finite number of investors and the population size is normalized to one. Each period consists of two halves. In the first half-period, after output and financial services are produced, a representative investor chooses the first-half-period money holdings and other assets. In the second half-period, the money supply is set, the commodity market and the financial capital good market are open, and the investor chooses consumption, investment into the financial capital stock, and the second-half-period money holdings.

The per-investor money supply in period t, M_t, is governed by the following rule:

$$M_t = \overline{P}_t (T_t^Y)^{\alpha} (L_t / \overline{P}_t)^{1-\alpha} e^{-\lambda H_t}, \ \mu \geq 0, \ 0 < \alpha < 1, \ \lambda > 0, \tag{1}$$

where $\overline{P}_t$ is the Fed's presumed price level, T_t^Y is a trend consistent with the long-run output, $L_t / \overline{P}_t$ is the first-half-period money demand perceived by the Fed, and H_t is the policy stance following a stationary process with zero mean.[2] α and $1-\alpha$ indicate the elasticities of the money supply with respect to the trend and to the first-half-period money demand, respectively. We assume that M_t is neither fully endogenous nor fully exogenous ($0<\alpha<1$). Money supply rule (1) comprises a trend component, T_t^Y, reflecting a Friedman-type money growth rule if we set $T_t^Y = \overline{m}e^{\mu}$, an accommodative component reflecting the money supply on demand[3], and an exogenous component due to the Fed's policy stance or shock.

The Fed has imperfect knowledge about the investor's decision on the second-half-period money holdings due to information asymmetry. The Fed can make forecast errors for

[2] More generally, H_t may also depend on inflation and output growth that are predetermined as of the time when the Fed sets the money supply as well as a stochastic shock. This extension, nonetheless, does not alter our analysis.

[3] A way to rationalize the Fed's response to the investor's demand is to consider the delegation role played by a small number of banks in allocating money from the single supplier (the Fed). Each bank represents the investors' money demand collectively and is large enough to call for the Fed's response. With such a delegation, the investor takes into account the Fed's response in making portfolio decisions. Also, upon a shock identified by investors as global, each investor may adjust portfolio with the perception that other investors also have such adjustments.

the price level.[4] Suppose that $\overline{P}_t$ is a predetermined variable consistent with a long-run target level, P_t^*, i.e., allowing for the partial adjustment to the long-run target level, $\overline{P}_t = (1-g)P_t^* + g\overline{P}_{t-1}$ $(0 < g < 1)$.[5] We assume that $P_t = \overline{P}_t e^{w_t}$, where P_t is the price level that balances the supply of money and the demand for money and w_t is a stochastic error.

Also note that policy rule (1) reconciles the notion that the Fed follows a kind of the Taylor rule, which is known to perform well since the mid-1980s (Taylor, 1993; Judd and Taylor, 1998). The Taylor rule suggests that the Fed should *increase* the interest rate, resulting in a decrease in *M*, if inflation rises above a target or if output is above its trend. Similarly, policy rule (1) requires a tightening policy with respect to higher inflation and a loosening policy with respect to larger long-term output.[6] As regards a demand-side real factor, the Taylor rule requires the Fed to increase the interest rate as output exceeds its trend, resulting in a decrease in *M*. Policy rule (1), however, does not consider this case since the output process here is endowment determined, while it can be extended to reflect a tighter policy stance in response to higher output growth to reduce *M* (see footnote 2).

The representative investor's preferences over consumption and liquidity are given by

$$E_0 \sum_{t=0}^{\infty} \beta^t U(X_t, \overline{L_t / P_t}), \tag{2}$$

where E_0 is the expectations operator conditional on period 0 information and β is a discount factor. The utility function is assumed to be [7]

$$U(X_t, \overline{L_t / P_t}) = s \ln X_t + (1-s)\ln(\overline{L_t / P_t}),\ 0 < s < 1,$$

[4] The information asymmetry assumption is plausible since the investor faces many specific factors that determine the external finance premium or the adjustment cost of funds. Even without the information asymmetry, price forecast errors appear unless the presumed price level is exactly consistent with the money supply rule that supports the long-run output and the target price over the long run.

[5] Money supply rule (1) is similar to that in Ramey (1993) or Goodfriend (1997) in that it supports stable prices. Unlike Ramey and Goodfriend where price rigidity is introduced, our rule includes $\overline{P}_t$ instead of P_t.

[6] Policy rule (1) requires the Fed to reduce *M* as the interest rate rises and thus money demand falls with inflation, and to increase *M* as the long-term output T^Y increases.

[7] The results for the money demand function in this paper remain qualitatively unaffected when constant relative risk aversion is assumed.

where X_t and $\overline{L_t / P_t}$ denote real consumption and *effective* real liquidity in period *t*, respectively. Let $\overline{L_t / P_t} \equiv L_t / P_t + nF_t$, where L_t indicates the first-half-period money balance, and F_t is the real financial service. We assume that financial services are imperfect substitutes for money balances ($0 < n < 1$). We attribute the utility gains from money to liquidity services that are provided by money balances and financial services, both of which are acquired *in advance*. This specification of money can be regarded as imposing a weak cash-in-advance constraint since the utility function is specified in a more general form (see Feenstra, 1986).

The representative investor enters each period *t* carrying over money balances and bonds from the previous period, a proportion Z_{t-1} of the equity claim to all future output, and the nominal receipt of dividends, $P_t Y_t Z_{t-1}$. In the first half-period, the investors borrow and lend among themselves by trading one-period nominal discount bonds, which are available in zero net supply. The investor purchases equity claim Z_t at the price of Q_t, discount bonds D_t at the nominal interest rate R_t, and chooses the first-half-period money balance L_t. The output Y_t and the financial service F_t are produced: Y_t is not storable for future consumption but can be invested into the financial capital, and F_t is perishable. In the second half-period, the government makes a monetary injection, ΔM_t, in the form of a lump sum transfer to the investor. After the monetary injection, the investor chooses consumption and investment into the financial capital and adjusts money holdings. The investor ends up with the total money holdings, $L_t \tau_t$, in the second half-period, where τ_t (>0) is the ratio of the total money holdings to the first-half-period money holdings.

The accumulation of the financial capital stock K_t is given by

$$K_t = (1-\delta)K_{t-1} + I_t, \; K_0 \geq 0, \tag{3}$$

where I_t is the investment in period *t* and $\delta \in (0,1)$ is the depreciation rate. The financial service flow F_t is determined by the financial capital stock operational during the period,

$$F_t = F(K_{t-1}) = K_{t-1}^{\psi}, \; 0 < \psi < 1. \tag{4}$$

Dotsey (1984) argues that a financial innovation involves a ratchet effect since it requires start-up costs and will remain in place until it is replaced by more advanced technology. In our model, past investments are subsumed into the current financial capital stock with a decaying factor and, hence, the financial service flow reflects Dotsey's ratchet effect.

Finally, a costly adjustment occurs when the first-half-period money holdings differ from the money supply. The cost function is assumed to be symmetric for simplicity

$$J_t = \frac{\theta}{2}\frac{L_t}{P_t}\left(\frac{M_t}{L_t} - 1\right)^2, \tag{5}$$

where $\theta (> 0)$ is a cost parameter and M_t, the money supply in period *t*. A change in the first-half-period money holdings induces a change in the money supply, which, however, may not where $\theta(>0)$ is a cost parameter and M_t, the money supply in period *t*. A change in the first-half-period money holdings induces a change in the money supply, which, however, may not accommodate fully the change in money demand. Discrepancies between the first-half-period money holdings and the money supply generate an adjustment cost. The adjustment cost reflects real resources spent for the reallocation to match the money holdings with the money stock available. The adjustment cost is assumed to be substantial so that it is not easy to nullify the first-half-period money demand. Otherwise, the investor can increase utility arbitrarily by increasing the first-period money balance and then shifting it into other uses in the second half-the period. Note that the investor cannot rearrange other financial assets.

Let's define $q_t \equiv Q_t / P_t$, $d_t \equiv D_t / P_t$, and $\pi_t \equiv P_t / P_{t-1}$. Then the budget constraint is

$$X_t + \frac{L_t}{P_t}\tau_t + q_t Z_t + d_t + I_t + \frac{\theta}{2}\frac{L_t}{P_t}(\frac{M_t}{L_t} - 1)^2 \tag{6}$$
$$= (Y_t + q_t)Z_{t-1} + \frac{M_t}{P_t} - \frac{1}{\pi_t}\frac{M_{t-1}}{P_{t-1}} + \frac{1}{\pi_t}\frac{L_{t-1}}{P_{t-1}}\tau_{t-1} + (1 + R_{t-1})\frac{1}{\pi_t}d_{t-1}.$$

The market-clearing conditions for the goods market, the money market, the equity market, and the credit market are $X_t + I_t = Y_t$, $L_t\tau_t = M_t$, $Z_t = 1$, and $D_t = 0$, respectively.

