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FINANCIAL INNOVATION AND THE CONTROL OF MONETARY AGGREGATES: SOME EVIDENCE FROM CANADA

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ABSTRACT

This paper presents an empirical test of the proposition that control of a monetary aggregate will generate a rise in its velocity. The test is carried out utilizing the Canadian experience of controlling M1 growth from 1975:3 to 1982:3. Section One of the paper presents evidence of the instability of the Canadian demand from M1 money since 1975:3. Section Two develops a specific form of the proposition which emphasizes the role of asset substitution between classes of chartered bank deposits. A relative asset demand equation is derived from a wealth maximization model subject to a technological transactions constraint and this equation is estimated from 1961 through 1982. The results lend support to the proposition that central bank control of M1 generated a rise in M1 velocity.

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Introduction

Over the past nine years there has been extensive documentation of instability in the U.S. M1 demand for money function. Since this instability implies less predictability between money and nominal income, it raises serious doubts about the ability of the monetary authorities to control nominal income through a policy of targeting the growth rates of monetary aggregates.

The evidence on the instability of the money demand function has been reviewed recently by Judd and Scadding [1982] and they note that it has spawned two distinct research agendas. First, it is argued that financial innovation has changed the meaning of the monetary aggregates which the authorities are attempting to control. The solutions implied by this argument are either to redefine the aggregates to include the new instruments which are substituting for M1 money in the payments mechanism, as the Federal Reserve has done recently, or to attempt to model the process of innovation thereby restoring predictability to the relationship between money and nominal income.

The second line of research has followed the suggestion that the perceived stability of the demand for money prior to 1973 was in fact a misconception. The reality, it is maintained, was that a number of issues regarding the appropriate specification of money demand were swept under the rug because the data could not generate a resolution of them.[1] Consequently the pre-1973 debate has been reopened in the hope of generating a more robust specification of the demand for money.

In this paper I focus on the first approach and attempt to shed some light on the process of innovation utilizing Canadian data. The basic framework is that of Silber [1975, 1982] who suggests that innovation results from the attempt of banks to circumvent constraints imposed upon the banking industry or its customers. These constraints can arise either through centralized policy of the government or through the normal functioning of markets. Regardless of the source, however, this framework suggests that if changing conditions increase the shadow price of adhering to a constraint firms will have an incentive to undertake or intensify the search for new financial instruments.

One application of this constraint-induced innovation hypothesis has been carried out by Simpson and Porter [1980] who concentrate on the effects of high interest rates on the interest elasticity of M1 demand. Given higher opportunity costs of holding money, individuals will impute a higher rate of return to investment in new techniques of money management. In addition, the restriction on the payment of interest on demand deposits will induce banks to innovate in this area. This twofold effect, it is alleged, results in new instruments such as lock boxes or ATS accounts which lower the demand for M1. Furthermore, should interest rates subsequently decline the process will not reverse itself because the resource costs of the new techniques and instruments have been incurred. Simpson and Porter attempt to capture this effect through an interest rate ratchet variable which allows for a lag between peaks in interest rates and subsequent innovations. Their results, however, do not suggest that this technique is sufficient to restore confidence in the M1 function.

A more fundamental interpretation of circumventive innovation is suggested by the Kaldor hypothesis [1970] that an attempt by the authorities to control the growth rate of a monetary aggregate will result in a rise in the velocity of that money as private agents substitute to another payments mechanism.[2]

The choice of interpretation is of more than passing interest because of the profoundly different policy implications. The former suggests either a respecification of M1 demand or the choice of a broader aggregate to internalize any substitution which might occur. While it is recognized that during a period of rapid innovation it may be difficult to achieve desired stability and predictability in the money demand function, the basic thrust of this interpretation does not appear to undermine the case for control of a particular monetary aggregate. The latter interpretation, in contrast, does appear to question the basic policy. Bluntly put, any single-minded pursuit of a particular monetary target is destined to fail because private agents will innovate to escape the cutting edge of the control.