The first-order conditions of the constrained maximization problem and the market-clearing conditions yield the equilibrium conditions that are described in Appendix A. We assume, for simplicity, that the investor has perfect foresight.[8] Combining the equilibrium condition for the first-half-period-money holding and that for bond holdings (conditions (A1) and (A3) in the appendix), we obtain the following equilibrium relation:

$$\frac{1-s}{s}\frac{X_t}{L_t / P_t + nF_t} = \alpha\frac{R_t}{1+R_t}\tau_t + \Gamma(\tau_t), \tag{7}$$

where the marginal adjustment cost (dollar value) is given by $\Gamma(\tau_t) = \theta[(1/2-\alpha)\tau_t^2 - (1-\alpha)\tau_t + 1/2]$ and $\tau_t = M_t / L_t$. If $\alpha \to 1$ (i.e., the money supply becomes fully exogenous), money supply factors become irrelevant in equation (7) as $\tau_t \to 1$ (see condition (A5) in the appendix). Without financial innovations and adjustment costs (despite the money supply rule contingent on the first-half-period money holdings), equation (7) collapses to $M_t / P_t = \frac{(1-s)}{s\alpha} X_t (R_t / 1 + R_t)^{-1}$, similar to a conventional money demand function. Note that the opportunity cost of the first-half-period money holdings is multiplied by τ_t because an additional unit of L_t induces a change in M_t by the factor τ_t. Equation (7) indicates that the first-half-period money holdings, L_t / P_t, will be lowered with τ_t around the steady state unless the adjustment cost is excessively high, since the rise in the opportunity

[8] Under uncertainty, the model will provide more complicated expressions for money demand including covariance terms among the variables, but it will generate a similar result that money supply factors enter into the money demand function. Thus, the perfect foresight assumption is innocuous for the purpose of this paper.

cost dominates the decrease in the marginal adjustment cost as τ_t increases.[9] Since equation (1) implies that a policy tightening lowers the real money stock, tighter policy decreases τ_t and thus increases the first-half-period money demand. Equation (7) also indicates that the first-half-period money demand will be lowered with F_t.

Taking a log-linear approximation of equation (7) around the steady-state values of variables and using condition (A5) in Appendix A to eliminate τ_t, we obtain the money demand function

$$m_t - p_t = \beta_0 + \beta_x x_t - \beta_r r_t - \beta_f f_t + \beta_H H_t + e_t, \tag{8}$$

where the lower case letter indicates the logarithm of a variable (e.g., $m_t = \ln M_t$) and e_t is an error. The derivation of equation (8) is described in Appendix B. Coefficients β_x, β_r, and β_f are all unambiguously positive. The coefficient of H_t is given by $\beta_H = (\beta_f - \beta_x \alpha\theta\phi^{-1})\ (1-\alpha)^{-1}\lambda$, where ϕ>0. Given that the adjustment cost is substantial, β_H decreases with the money supply exogeneity α: specifically, $\partial\beta_H/\partial\alpha$<0 if θ>$\phi\beta_f\beta_x^{-1}$. Also, β_H approaches $\beta_f\lambda$(>0) as $\alpha \to 0$ and remains positive as long as the money supply endogeneity is substantial and the adjustment cost is not too excessive.[10]

The derived money demand function has the following characteristics. First, the money supply process affects money demand. In particular, when anticipating tighter monetary policy, the investor increases the first-half-period money holdings if the adjustment cost of the money holdings is not too excessive. The total money holdings, M_t / P_t, increase upon tighter policy when the increase in the first-half-period money demand outstrips a fall in τ_t given that the financial service cannot be adjusted during the period. Second, coefficients in the money demand function including β_0 and β_H depend on the money supply endogeneity (1-α) and thus shift as α varies over regimes (see, for details, Appendix B).

For the robustness check of the derived money demand function, we consider two alternative model specifications. First, we may consider an endogenous output process introducing a production technology. Such an output process, however, does not affect the investor's decision rules much, and thus the policy stance still affects money demand.[11]

---

[9] Note that L_t/P_t decreases as τ_t rises, if $R^{ss}(1+R^{ss})^{-1}-\theta$>0, where superscript ^{ss} denotes the steady-state value.

[10] If the adjustment cost were too high, upon tighter policy, agents would comply with the policy stance to avoid excessive adjustment costs that outweigh the gains from utilizing the money supply endogeneity. We find through calibrations that the larger is θ, the more quickly β_H turns negative as α increases.

[11] Specifically, suppose that the investor has a production technology that transforms working hours into real output and that leisure is included in his utility function. This extension does not affect the form of equation (8). This line of model extension will also be (continued...)

Second, in contrast to our assumption that the single authority responds to the investor's demand, one may consider the case in which the authority responds to the *average economy-wide* demand but not to an individual investor's demand. In this case, the economy-wide money demand calls for the money supply response in the second half-period. This renders τ_t different from unity and requires the adjustment of money holdings through which the policy stance is fed into the total money holdings. The qualitative implication of equation (8) remains the same although the positive effect of the policy stance on money demand becomes smaller.

III. Empirical Investigations

A. Data and Estimation Methods

The data used are quarterly and monthly U.S. observations (for data sources, see Appendix C). We measure money (M) by M1 to capture how agents adjust highly liquid, short-term monetary assets upon their recognition of policy shocks. The scale variable for money demand (X) is measured by real GNP for quarterly data and by the real personal consumption expenditure for monthly data. The price level (P) is the implicit GNP deflator for quarterly data and the implicit personal consumption expenditure deflator for monthly data. The opportunity cost of money holdings (R) is defined as the market rate minus the own rate of return on M1. The market rate is measured as the three-month Treasury bill rate. The own rate of M1 is calculated using series of individual assets in M1 and their own rates of return.

We construct a proxy for the financial service since its direct measure is not available while the capital stock of depository institutions are available only on a yearly basis. Following Dotsey (1984) and Choi and Oh (1999), a proxy for the quarterly financial service in logarithm is measured as $\ln\sum_{i=0}^{t-1}(1-\delta)^i R^*_{t-i}$, where R^* is the quarterly rate of return on the ten-year Treasury bond (TB) ($R10$), the initial period (t=0) is 1954:2 being dictated by the availability for $R10$, and δ=0.0212. Finally, the policy stance is measured by the (inverted) Boschen-Mills (1995) index ($H1$) or by the change in the Federal funds rate ($H2$), both of which are known as good proxies for the monetary policy stance (Bernanke and Blinder, 1992; Christiano et al., 1996).

Figure 1 depicts the monthly movements of the aforementioned variables. Panels A and B show the real balance, the real output, and the opportunity cost variable, all in logarithms. Panel C displays the measured financial service (f). The logarithm of the capital stock of depository institutions (k) and of the number of automated teller machines (ATMs)

---

compatible with a pricing scheme that results in price rigidity, which is employed by Goodfriend (1997) and Galí (1999).

(*atm*) are also depicted for comparison, both being available on a yearly basis (see Appendix C). Both f and k grew rapidly until the mid-1980s. After that, the former stagnated while the latter grew at a lower speed. The number of ATMs sharply increased before 1985, and since then its growth has slowed down, consistent with other financial service measures. Finally, panel D shows policy stance measures $H1$ and $H2$. A higher value in both measures represents a tighter policy stance.

We begin by testing whether each series in equation (8) possesses a unit root. Test results indicate that m_t–p_t, x_t, r_t, and f_t are integrated of order one (I(1)), but H_t is stationary.[12] A long-run or cointegrating relationship among the I(1) variables must exist to support our model. Also, deviations from the cointegrating relationship should be caused by the stationary policy stance if the money supply is partially endogenous. To examine this effect of the policy stance on money demand, we follow a two-step approach. First, after identifying whether there exists a long-run relationship among the I(1) variables, we estimate the long-run money demand relationship. Next, we examine how monetary policy affects money demand by regressing deviations from the long-run money demand on the policy stance.[13]

Our theoretical model and unit root test results imply the following cointegrating system:

$$(m_t - p_t) - \beta_x x_t - \beta_r r_t - \beta_f f_t = u_t, \tag{9}$$

where u_t is a stochastic error that is stationary. Equation (8) implies that u_t embodies $\beta_H H_t$. We assume for the moment the same degree of money supply endogeneity for the whole sample period. To test the null hypothesis that the variables in equation (9) are not cointegrated, we employ Johansen's (1988, 1991) trace test ($\mathrm{J}^{\mathrm{T}}(1)$) and maximum eigenvalue test ($\mathrm{J}^{\mathrm{Max}}(1)$). These tests, accounting for the simultaneity among variables, are

---

[12] The augmented Dickey-Fuller (ADF) tests with lag lengths up to eight are performed on both the monthly and quarterly data. The unit root hypothesis is not rejected at the 5 percent significance level for each variable except for measures of the policy stance (test results are available from the authors upon request).

[13] As shown by Park and Phillips (1989, p. 102), once the first-step estimation supports cointegration among I(1) variables, equation (8) can be estimated by the standard ordinary least squares method and a statistical inference can be made for β_H. We instead follow the two-step approach because, as noted by Park and Phillips, the least squares estimator of β_H is asymptotically equivalent to the regression coefficient from the second-step regression. Another merit for this approach is that it can be easily adapted into other methods for estimating cointegrating vectors.

based on a vector autoregression (VAR), which include unrestricted intercepts with no trends or unrestricted intercepts with restricted trends.

Although we can obtain a consistent estimator of the cointegrating vector by the standard ordinary least squares (SOLS) method following Engle and Granger (1987), the SOLS estimator may have a severe small sample bias when regressors are correlated with the error term as noted by Stock and Watson (1993). Thus we also employ the Phillips-Hansen (1990) fully modified estimator (PHFM), which has semiparametric corrections for serial correlation and endogeneity to assure asymptotically median-unbiased estimators.[14]

To examine how the policy stance causes deviations from the long-run money demand relationship, we regress the estimated *percentage* deviations from long-run equilibrium values on the intercept and a policy stance measure: [15]

$$\hat{u}_t = c_0 + \beta_H H_t + \xi_t, \tag{10}$$

where $\hat{u}_t$ is the estimated residual of equation (9) multiplied by 100 and H_t, a policy stance measure. Since the error term in equation (8) may involve strong persistence due to the partial adjustment of money holdings, we assume that the error term ξ_t is serially correlated.

The slope parameter, β_H, can be estimated consistently since $\hat{u}_t$ is a consistent estimate of u_t.[16] The use of $\hat{u}_t$ may introduce approximation errors that can be correlated with H_t (as implied by (8)). To deal with this possibility, we use an instrumental variables (IV) procedure to obtain asymptotically correct estimates. We use well-known indicators for the policy stance as instrumental variables. The instrument sets include {intercept, $H1_{t-1}$, Δnbr_t} and {intercept, $\Delta RDFFR_t$, Δnbr_t}, where *nbr* is nonborrowed reserves (NBR) in logarithm, and *RDFFR* is the discount rate before 1974:09 and the target funds rate since

---

[14] It is well known that Johansen's (1991) estimator (JOH), although unbiased, contains poor finite sample properties compared to efficient estimators such as the PHFM and Stock and Watson's (1993) dynamic OLS (DOLS) estimator. The DOLS method applied to our system, while resulting similar point estimates of cointegrating vectors, however, involves a multicollinearity problem because leads and lags of the first-differenced financial service variable are closely correlated with the opportunity cost variable.

[15] We avoid estimating the regression in a dynamic form because a high correlation between differenced values of the opportunity cost variable and policy shock variable causes a multicollinearity problem.

[16] We find that Monte Carlo simulation assures that the two-step procedure with SOLS and PHFM methods captures correctly the effect of an I(0) regressor on an I(1) dependent variable when regressors are exogenous and independent of the model error that follows an AR(1) process.

1974:09.[17] To account for the serial correlation in ξ_t, we compute standard errors of β_H by Newey and West's (1987) method with instrumental variables.

B. Effects of the Policy Stance on Money Demand

First we conduct cointegration tests and the two-step estimation for the whole sample, assuming a single regime. Table 1 reports the results of cointegration tests with quarterly and monthly series. The null hypothesis of no cointegrating vector against more than one cointegrating vector is rejected by the Johansen tests at the 5 percent significance level in both the no trends and restricted trends cases.[18] Also, the choice of data frequency does not affect the results qualitatively.

Table 2 (columns 2–4) contains cointegrating vector estimates. The PHFM estimates are quite similar to the SOLS estimates. Notably, the income elasticity estimate is close to unity, regardless of the choice of data frequency. Specifically, with monthly data, the PHFM and SOLS estimates for $[\beta_x, -\beta_r, -\beta_f]$ are [1.098, –0.054, –0.457] and [1.096, –0.049, –0.456], respectively. The similarity of these estimates suggests that the possibility of a small-sample bias in the SOLS estimate will be small. Figure 2 depicts monthly deviations from the long-run demand based on the PHFM estimator. The deviations show some persistence, presumably reflecting parameter instability (see, for related episodes, Choi and Oh, 1999).

Table 2 (columns 5–8) summarizes the estimated results with two alternative measures of the policy stance. The results for quarterly and monthly series are qualitatively the same. The estimated coefficient of the policy stance, $\hat{\beta}_H$, is positive, but insignificant. The generalized R^2, GR^2, proposed by Pesaran and Smith (1994) as an appropriate measure of fit for IV regressions, is quite small and in the range of 0.00–0.03. Sargan's (1964) misspecification test result suggests that instrumental variables are valid instruments, as indicated by its *p*-values. Thus, we find little evidence of the effect of the money supply factor on money demand for the whole sample.

---

[17] We have chosen instrumental variables to satisfy the following conditions. First, instruments are highly correlated with the regressors. The first stage F statistic is well above 10 (see Staiger and Stock, 1997). Second, instrumental variables are independent of the residual. Although lagged unemployment and inflation rates may affect the policy stance, we do not use them since they tend to be correlated with the estimated residuals.

[18] There is no clear-cut evidence about the number of cointegrating vectors. For the monthly data, the null hypothesis of one cointegrating vector is not rejected. For the quarterly data, the null is not rejected when the intercept and trends are unrestricted but rejected in favor of two cointegrating vectors when no trends or restricted trends are included. The second cointegrating vector may arise due to financial service demand. We employ estimation methods to account for the possible endogeneity of the financial service in a single equation approach.

However, the whole sample analysis cannot capture different degrees of the money supply endogeneity under different monetary regimes. It is well known that the degree of the money supply endogeneity is affected by the Fed's targeting procedure. In particular, interest rate targeting results in the procyclical and/or endogenous behavior of the money supply. We now explicitly take into account shifts in the targeting procedure.

We consider the following regimes: money market condition targeting (1959:02–1974:08) [Regime 0: RG0]; Federal funds rate (FFR) targeting (1974:09–1979:09) [Regime 1: RG1]; nonborrowed reserve (NBR) targeting (1979:10–1982:09) [Regime 2: RG2]; and post-NBR targeting (1982:10–1996:06) [Regime 3: RG3]. In October 1982, the Fed adopted borrowed reserve (BR) targeting. This procedure is close to interest rate targeting since borrowed reserves are affected by the spread between the market rate and discount rate. In addition, the Fed gradually moved back to FFR targeting during the mid-1980s. FFR targeting was quite explicit from 1984:03 (Rudebusch, 1995).[19] Since an explicit FFR targeting may have strengthened the money supply endogeneity, we additionally consider the recent FFR targeting (1984:03–1996:06) [Regime 4: RG4].

To check the varying endogeneity of the money supply, we conduct a simple vector autoregression (VAR) analysis as a diagnostic test. The VAR contains three monthly variables, a policy stance measure, real M1 balance growth, and NBR growth.[20] We presume that the policy stance ($H2$) is contemporaneously exogenous to the other variables and that innovations to the real balance and NBR reflect changes in money demand and money supply, respectively. Figure 3 displays impulse responses (annualized in percentage) under different regimes. The first two rows show the NBR growth response to a positive, one-percentage point shock in the real balance growth with alternative orderings. With the ordering of $\{H2_t, \Delta(m_t - p_t), \Delta nbr_t\}$ allowing for a contemporaneous NBR response to a real balance shock, the NBR growth positively responds to the real balance shock under all regimes, and its impact is notably high under RG3 and RG4. With the ordering of $\{H2_t, \Delta nbr_t, \Delta(m_t - p_t)\}$ not allowing for the contemporaneous NBR response to a real balance shock, the real balance shock is subsequently followed by money supply expansions to a significant extent under RG3 and RG4 only. The last row shows real balance responses to a one-percentage point shock in $H2_t$. The real balance growth increases significantly within the period when tighter policy is implemented under RG3 and RG4. Note that tighter policy later on leads to decreases in the real balance. In sum, the result supports a strong money supply

[19] The transition toward funds rate targeting was completed by 89:12 when the Federal Open Market Committee's borrowing objective became unavailable. There are some variations in the starting date of the recent FFR targeting: e.g., it is 84:03 in Rudebusch (1995), 87:01 in Sellon (1994), and 87:10 in Meulendyke (1998, p. 55), respectively.

[20] The lag length is chosen at 3 months based on the Hannan-Quinn information criterion over the whole sample. Different lag lengths provide qualitatively similar results.

endogeneity and the positive initial impact of tighter policy on money demand, especially under RG3 and RG4.

We note the price puzzle that tighter policy apparently leads to a high price level when a commodity price index (PCOM) is not controlled in a VAR analysis (see Sims, 1992; Christiano et al., 1996). Controlling for the price effect may reinforce the initial impact of tighter policy on the real balance. To deal with the possible influence of the price puzzle, we include inflation and the growth rate of PCOM (ordered fourth and fifth in the orthogonalization, respectively) in the five-variable VAR analysis. This exercise provides qualitatively the same results, while the positive initial impact of tighter policy on real balances is slightly strengthened under RG3-RG4.

We now conduct cointegration tests for different regime periods. Table 3 summarizes the results from Johansen's trace test and maximum eigenvalue test. Although quarterly observations are not enough for some sub-samples, we nevertheless conduct the tests except for RG2 that contains too few observations. The results are insensitive to the choice between the quarterly and monthly data. The test results indicate that the null hypothesis of no cointegration is rejected at the 5 percent significance level in most cases, suggesting the existence of cointegration among the variables under all the regimes.

Cointegrating vectors are estimated under different regimes. The SOLS and PHFM estimators generate very similar estimated residuals, and we report the results based on the PHFM estimator in Tables 4 and 5 (columns 2–4). Both the quarterly and monthly models provide similar implications. The quarterly model for RG2 is not estimated because the sample contains too few observations. The estimated cointegrating vector is rather sensitive to regimes, showing the wrong sign in $\hat{\beta}_r$ under RG1 and smaller values in $\hat{\beta}_x$ under RG0 and RG1. This sensitivity is in line with Lucas' (1988) and Stock and Watson's (1993) findings that the presence of multicollinearity between income and interest rate (particularly in 1973–1982) causes income and interest elasticities to be imprecise. The income elasticity became larger and close to one after the late 1970s. The demand for money became more sensitive to interest rates after 1982.[21] On the other hand, $\hat{\beta}_f$ has become imprecise under RG3 and RG4, plausibly due to flat movements of financial services since the mid-1980s. Figure 4 depicts the residuals from the cointegrating vectors over the regimes. The residuals under RG0–RG2, compared with those in Figure 2, are much reduced. A big swing in the residual under RG3–RG4 reflects the recent velocity puzzle for 1992-1995, which may be partly attributed to shifts from M1 into other assets.[22]

[21] This is possibly due to the fact that rates of return on some components of M1 after financial deregulations were close to market rates but slow to change (Meulendyke, 1998, p. 53).