An attempt to provide empirical evidence for or against the latter hypothesis is hampered by the fact that the experience of the Federal Reserve with targeting on monetary growth rates has coincided with an environment of high market rates of interest and restrictions on the payment of interest on deposits. As argued above, this combination could be the trigger for innovation. In addition, the move to deregulation has spawned new instruments such as the NOW account and the new money market account; innovations which a priori should lead to a reduction in M1 demand. As a result, it does not appear possible to separate the effect, if any, of the Kaldor hypothesis from the forces mentioned above.

The Canadian experience with controlling monetary aggregates appears to be appropriate for an empirical test of the Kaldor hypothesis for the following reasons. First, the stated policy of the Bank of Canada from the fourth quarter of 1975 until the fourth quarter of 1982 was a gradual reduction in the rate of growth of M1 and, of equal importance, this policy was effected to a considerable degree.[3] Second, during this period there was no restriction on the payment of interest on deposits nor was there any substantial change in the regulatory environment facing the banking industry. Third, over the period in question there has been substantial financial innovation and, unlike the U.S. experience, this has occurred within the chartered banking system.[4] Finally, during this same period there has been a significant rise in M1 velocity after accounting for both trend and the rapid rise in interest rates.

Indeed, for the above reasons one could argue that the Canadian environment of 1975 through 1982 will closely resemble that of the United States in the decade ahead and should therefore be able to provide clues as to what difficulties await the Federal Reserve.[5]

In what follows, Section I discusses the nature of financial innovation in Canada and presents some evidence relating to the instability of the money demand function. Section II models the Kaldor hypothesis in the context of circumventive innovation and presents a test of the hypothesis. Section III summarizes the main conclusions of the paper.

Section I

A. Financial Innovation in Canada[6]

The financial innovations of the past eight years which have influenced the demand for M1 have occurred both in the corporate and household sectors. In the former category, the first major innovations occurred in the mid-seventies and they took the form of new cash management techniques which allowed corporations to minimize daily working balances. One of the more important of these is the centralized concentration account which allows for consolidation of several, perhaps geographically dispersed, accounts. A report is issued to the corporate treasurer the morning following deposits and he may allocate these funds as he sees fit. In addition, this period witnessed the introduction of regional lock boxes and preauthorized account withdrawals which reinforced the tendency to minimize working balances.

In the past two years, banks began to accept standing orders on how to employ surplus funds overnight. Two of the options offered are interest-bearing notice deposits (which are not included in the definition of M1) or the automatic paydown of outstanding demand loans. Either of these options would tend to reduce the demand for M1 balances.

On the household side, the major innovations have been the introduction of daily interest savings accounts and daily interest chequable savings accounts. Prior to the third quarter of 1979, chartered banks calculated interest on savings accounts based on the minimum monthly balance. Consequently, funds received during the month such as salary payments were deposited typically in personal chequing accounts. With the advent of daily interest accounts in August and September, 1979, individuals would have a much greater incentive to economize on demand deposits within the month.

The second innovation occurred in the latter half of 1981 with the introduction of daily interest chequable savings accounts. Prior to this change savings accounts could not as a rule be used for transactions purposes. The new accounts are actually hybrids of saving and demand in that interest is paid above a minimum balance and withdrawals, while subject to a fee, are not restricted. Since this is technically a notice deposit, it is not included in the definition of M1 and the spread of this account would tend to reduce M1 balances.[7]

The combined effect of these innovations has had a significant effect on the M1 demand function and the next section presents some evidence of this effect.

B. Stability of the Demand for Money

Landy [1980] presents some evidence on the stability of the Canadian M1 demand function since 1975. Utilizing the technique of out-of-sample dynamic forecasting, she identifies a break (downshift) about the second quarter of 1976. Hern [1980] has noted that dynamic forecasting tends to exaggerate the duration of any shift since it captures the lags in adjustment of money to the new level of demand. He argues that static forecasts will present a more accurate picture of any shift in the demand function. Consequently, it seems appropriate to reconsider the evidence.