[22] First, the surge in popularity of bonds and mutual funds around 1992 may have led to shifts from M1 and time and savings deposits into mutual funds (Pakko, 1995). Second, the introduction of sweep accounts in 1995 caused a shift from M1 into time and savings

(continued...)

Tables 4 and 5 (columns 5–8) also report the second-step regression results under different regimes. Again, both the quarterly and monthly models provide similar implications. All regressions pass the Sargan test at the 5 percent significance level, indicating that instrumental variables are reasonably chosen. In this regression, GR^2 under RG3 and RG4 is substantially high: e.g., under RG4, it is 0.381 for the quarterly model and 0.246 for the monthly model when the policy stance is measured by the funds rate change. For both policy stance measures, $\hat{\beta}_H$ is positive for all regimes and highly significant for RG3 and RG4. As an exception, it is also statistically significant with *H*2 under RG2 in the monthly model. Most importantly, $\hat{\beta}_H$ is at its highest under RG4 where the money supply endogeneity is expected to be most substantial: a percentage point rise in the funds rate change contemporaneously increases monthly money demand by 7.3 percent. That is, the money demand response to tighter policy is strongly significant and notably sizable under the recent FFR regime.

Different estimates of $\hat{\beta}_H$ over different regimes presumably reflect the following. The money supply endogeneity under RG3 and RG4 will be much stronger due to the explicit funds rate targeting as noted earlier, although interest rates are emphasized in pursuing monetary policy under RG0. Money supply endogeneity should be low under RG1 and RG2, since a strong emphasis is placed on monetary aggregates as the intermediate target under RG1 despite the funds rate being an operating target and under RG2 with NBR targeting.[23]

Our empirical findings provide a clue for the 'liquidity puzzle' —the rise of interest rates immediately following a positive shock in (broad) money in VARs. That is, tighter monetary policy accompanies the initial rise in the real balance on the one hand and increases the interest rate on the other. With sluggish price responses to monetary shocks, this comovement will show up as a positive contemporaneous correlation between money and interest rates.

Since the liquidity puzzle in the VAR analysis pertains to unanticipated shocks in M1 and interest rates, we further regress the deviation in the real balance, which can be taken as the (induced) innovation in M1, on anticipated and unanticipated components of the funds rate change. The estimated impacts, reported in Appendix D, are statistically highly significant (while the anticipated policy stance has a stronger impact and is extremely

---

deposits because sweep accounts enabled banks to sweep funds automatically out of checkable accounts to avoid the reserve requirement (Gavin, 1996).

[23] Karamouzis and Lombra (1989, p. 53) argue that the NBR procedure represented 'a middle-ground between the perfectly interest elastic short-run supply of reserves under the old funds rate procedure and a completely interest inelastic short-run supply of reserves.' Mishkin (1997, p. 492) argues that the NBR procedure enabled Volcker to use the funds rate to fight inflation. Both imply that the money supply was not fully exogenous.

significant), suggesting that the innovation in M1 comoves with interest rates.[24] This result is also consistent with panel C of Figure 3, i.e., a percentage point positive shock in the funds rate change initially raises the real balance by about four percentage points under RG3–RG4.

C. Robustness Checks

To assure the robustness of our results, we first use various methods to estimate the cointegrating vector for money demand. We have shown that the SOLS and PHFM estimators provide quite similar results with both the quarterly and monthly data. Also, the use of the nonlinear least squares estimator as in Baba et al. (1992) or of the autoregressive distributed lag estimator suggested by Pesaran and Shin (1995) does not change the results qualitatively. They all show that tighter monetary policy affects money demand positively under the recent FFR targeting. This result is robust to the choice of a policy stance measure between the BM index and the funds rate change. Not surprisingly, when the recent FFR targeting period is further confined to the post-1987 period (see footnote 19), the positive effect of policy shocks on money demand becomes more evident. Moreover, the estimation of equation (10) by the OLS method with the Newey-West adjusted standard errors, treating the generated residual as its true value,[25] and the estimation of equation (8) by least squares (see footnote 13) based on Park and Phillips (1989) yield qualitatively the same results.

Additional robustness checks are carried out. First, we examine the conventional system excluding the financial service variable. The income elasticity estimate becomes close to 0.5 for the whole period. Its estimated cointegrating vectors appear severely unstable for the pre-1984 period: the income or interest elasticity appears to be insignificant or to have a wrong sign under RG0-RG2. In contrast, the income and interest elasticities under RG4 are quite similar to those in Tables 4 and 5.[26] Nevertheless, estimating equation (10) based on the conventional money demand function yields qualitatively the same results. Second, a broader view of money demand argues for a long-term rate (see, e.g., Goldfeld and Sichel, 1990) as

[24] The actual policy stance is fed into money demand in the uncertainty case, as well. Thus, the unanticipated component of the policy stance also affects money demand.

[25] The cointegrating vector is consistently estimated at a convergence speed higher than the coefficient on the I(0) variable of the second-step regression. Thus, the approximation error introduced into equation (10) by the generated residuals would be of a negligible order.

[26] This evidence suggests that the financial service maintained a substantial explanatory power for money demand until the mid-1980s but it became ineffective as its movements became flat. Hansen's (1992) parameter stability test for the cointegrating vector suggests that our money demand relation outperforms the conventional one for the pre-1984 sample whereas both involve parameter instability for the post-1984 sample.

the opportunity cost, although much of the literature based on a transactions view uses a short-term rate. The use of the ten-year TB rate in measuring the opportunity cost provides qualitatively similar results. Third, alternative measures of the policy stance such as the spread between the funds rate and the ten-year TB rate (e.g., Bernanke and Blinder, 1992) and the mix of nonborrowed and total reserves suggested by Strongin (1995) provide results that are fairly supportive of our model. Fourth, it could be of interest to use the capital stock of depository institutions in measuring the financial service. Using the interpolated series of the capital stock, we find that money demand significantly increases in response to tighter policy under RG3 and RG4 but not under other regimes.[27] Finally, there may be a trend component in the money demand function (as implied by equation (A6) in Appendix B). That trend, however, may share trend components in x_t and/or f_t and reflect some unspecified trend in m_t-p_t under each regime. Nonetheless, we check whether the inclusion of a time trend in equation (9) changes the result. We obtain similar results that money demand responds positively to the policy stance under RG3 and RG4, while the trend varies over regimes.

Taken together, our results, robust to alternative estimation methods, data frequencies, and alternative measures of variables, provide evidence that tighter monetary policy generates strong positive impacts on money demand under the recent FFR targeting.

IV. Concluding Remarks

Despite the evidence that the money supply is at least partially endogenous, not much research in the money demand literature has focused upon its endogeneity. Moreover, no previous studies have explicitly accounted for the effect of the partial endogeneity of monetary policy on money demand. This paper develops a theoretical model that incorporates both the partial endogeneity of monetary policy and endogenous financial innovations and empirically investigates the money demand function derived from the model.

In a general equilibrium framework, we attribute the utility gains from money to the liquidity service provided by the real balance and financial service acquired before transactions. The partial endogeneity of policy is incorporated by the money supply rule that potentially reconciles the notion of a Taylor rule. Money holdings are chosen sequentially by half-periods with an adjustment cost for changing money balances in the second half-period. The model analysis suggests that money supply factors affect money demand due to money supply endogeneity and that money demand parameters are regime-dependent.

[27] The estimated cointegrating vectors by PHFM and DOLS suggest that the income and interest elasticities are similar to those reported in Tables 2, 4, and 5 for the whole sample and RG3 and RG4 but not for RG0-RG2.

The estimation results with U.S. data indicate that tighter policy has a positive initial impact on money demand, which is the most pronounced under the recent FFR targeting. This evidence reflects that an increase in money demand induces an accommodating increase in the money supply, which countervails or mitigates the potential impact of a policy tightening on real balances. Our finding is in line with the evidence that firms' demands for net funds initially rise upon tighter policy (e.g., Christiano et al., 1996). We argue that money supply endogeneity motivates people to hold money with a preemptive motive to cushion the impact of policy shocks. Consequently, policy shocks are fed into money demand perturbations. This reflection of policy shocks in money demand helps explain the liquidity puzzle.

Our regressions assume that the policy stance causes money demand. The reverse causation that a tighter policy reflects the Fed's response to a positive money demand shift, however, is not consistent with interest rate targeting and hence the endogeneity of money supply. For the policy stance, especially, a change in the funds rate should not take place in respond to a money demand shift under interest rate targeting. Rather the money supply must increase to accommodate such a shift.

Finally, our finding suggests that financial innovations and turbulence in monetary policy deteriorate the stability of the conventional demand function. Given the importance of money supply endogeneity, it is of crucial importance to take into these factors when we assess the effects of monetary policy. For this purpose, money supply endogeneity can be examined in a sophisticated VAR analysis that restricts separately money demand and money supply shocks (e.g., Gordon and Leeper, 1994), which is left for future research.

Table 1. Money Demand Cointegration Tests, 1959:02–1996:06

Test	Quarterly	Monthly
	Intercepts / Trends	
$J^T(1)$	73.56* / 82.95* (48.88) / (63.00)	63.17* / 71.49* (48.88) / (63.00)
$J^{Max}(1)$	33.86* / 36.28* (27.42) / (31.79)	33.09* / 33.27* (27.42) / (31.79)

Notes: (i) $J^T(1)$ and $J^{Max}(1)$ are Johansen's trace test and maximum eigenvalue test, respectively, of no cointegration against more than one cointegrating vector. (ii) The VAR with the order of 4 includes unrestricted intercepts and no trends or alternatively unrestricted intercepts and restricted trends. Critical values at the 5% significance level (Osterwald–Lenum, 1992; Pesaran et al., 1996) are in parentheses. * significant at .05 level.

Table 2. Cointegrating Vector and Money Supply Factor Coefficients

Estimator	Estimated cointegrating vectors			Estimated effect of money supply factor			
	β_x	$-\beta_r$	$-\beta_f$	β_{H1}	β_{H2}	GR^2	χ^2
Quarterly							
SOLS	1.307	−0.082	−0.512	0.296	·	0.004	0.94
	·	·	·	(0.628)	·		[V1]
				·	0.201	0.002	0.88
				·	(0.445)		[V2]
PHFM	1.278**	−0.095**	−0.492**	0.794	·	0.027	0.75
	(0.099)	(0.014)	(0.061)	(0.617)	·		[V1]
				·	0.404	0.008	0.97
				·	(0.434)		[V2]
Monthly							
SOLS	1.096	−0.049	−0.456	0.506	·	0.014	0.87
	·	·	·	(0.378)	·		[V1]
				·	0.663	0.006	0.76
				·	(0.584)		[V2]
PHFM	1.098**	−0.054**	−0.457**	0.643^{+}	·	0.022	0.95
	(0.055)	(0.008)	(0.037)	(0.377)	·		[V1]
				·	0.717	0.007	0.77
				·	(0.561)		[V2]

Notes: (i) All regressions include a constant. Since the asymptotic distribution of the SOLS estimator involves the unit root distribution and is nonstandard (Campbell and Perron, 1991), its standard errors are not reported. For the PHFM estimator, the frequency zero spectral estimator allowing for trends was computed using a Parzen kernel with 5 lags. The sample period for the second-step regressions with *H*1 ends at 1995:03 due to the BM index availability (updated from Boschen). (ii) In IV regressions, standard errors in parentheses are corrected by the Newey-West (1987) method with Parzen weights (truncation lags=5). GR^2 is Pesaran and Smith's (1994) generalized R^2. (iii) The column for χ^2 reports *p*-values of Sargan's (1964) test of misspecification that follows $\chi^2(n\text{-}2)$ under the null hypothesis of correct specification with *n* valid instruments. Each regression uses an instrument set indicated in a square bracket: V1={a, $H1_{t-1}$, Δnbr_t}, V2={a, $\Delta RDTFFR_t$, Δnbr_t}, where *a* denotes an intercept. $^{+}$significant at .10 level. *significant at .05 level. **significant at .01 level.

Table 3. Different Monetary Regimes and Money Demand Cointegration

Regime [Sample period]	Test	Quarterly Intercepts / Trends	Monthly Intercepts / Trends
RG0	$J^T(1)$	142.6* / 171.8*	50.81* / 66.29*
		(48.88) / (63.00)	(48.88) / (63.00)
[59:02–74:08]	$J^{Max}(1)$	106.6* / 128.4*	27.67* / 42.71*
		(27.42) / (31.79)	(27.42) / (31.79)
RG1	$J^T(1)$	70.66* / 79.95*	73.45* / 100.5*
		(48.88) / (63.00)	(48.88) / (63.00)
[74:09–79:09]	$J^{Max}(1)$	32.47* / 32.55*	44.78* / 44.89*
		(27.42) / (31.79)	(27.42) / (31.79)
RG2	$J^T(1)$	–	90.17* / 110.2*
			(48.88) / (63.00)
[79:10–82:09]	$J^{Max}(1)$	–	59.97* / 65.27*
			(27.42) / (31.79)
RG3	$J^T(1)$	57.72* / 75.84*	72.64* / 96.43*
		(48.88) / (63.00)	(48.88) / (63.00)
[82:10–96:06]	$J^{Max}(1)$	29.25* / 29.57$^+$	38.34* / 51.23*
		(27.42) / (31.79)	(27.42) / (31.79)
RG4	$J^T(1)$	49.38* / 68.73*	68.08* / 92.08*
		(48.88) / (63.00)	(48.88) / (63.00)
[84:03–96:06]	$J^{Max}(1)$	31.07* / 36.70*	32.88* / 48.18*
		(27.42) / (31.79)	(27.42) / (31.79)

Notes: (i) $J^T(1)$ and $J^{Max}(1)$ are Johansen's trace test and maximum eigenvalue test, respectively, of no cointegration against more than one cointegrating vector. (ii) The quarterly model is based on the VAR with the order of 2. The monthly model is based on the VAR with order of 6 for RG0 and with order of 4 for RG1-RG4. The VAR includes unrestricted intercepts and no trends or alternatively unrestricted intercepts and restricted trends. Critical values at the 5% significance level (Osterwald–Lenum, 1992; Pesaran et al., 1996) are in parentheses. $^+$significant at .10 level. *significant at .05 level.

Table 4. Cointegrating Vector, Money Supply Factor Coefficient
Under Different Regimes: Quarterly data

Regime	Estimating cointegrating vectors			Estimated effect of money supply factor			
[Sample period]	β_x	$-\beta_r$	$-\beta_f$	β_{H1}	β_{H1}	GR^2	χ^2
RG0	0.426**	−0.001	−0.053	0.196	·	0.016	0.15
[59:1−74:3]	(0.072)	(0.005)	(0.035)	(0.197)	·		[V3]
				·	0.295	0.035	0.31
				·	(0.187)		[V3]
RG1	0.608**	0.005	−0.531**	0.010	·	0.000	0.98
[74:4−79:3]	(0.088)	(0.003)	(0.087)	(0.163)	·		[V1]
				·	0.017	0.000	0.20
				·	(0.098)		[V3]
RG3	1.140**	−0.085**	−0.178	3.100**	·	0.318	0.53
[82:4−96:2	(0.071)	(0.016)	(0.217)	(0.747)	·		[V4]
(95:1)]				·	3.444**	0.311	0.18
				·	(1.100)		[V2]
RG4	1.163**	−0.080**	−0.071	3.496**	·	0.402	0.41
[84:2−96:2	(0.080)	(0.017)	(0.358)	(0.763)	·		[V4]
(95:1)]				·	4.208**	0.381	0.11
				·	(1.276)		[V4]

Notes: (i) Results are based on the PHFM estimator using a Parzen kernel with 5 lags. The sample period for the second-step regressions with $H1$ for RG3–RG4 ends at 1995:1 due to the BM index availability. (ii) In IV regressions, standard errors in parentheses are corrected by the Newey-West (1987) method with Parzen weights (truncation lags=5). GR^2 is Pesaran and Smith's (1994) generalized R^2. (iii) The column for χ^2 reports p-values of Sargan's (1964) test statistic that follows $\chi^2(n\text{-}2)$ under the null hypothesis of correct specification with n valid instruments. The instrument set is indicated in a square bracket: V1=$\{a, H1_{t-1}, \Delta nbr_t\}$, V2=$\{a, \Delta RDTFFR_t, \Delta nbr_t\}$, V3=$\{a, H2_{t-1}, \Delta RDTFFR_t\}$, and V4=$\{a, \Delta RD_t, \Delta TFFR_t\}$, where a denotes an intercept. $^{+}$significant at .10 level. *significant at .05 level. **significant at .01 level

Table 5. Cointegrating Vector, Money Supply Factor Coefficient
Under Different Regimes: Monthly Data

Regime	Estimated cointegrating vectors			Estimated effect of money supply factor			
[Sample period]	β_x	$-\beta_r$	$-\beta_f$	β_{H1}	β_{H1}	GR^2	χ^2
RG0	0.693**	−0.007*	−0.171**	0.021	·	0.000	0.82
[59:02–74:08]	(0.045)	(0.003)	(0.024)	(0.125)	·		[V1]
				·	0.655	0.028	0.08
				·	(0.425)		[V2]
RG1	0.572**	0.015**	−0.538**	0.060	·	0.017	0.21
[74:09–79:09]	(0.084)	(0.002)	(0.083)	(0.071)	·		[V1]
				·	0.218	0.040	0.21
				·	(0.132)		[V4]
RG2	1.064**	−0.008	−0.227**	1.168	·	0.073	0.12
[79:10–82:09]	(0.220)	(0.006)	(0.031)	(1.155)	·		[V4]
				·	0.165**	0.082	0.54
				·	(0.059)		[V4]
RG3	1.016**	−0.063**	−0.082	3.223**	·	0.242	0.29
[82:10–96:06	(0.047)	(0.011)	(0.154)	(0.527)	·		[V4]
(95:03)]				·	7.097**	0.232	0.46
				·	(1.493)		[V2]
RG4	0.953**	−0.068**	−0.387	3.346**	·	0.255	0.28
[84:03–96:06	(0.061)	(0.012)	(0.247)	(0.606)	·		[V4]
(95:03)]				·	7.341**	0.246	0.34
				·	(1.510)		[V2]

Notes: (i) Results are based on the PHFM estimator using a Parzen kernel with 5 lags. The second-step regressions with $H1$ for RG3–RG4 ends at 1995:1 due to the BM index availability. (ii) In IV regressions, standard errors in parentheses are corrected by the Newey-West (1987) method with Parzen weights (truncation lags=5). GR^2 is Pesaran and Smith's (1994) generalized R^2. (iii) The column for χ^2 reports p-values of Sargan's (1964) test statistic that follows $\chi^2(n\text{-}2)$ under the null hypothesis of correct specification with n valid instruments. The instrument set is indicated in a square bracket: V1={a, $H1_{t-1}$, Δnbr_t}, V2={a, $\Delta RDTFFR_t$, Δnbr_t}, and V4={a, ΔRD_t, $\Delta TFFR_t$}, where a denotes an intercept. $^+$significant at .10 level. *significant at .05 level. **significant at .01 level.