The conventional money demand function is given by equation (1).[8]

$$\ln(M/P)_t = \alpha_0 + \alpha_1 \ln(M/P)_{t-1} + \alpha_2 \ln(RP)_t + \alpha_3 \ln(y)$$
$$+ \alpha_4 DUM1 + \alpha_5 DUM2 + \varepsilon_t \ , \qquad (1)$$

where $M \equiv$ currency plus chartered bank demand deposits (M1);

$P \equiv$ implicit GNP deflator (1971 = 100);

RP $\equiv$ 90 day prime corporate paper rate;

y $\equiv$ real GNP;

and

DUMi, i = 1, 2 ; $\equiv$ dummy variables to control for the effect of the interruption of the payments mechanism due to postal strikes.

The addition of dummy variables is necessitated by the apparent willingness of the Bank of Canada to accommodate the temporary increase in demand for liquidity which occurs during a postal strike. A strike tends to delay households' payments to firms but it does not interrupt some of firms' financial obligations such as payrolls. Consequently, an uncorrected money demand function will tend to underestimate money at the time of a strike.

Gregory and MacKinnon [1980] argue that the partial adjustment model which underlies equation (1) requires the addition of a lagged dummy whenever a dummy is included. The coefficient of the lagged dummy must be constrained to equal the negative of the product of the coefficient of the lagged dependent variable and the contemporaneous dummy. This procedure was followed in the estimation of equation (1).

Since the nature and duration of postal strikes has differed significantly, separate dummies were added for each strike. The coefficients for the strikes of 1965:3, 1968:3, 1970:2, 1975:1, 1980:3 and 1981:3 are insignificant are the ten percent level and are not included in the final regression. The coefficients for the strikes of 1974:2 and 1975:4 are significant at the five percent level and are essentially identical. In the final regression they are constrained to be equal. The coefficient for the strike of 1978:4 is significant at the ten percent level.[9]

Table 1 presents estimates of equation (1) for the 1956:1 - 1976:1 and the 1956:1 - 1982:3 sample periods. The choice of sample period follows that of Landy. While the differences between the samples are not as pronounced as that for the U.S. equation, the general pattern is maintained. The coefficient on the lagged dependent variable rises from 0.736 to 0.900. This implies that the mean adjustment lag increases from 3.8 to 10 quarters which suggests an implausible lag. The impact elasticity for real income is halved and the long-run interest elasticity rises from 0.20 to 0.55. The standard error of the equation also increases by 25 percent.

A standard F test for structural stability allows one to reject the hypothesis of stable coefficients across the hypothesized break point of 1976:1. The calculated F statistic of 5.09 exceeds the one percent critical value of 3.51.

Further evidence of the instability of the function is demonstrated by an analysis of the equation's forecasting ability. Post-sample static forecasts of equation (1), based on the coefficient estimates from the 1956:1 - 1976:1 regression, are presented in Table 2. In every quarter the equation over-predicts M1 demand. The root mean square error is more than three times the standard error of the in-sample equation and the fraction of error attributed to bias (one-sided prediction) is 58%.

Table 1

Equation 1 Regression Results

	Coefficients[1]						Summary Statistics	
Period	Constant	ln(M/P)-1	ln(RP)	ln(y)	DUM1	DUM2	S.E.E.	D.W.
1956:1 - 1976:1	-0.293 (4.99)	0.736 (16.40)	-0.053 (9.22)	0.201 (7.28)	0.030 (4.61)		0.0114	1.64
1956:1 - 1982:3	-0.378 (6.12)	0.900 (26.52)	-0.055 (8.49)	0.102 (4.63)	0.029 (3.84)	0.020 (1.91)	0.0143	1.76

1. The numbers in parentheses are absolute values of t-statistics. DUM1 set equal to 1 for 1974:2 and 1975:4, zero otherwise; DUM2 set equal to 1 for 1978:4, zero otherwise.

Table 2

Post-sample Static Forecast Errors 1976:2 - 1982:3[1]

Year/Quarter	Forecast Error (%)	Year/Quarter	Forecast Error (%)	Summary Statistics RMSE	BIAS
76:2	0.60	79:3	0.36	0.340	0.58
:3	0.21	79:4	0.69		
:4	0.61	80:1	0.25		
77:1	0.22	80:2	0.96		
:2	0.36	80:3	0.53		
:3	0.46	80:4	0.49		
:4	0.33	81:1	0.72		
78:1	0.31	81:2	0.36		
:2	0.24	81:3	1.00		
:3	0.13	81:4	1.90		
:4	0.34	82:1	0.69		
79:1	0.34	82:2	0.64		
:2	0.24	82:3	2.04		

[1] One-period ahead forecasts are based on the coefficient estimates from the 1956:1 - 1976:1 period. The forecast errors are the predicted less actual logs of real money balances.

While the above evidence would tend to support a claim that the demand for money has altered over the past seven years, it does not provide any clues as to the nature of the breakdown. If the function simply experienced a one-time change in intercept without a change in slope coefficients, then one could still retain confidence in the underlying economic relationships. However, if the instability were due to the omission of variables or to fundamental changes in relationships, then the implications for control of monetary aggregates are much more serious.

In order to gain a better understanding of the nature of the instability one can utilize the Breusch and Pagan [1979] test for random coefficient variation. The test statistic is one half the R^2 times the sample size of a regression of the squared residuals from equation (1) on the squared values of the explanatory variables.[10] This statistic is χ^2 with degrees of freedom equal to the number of explanatory variables. The calculated statistic of 13.82 exceeds the one percent critical value of 11.34. Therefore, we can reject the null hypothesis of zero coefficient variation over this period.

Since the post 1976:1 sample has been characterized by substantially higher interest rates on average, a log-linear form which constrains the interest elasticity to be constant might be expected to perform more poorly than a semilog form. Equation (1) was reestimated using the natural value rather than the log of the interest rate and was subjected to the same tests as outlined above. The F test statistic of 4.22 supports a rejection of the null hypothesis of structural stability at the one percent level. The out-of-sample predictions show less bias than those of equation 1 but there is a tendency to overpredict from 1976:2 to 1980:3 (sixteen out of eighteen quarters) and underpredict after 1980:3 (six out of eight quarters). In addition the Breusch and Pagan test statistic of 14.42 supports a rejection of the hypothesis of zero coefficient variation again at the one percent level. We conclude that the money demand function cannot be rehabilitated simply by a change in functional form.

Since the hypothesis under investigation deals with the velocity of M1, we have investigated this variable directly. Table 3 presents the results of a regression of the log of M1 velocity on a constant term, the log of the three-month treasury bill rate, and time for four subperiods between 1954 and 1982. The most striking feature of the results is that the estimated quarterly growth rate in M1 velocity averages 0.47 percent from 1954:4 to 1975:3 and 1.1 percent since the advent of targeting of M1 in 1975:3. Also, the interest rate variable for the last period is not significant. The F test statistic supports a rejection of the hypothesis of structural stability about 1975:3 at the one percent level of significance.

In anticipation of the discussion of Section II, Table 4 presents the results of an analysis of the velocities of currency, demand deposits and time deposits. This data demonstrates that the rapid acceleration in M1 velocity over the past seven years can be attributed primarily to the increase in demand deposit velocity. This result together with the fact that the time deposit velocity growth rate increases in absolute value by 150 percent over this same period indicates that significant deposit substitution has occurred since the targeting of M1 began.