Figure 1. Variables in Money Demand Relationships

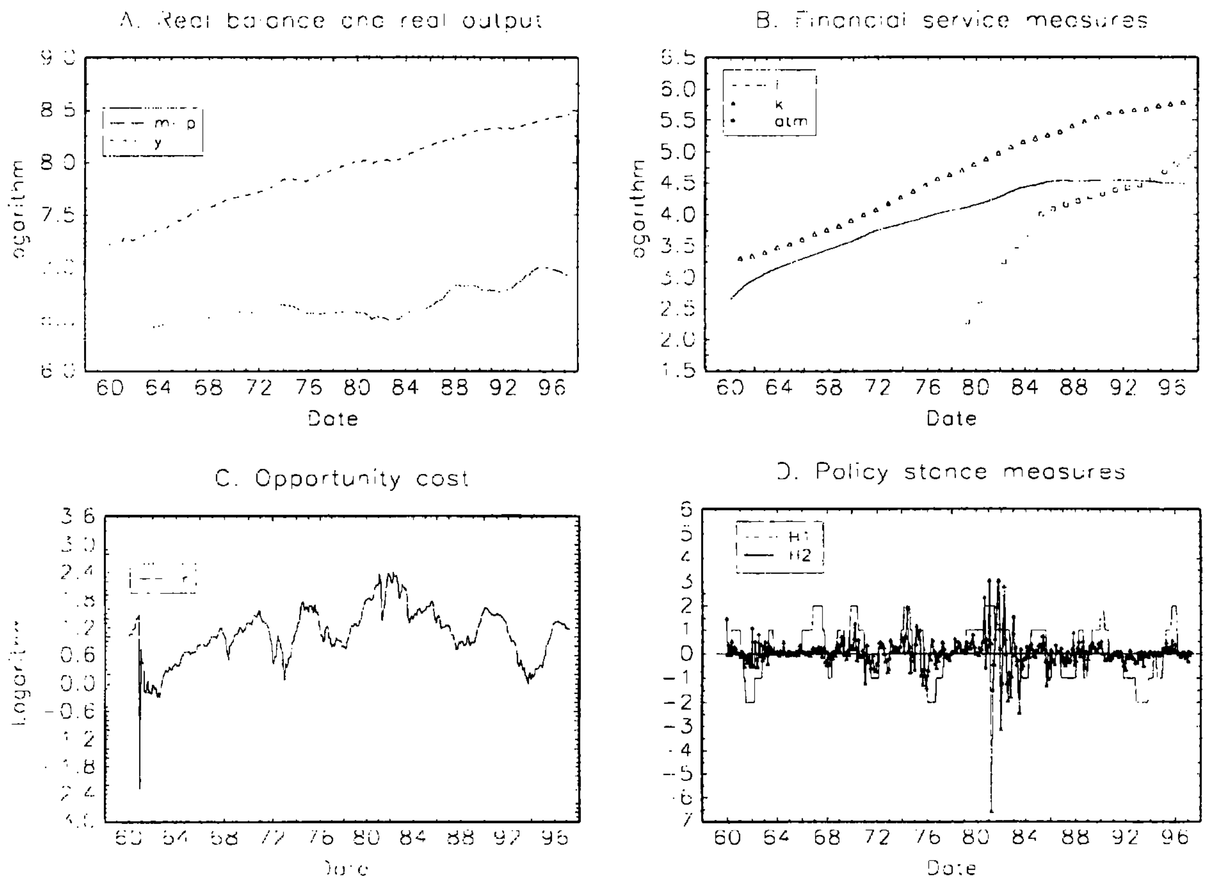


Figure 2. Residual ($\hat{u}_2 \times 10^2$) from the whole Sample Cointegrating Regression

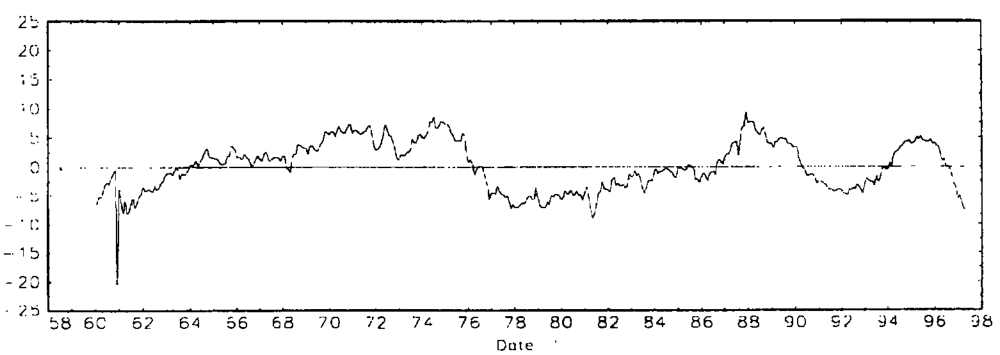
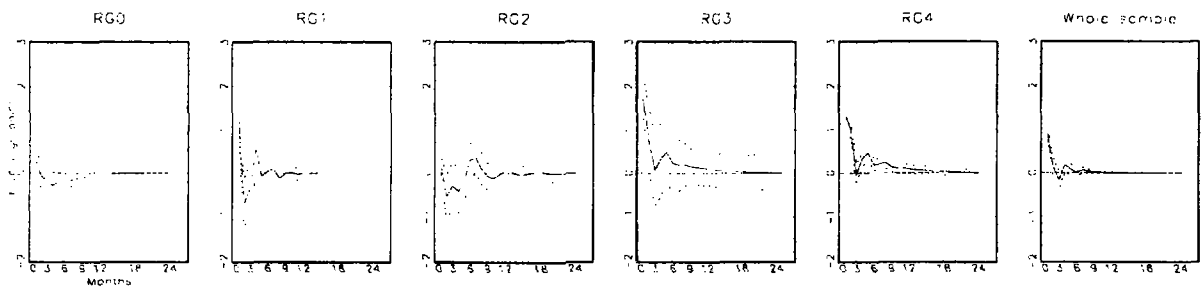


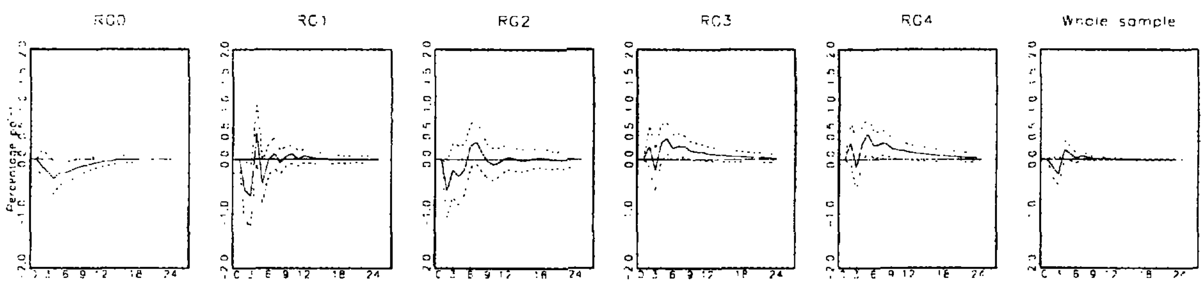
Figure 3. Impulse Responses Under Different Regimes.

The dashed lines are one-standard-error bands computed from the bootstrapping with 1000 replications

A. Nonborrowed Reserves Responses to a Shock to the Real Balance: $\{H2_t, \Delta(m_t-p_t), \Delta nbr_t\}$



B. Nonborrowed Reserves Responses to a Shock to the Real Balance: $\{H2_t, \Delta nbr_t, \Delta(m_t-p_t)\}$



C. Real Balance Responses to a Shock to the Policy Stance: $\{H2_t, \Delta(m_t-p_t), \Delta nbr_t\}$

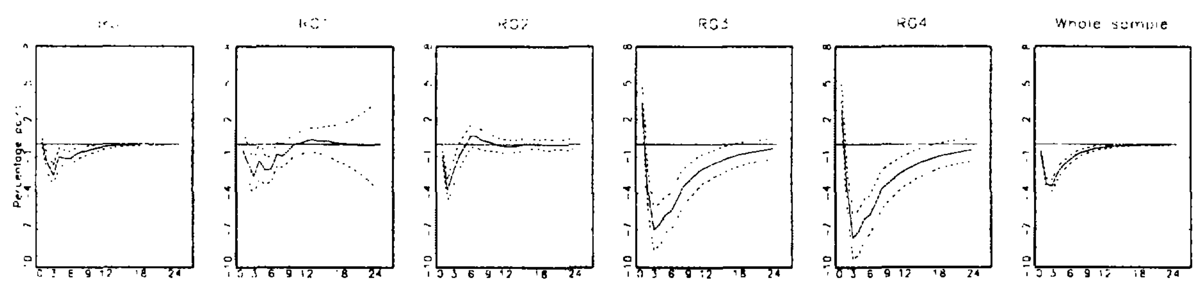
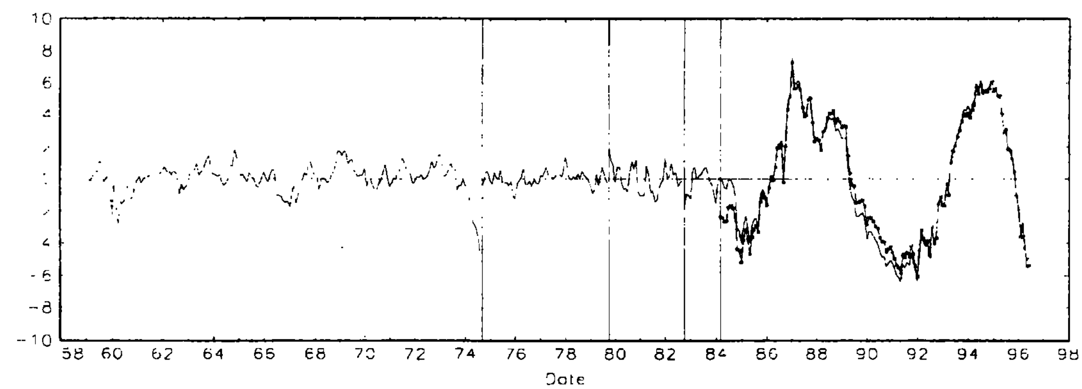


Figure 4. Residuals ($\hat{u}_2 \times 10^2$) from Cointegrating Regressions Under Different Regimes.

Vertical reference lines indicate the end (or start) of regimes. The line with symbols is the residual under RG4.



References

Baba, Yoshihisa, David F. Hendry and Ross M. Starr, 1992, "The Demand for M1 in the U.S.A.," *Review of Economic Studies* 59 (January), pp. 25–61.

Bernanke, Ben and Alan Blinder, 1992, "The Federal Funds Rate and the Channels of Monetary Transmission," *American Economic Review* 82 (September), pp. 901–921.

Boschen, John F. and Leonard O. Mills, 1995, "The Relationship between Narrative and Money Market Indicators of Monetary Policy," *Economic Inquiry* 33 (January, pp. 24–44.

Boyle, Glenn W., 1990, "Money Demand and the Stock Market in a General Equilibrium Model with Variable Velocity," *Journal of Political Economy* 98 (October), pp. 1039–53.

Campbell, John Y. and Pierre Perron, 1991, "Pitfalls and Opportunities: What Macroeconomists Should Know about Unit Roots," *National Bureau of Economic Research Macroeconomics Annual* 6, pp. 141–201.

Choi, Woon Gyu, 1999, "Estimating the Discount Rate Policy Reaction Function of the Monetary Authority," *Journal of Applied Econometrics* 14 (July–August), pp. 379–401.

——— and Yungsan Kim, 2000, "Monetary Policy and Corporate Liquid Asset Demand," Working paper, Hong Kong University of Science and Technology (March).

——— and Seonghwan Oh, 1999, "A Money Demand Function with Output Uncertainty, Monetary Uncertainty, and Financial Innovations," Working paper, Hong Kong University of Science and Technology (August).

Christiano, Lawrence J., Martin Eichenbaum and Charles Evans, 1996, "The Effects of Monetary Policy Shocks: Evidence from the Flow of Funds," *Review of Economics and Statistics* 78 (February), pp. 16–34.

Cogley, Timothy, 1993, "Adapting Instability in Money Demand: Forecasting Money Growth with a Time-Varying Parameter Model," Federal Reserve Bank of San Francisco *Economic Review* (No. 3), pp. 35–41.

Danthine, Jean-Pierre and John B. Donaldson, 1986, "Inflation and Asset Prices in an Exchange Economy," *Econometrica* 54 (May), pp. 585–605.

Dotsey, Michael, 1984, "An Investigation of Cash Management Practices and Their Effects on the Demand for Money," Federal Reserve Bank of Richmond *Economic Review* 70 (September–October), pp. 3–12.

Engle, Robert F. and Clive W. J. Granger, 1987, "Cointegration and Error Correction: Representation, Estimation, and Testing," *Econometrica* 55 (March), pp. 251–276.

Feenstra, Robert C., 1996, "Functional Equivalence between Liquidity Costs and the Utility of Money," *Journal of Monetary Economics* 17 (March), pp. 271–291.

Galí, Jordi, 1999, "Technology, Employment, and the Business Cycle: Do Technology Shocks Explain Aggregate Fluctuations?" *American Economic Review* 89 (March), pp. 249–271.

Gavin, William T., 1996, "The FOMC in 1995: A Step Closer to Inflation Targeting," Federal Reserve Bank of St. Louis *Review* (September–October), pp. 29–47.

Goldfeld, Stephen M. and Daniel E. Sichel, 1990, "The Demand for Money," in *Handbook of Monetary Economics* ed. by Benjamin M. Friedman and Frank H. Hahn, Vol. 1, North–Holland, Amsterdam, pp. 299–356.

Goodfriend, Marbin, 1997, "A Framework for the Analysis of Moderate Inflations," *Journal of Monetary Economics* 39 (June), pp. 45–65.

Gordon, David B. and Eric M. Leeper, 1994, "The Dynamic Impacts of Monetary Policy: An Exercise in Tentative Identification," *Journal of Political Economy* 102 (December), pp. 1228–47.

Hansen, Bruce, 1992, "Tests for Parameter Instability in Regressions with I(1) Processes," *Journal of Business & Economic Statistics* 10 (July), pp. 321–335.

Ireland, Peter N., 1995, "Endogenous Financial Innovation and the Demand for Money," *Journal of Money, Credit, and Banking* 27 (February), pp. 107–123.

Johansen, Sϕren, 1988, "Statistical Analysis of Cointegrated Vectors," *Journal of Economic Dynamics and Control* 12 (June–September), pp. 231–254.

———, 1991, "Estimation and Hypothesis Testing of Cointegrating Vectors in Gaussian Vector Autoregressive Models," *Econometrica* 59 (November), pp. 1551–80.

Judd, John P. and Glenn D. Taylor, 1998, "Taylor's Rule and the Fed: 1970-1997," Federal Reserve Bank of San Francisco *Economic Review* (3), pp. 2-16.

Karamouzis, Nicholas and Raymond E. Lombra, 1989, "Federal Reserve Monetary Policy: An Overview and Analysis of the Policy Process," *Carnegie-Rochester Conference Series on Public Policy* 30 (Spring), pp. 7–74.

Laidler, David E. W., 1993, *The Demand for Money: Theories, Evidence & Problems*, Fourth ed., Harper Collins College Publishers, New York.

Lucas, Robert E., Jr., 1988, "Money Demand in the United States: A Quantitative Review," *Carnegie-Rochester Conference Series on Public Policy* 29 (Autumn), pp. 137–168.

———, 2000, "Inflation and Welfare," *Econometrica* 68, (March), pp. 247–274.

Meulendyke, Ann-Marie, 1998, *U. S. Monetary Policy & Financial Markets*, Federal Reserve Bank of New York, New York.

Mishkin, Fredric S., 1997, *The Economics of Money, Banking, and Financial Markets*. Fifth ed., Addison Wesley Longman Inc.

Newey, Whitney K. and Kenneth D. West, 1987, "A Simple, Positive Semi-Definite, Heteroskedasticity and Autocorrelation Consistent Covariance Matrix," *Econometrica* 55 (March), pp. 703–08.

Osterwald-Lenum, M., 1992, "A Note with Quantiles of the Asymptotic Distribution of the Maximum Likelihood Cointegration Rank Test Statistics," *Oxford Bulletin of Economics and Statistics* 54 (August), pp. 461–471.

Park, Joon Y. and Peter C. B. Phillips, 1989, "Statistical Inference in Regressions with Integrated Processes: Part 2," *Econometric Theory* 5 (April), pp. 95–131.

Pakko, Michael R., 1995, "The FOMC in 1993 and 1994: Monetary Policy in Transition," Federal Reserve Bank of St. Louis *Review* (March–April), pp. 3–25.

Pesaran, M. Hashem and Richard J. Smith, 1994, "A Generalized $\overline{R}^2$ Criterion for Regression Models Estimated by the Instrumental Variables Method," *Econometrica* 62 (May), pp. 705–710.

——— and Yongcheol Shin, 1995, "An Autoregressive Distributed Lag Modelling Approach to Cointegration Analysis," DAE Working Paper No 9514, University of Cambridge.

———, Yongcheol Shin, and Richard J. Smith, 1996, "Testing for the Existence of a Long-Run Relationship," DAE Working Paper No. 9622, University of Cambridge.

Phillips, Peter C. B. and Bruce E. Hansen, 1990, "Statistical Inference in Instrumental Variables Regression with I(1) Processes," *Review of Economic Studies* 57 (January), pp. 99–125.