Table 3

Estimated Rate of Growth in M1 Velocity

	Coefficients[1]					Summary Statistics[2]	
Period	Constant	Ln(RT)	Time	DUM1	DUM2	$\bar{R}^2$	ρ
54:4 - 61:3	1.96 (24.60)	0.030 (1.65)	0.005 (2.77)			0.97	0.97 (7.14)
61:4 - 68:3	2.23 (27.8)	0.068 (3.02)	0.005 (5.87)			0.99	0.67 (4.77)
68:4 - 75:3	2.33 (31.9)	0.049 (2.12)	0.004 (2.93)	-0.034 (2.51)		0.99	0.80 (7.30)
75:4 - 82:3	2.33 (23.90)	0.011 (0.29)	0.010 (6.18)	-0.039 (1.82)	-0.025 (1.50)	0.99	0.72 (5.30)

1. RT is the 3 month treasury bill yield. DUM1 is set equal to 1 for 1974:2 and 1975:4, zero elsewhere. DUM2 is set equal to 1 for 1978:4, zero elsewhere. The numbers in parentheses are absolute values of t-statistics.

2. ρ represents the value for the Beach and McKinnon [1978] adjustment for serial correlation.

Table 4

Estimated Rates of Growth of Currency, Demand and Time Deposit Velocities

	Velocity Growth Rates (% per quarter)		
Period	Currency	Demand Deposits	Time Deposits[1]
61:4 - 68:3	0.62	0.49	-0.42
68:4 - 75:3	0.19	0.54	-0.66
75:4 - 82:3	0.65	1.14	-1.65

1. Time deposits are defined as M3 - M1.

Section II

The maintained hypothesis under investigation is that the rise in M1 velocity above trend over the past eight years is a result of the attempt by private agents to circumvent Bank of Canada policy of a gradual reduction in M1 growth. The rationalization for this attempt at circumvention is straightforward. An announced policy of a reduction in the growth rate of a particular monetary aggregate should generate a reduction in expectations of future inflation, assuming that these expectations are formed in a rational manner, and subsequently, to a decline in nominal interest rates. If, however, the stance of fiscal policy in inconsistent with this announced policy (as it was in Canada over the period in question) then rational individuals may find the monetary policy lacking in credibility and they may not revise their expectations of inflation.[11] As a result, the ensuing reduction in liquidity will raise interest rates.

As individuals perceive higher opportunity costs of holding money balances they may attempt to substitute towards interest-bearing deposits. Substitution will be constrained by the fact that interest-bearing deposits are a less efficient payments mechanism than M1 money and individuals may look outside the banking system for alternatives.

Chartered banks have two options open to them to prevent competitors from attracting their deposits. They can pay interest on checking accounts or they can enhance the efficiency of other types of deposits as mediums of exchange through innovation. The tradeoff banks face in choosing between these alternatives is the eight percent differential in reserve requirements between demand and time deposits and the cost of innovation. Presumably some combination of the two options will be optimal which suggests that innovations will generate greater substitution from M1 to time deposits than in previous periods.

The above argument suggests that the rise in M1 velocity is linked directly to deposit substitution. It follows, therefore, that to test the proposition that M1 velocity is a function of the degree of control of M1 one requires a model of the decision process by which private agents determine the relative holdings of various monies.

Following Chetty [1969] and Moroney and Wilbratte [1976], we assume that households maximize financial wealth subject to a monetary transaction constraint.[12] Formally, define financial wealth in period t as

$$W(t) = M(t) + \sum_i X_i(t)[1+r_i(t)], \tag{2}$$

where $M(t)$ $\equiv$ nominal M1 money balances;

$X_i(t)$ $\equiv$ nominal holdings of the i^{th} class of interest-bearing assets;

and $r_i(t)$ $\equiv$ nominal interest rate of the i^{th} asset.

We assume the technology by which households combine money and interest-bearing financial assets is given by equation (3).

$$TS(t) = [\beta(t)M^{-\rho}(t) + \sum_i \beta_i(t)X_i^{-\rho_i}(t)]^{-\frac{1}{\rho}}, \tag{3}$$

where $TS(t)$ $\equiv$ volume of transactions services undertaken in period t;

$\beta(t), \beta_i(t)$ $\equiv$ technical coefficients on money and interest-bearing assets respectively;

and ρ, ρ_i $\equiv$ substitution parameters of money and interest-bearing assets respectively.