Poole, William, 1970, "Optimal Choice of Monetary Policy Instruments in a Simple Stochastic Macro Model," *Quarterly Journal of Economics* 84 (May), pp. 197–216.

Ramey, Valerie, 1993, "How Important is the Credit Channel in the Transmission of Monetary Policy?" *Carnegie-Rochester Conference Series on Public Policy* 39 (December), pp. 1–45.

Rudebusch, Glenn D., 1995, "Federal Reserve Interest Rate Targeting, Rational Expectations, and the Term Structure," *Journal of Monetary Economics* 35 (April), pp. 245–274.

Sargan, John. D., 1964, "Wages and Prices in the United Kingdom: A Study in Econometric Methodology," in *Econometric Analysis for National Economic Planning* ed. by P. E. Hart, G. Mills, and J. K. Whitaker, Butterworth, London.

Sellon, Gordon H., 1994, "Measuring Monetary Policy," Research Working Paper, RWP 94–12, Federal Reserve Bank of Kansas City.

Sims, Christoper A., 1992, "Interpreting the Macroeconomic Time Series Facts: The Effects of Monetary Policy," *European Economic Review* 36 (June), pp. 975–1000.

Staiger Douglas and James H. Stock, 1997, "Instrumental Variables Regression with Weak Instruments," *Econometrica* 65 (May), 557–586.

Stock, James H. and Mark W. Watson, 1993, "A Simple Estimator of Cointegrating Vectors in Higher Order Integrated Systems," *Econometrica* 61 (July), pp. 783–820.

Strongin, Steven, 1995, "The Identification of Monetary Policy Disturbances: Explaining the Liquidity Puzzle," *Journal of Monetary Economics* 35 (August), pp. 463–497.

Taylor, John B., 1993, "Discretion versus Policy Rules in Practice," *Carnegie-Rochester Conference Series on Public Policy* 39 (December), pp. 195–214.

A. Equilibrium Conditions of the Model

The first-order conditions for the investor with respect to $\{L_t, z_t, d_t, K_t\}$ together with the market-clearing conditions yield the following equilibrium conditions:

$$E_{t-1}[U_x(X_t,O_t)\{\alpha\tau_t + \Gamma(\tau_t)\}/P_t]$$
$$= E_{t-1}[U_2(X_t,O_t)/P_t] + \beta\alpha E_{t-1}[U_x(X_{t+1},O_{t+1})\frac{1}{\pi_{t+1}}\tau_t/P_t], \tag{A1}$$

$$E_{t-1}[U_x(X_t,O_t)]q_t = \beta E_{t-1}[U_x(X_{t+1},O_{t+1})(Y_{t+1}+q_{t+1})], \tag{A2}$$

$$E_{t-1}[U_x(X_t,O_t)] = \beta E_{t-1}[U_x(X_{t+1},O_{t+1})\frac{1}{\pi_{t+1}}](1+R_t), \tag{A3}$$

$$U_x(X_t,O_t) = \beta E_t[U_x(X_{t+1},O_{t+1})(1-\delta) + U_2(X_{t+1},O_{t+1})\psi n K_t^{\psi-1}], \tag{A4}$$

$$M_t/P_t = T_t^Y e^{-(\lambda/\alpha)H_t - \alpha w_t}\tau_t^{-(1-\alpha)/\alpha}, \tag{A5}$$

where X_t and O_t denote the real consumption ($X_t = Y_t - I_t$) and the real effective liquidity ($O_t = L_t/P_t + nF_t$) in equilibrium, respectively, and $\Gamma(\tau_t) \equiv (\partial J_t P_t/\partial L_t)$. $\Gamma(\tau_t) = \theta[(1/2-\alpha)\tau_t^2 - (1-\alpha)\tau_t + 1/2]$ is obtained using $\partial M_t/\partial L_t = (1-\alpha)M_t/L_t$ from equation (1).

Condition (A1) says that the foregone benefit from an additional L_t equals the expected marginal utility of L_t plus the discounted, expected marginal utility from consumption in period t+1. The foregone benefits are the marginal adjustment cost of L_t and the marginal utility gain from the resulting total money balance. Note that the latter is discounted by the factor α. Due to the contemporaneous money supply on demand, an additional unit in L_t induces a transfer income increase through the money supply increase of $(1-\alpha)\tau_t$. Condition (A2) is the standard capital-asset-pricing equation for the value of equity claim to dividends from sales of output. Condition (A3) reflects the Fisherian decomposition of the nominal interest rate into a real rate (intertemporal marginal rate of substitution) and an expected inflation rate. Condition (A4) says that the foregone benefit from an additional financial capital in period t equals the discounted expected utility gain from the additional financial capital net of depreciation. Finally, condition (A5) reflects the money supply rule and money market clearing condition.

B. Derivation of Equation (8)

Taking the log-linear approximation of equation (7) around the steady-state values of variables,

$$m_t - p_t = \beta_0 + \beta_x x_t - \beta_r r_t - \beta_f f_t - \beta_\tau \ln\tau_t + \varepsilon_t, \tag{A6}$$

where $\beta_0 = -\ln \alpha$, $\beta_x = [(M/P + nF)/(M/P)]^{ss}$, $\beta_r = \beta_x[1/(1+R)]^{ss}$, $\beta_f = [nF/(M/P)]^{ss}$, and $\beta_\tau = (\beta_f - \beta_x \alpha\theta\phi^{-1})$ with $\phi = \{(1-s)/s\} \cdot [X/(M/P + nF)]^{ss}$. Superscript ^{ss} denotes the steady state, and $\tau^{ss} = 1$ is used. We now replace $\ln \tau_t$ with observables. Condition (A5) implies that $\ln \tau_t = \frac{\alpha}{1-\alpha}[\ln T_t^Y - (m_t - p_t) - w_t] - \frac{\lambda}{1-\alpha} H_t$. For T_t^Y, we assume that $\ln T_t^Y - (m_t - p_t) - w_t = \eta_t$, where η_t is a stationary error. Then equation (A6) can be written as equation (8), where $\beta_H = \beta_\tau \cdot \lambda / (1-\alpha)$ and $e_t = \varepsilon_t - \{\beta_\tau \alpha / (1-\alpha)\}\eta_t$. Alternatively, allowing T_t^Y to share a trend with the consumption, if merge T_t^Y into x_t, we obtain a money demand function similar to equation (8) although all coefficients depend on α.

C. Data Source and Variable Definitions

All U.S. data except for the Federal funds rate target and the annual capital stock are from FRED at the Federal Reserve Bank of St. Louis Web site. FRED code names and variable definitions follow: GNP (quarterly nominal GNP); GNPC92 (quarterly real GNP, chained 1992); PCE (nominal personal consumption expenditure); PCEC92 (real personal consumption expenditure, chained 1992); M1SL (money stock M1); TB3MS (three-month Treasury bill rate); GS10 (ten-year TB rate); BOGNONBR (nonborrowed reserves); TRARR (total reserves); FEDFUNDS (Federal funds rate, daily average); MDISCRT (discount rate, daily average); $RDFFR_t = RD_t \cdot I_{(t<1974:09)} + TFFR_t \cdot I_{(t\ 1974:09)}$, where RD is the discount rate, $I_{()}$ is an indicator function, and $TFFR$ is an average of the daily funds rate target compiled from Sellon (1994) and Rudebusch (1995). We constructed the explicit own rate of M1 by weighting explicit interest rates on all individual components of M1 by the lagged ratio of the individual components to M1 (compiled from FRED). All the above series are monthly data except for GNP and GNPC92. The monthly data are averaged to obtain the quarterly observations. The annual current (net) capital stock of depository institutions is taken from the 'Tangible Wealth Table 5KCU' (K1NFI601ES00) of the September 1998 *Survey of Current Business*. The capital stock (bil. \$) is deflated by the last quarter GNPC92 and then rendered in the quarterly series through interpolation, imposing a constant growth of the capital stock within the same year. The number of ATMs (annual figures, end of June) starting from 1978 is taken from *EFT Network Bank Data Book* (1988, 1999) (complied from *Bank Network News*).

D. Effects of Anticipated and Unanticipated Policy Stance on Money Demand

Anticipated and unanticipated values of $H2_t$ are generated from a monthly AR(6) model for each of RG3 and RG4. The estimated *percentage* deviation from equation (10), $\hat{u}_t$, is regressed against the anticipated policy stance, $H2_t^{AN}$, and the unanticipated policy stance, $H2_t^{UN}$ (see Table 6).

Table 6. Anticipated and Unanticipated Policy Stance Coefficients

Regime	β_H^{AN}	β_H^{UN}	GR^2	$\chi^2(1)$
RG3 [82:10–96:06]	12.949** (2.636)	4.070** (1.125)	0.25	0.08
RG4 [84:03–96:06]	13.841** (2.768)	4.250** (1.183)	0.27	0.06

Notes: (i) Entries are the coefficients obtained by regressing $\hat{u}_t$ (based on the PHFM estimator) against $H2_t^{AN}$ and $H2_t^{UN}$, using {*intercept*, $\Delta RDTFFR_t$, Δnbr_t, $H2_{t-1}$} as instruments. (ii) Standard errors in parentheses are corrected by the Newey-West (1987) method with Parzen weights (truncation lags=5). (iii) GR^2 is the generalized R^2. The column for χ^2 (1) reports *p*-values of Sargan's (1964) test. $^+$significant at .10 level. *significant at .05 level. **significant at .01 level.