As Moroney and Wilbratte note, the above formulation assumes that the decisions affecting the relative portfolio holdings of money and assets X_i are independent of the yields on physical and human capital; an assumption which Bisignani [1975] tests and accepts using U.S. data. If, in addition, we assume that $\rho = \rho_i = \rho_j$ for all i and j (an assumption which Moroney and Wilbratte could not reject) then we can approximate equation (3) by its CES form.

Maximizing equation (2) subject to the CES form of equation (3) yields the following first-order condition.

$$m_i(t) = \tilde{\beta}_i^{-\sigma}(t) g_i^{\sigma}(t), \tag{4}$$

where $m_i(t)$ $\equiv$ the optimal ratio of M1 to X_i;

$\tilde{\beta}_i(t) \equiv \beta_i(t)/\beta(t)$, the relative technology coefficient of the i^{th} asset;

$\sigma \equiv 1/(1+\rho)$, the elasticity of substitution between M1 and X_i;

and $g_i(t) \equiv 1/(1+r_i(t))$.

Finally, we note that there is no reason to assume that $\tilde{\beta}_i$ remains constant over time especially given our knowledge of the trend rate of increase in M1 velocity. As a working hypothesis we assume that $\tilde{\beta}_i$ is a function of permanent income as specified in equation (5).

$$\tilde{\beta}_i(t) = \tilde{\beta}_0 Y^{\alpha_i}(t), \tag{5}$$

where $Y(t)$ $\equiv$ permanent income

and $\alpha_i \equiv$ the difference between the coefficients on β_i and β respectively.

The justification for equation (5) is twofold. First, the inventory approach to modelling money demand suggests that there are economies of scale associated with the level of income. We assume that these scale economies may be approximated by a rise in $\tilde{\beta}_i$ which is the relative technology coefficient of the i^{th} asset. Second, permanent income is a trend-demoninated variable and as such it may be expected to capture the effects over time of changes in the transactions demand for M1 money which are unrelated to the control of M1.

Substituting equation (5) into equation (4), taking logs and adding a dummy variable and an error term yields the equation to be estimated.

$$\ln m_i(t) = a_i + b_i \ln Y(t) + c_i \ln g_i(t) + \delta \mathrm{DUM1} + \varepsilon_i, \qquad (6)$$

where $a_i \equiv -\sigma \ln \tilde{\beta}_{i0}$;

$b_i \equiv -\sigma\alpha_i$;

$c_i \equiv \sigma$, the elasticity of substitution between M1 and X_i;

and DUM1 $\equiv$ 1 for 1974:2 and 1975:4, zero otherwise.

We can utilize equation (6) to test directly whether banks pursued a policy of deposit substitution as a result of central bank control of M1. If the hypothesis is correct, some form of structural instability of equation (6) should appear about the breakpoint of 1975:4 and with the addition of a suitable proxy variable we should be able to model this instability.

Defining M3-M1 as the interest-bearing asset and the chartered bank three-month deposit rate as the relevant interest rate, we estimated equation

(6) using quarterly data from 1961:1 to 1982:3. Y(t) and g(t) were estimated in distributed lag form since the estimation results of equation (1) suggests a mean lag in adjustment of actual to desired money balances of close to four quarters prior to 1975:3. The lags were estimated with a second degree polynomial. We imposed an endpoint restriction of zero for the seventh and fourth lags for Y(t) and g(t) respectively.

The initial estimate by ordinary least squares yielded a D.W. statistic which indicated positive autocorrelated disturbances and hence a specification error. Two possible sources of error are the assumption that the transactions technology is of a CES type and that the portfolio relationship between M1 money and time deposits of chartered banks is independent of the yield on physical or human capital.

Utilizing equation (3) which implies a more general technology, we derived an equation corresponding to equation (6) and estimated it over the same sample period with M2-M1 and M3-M1 as the relevant alternative assets. Again, the D.W. statistic indicated positive autocorrelation and the correlation correction yielded estimates of rho which were not significantly different from that of equation (6). Therefore we conclude that the assumption of a CES technology is not a source of error in equation (6).

If the assets in the transactions constraint are not separable from physical or human capital, then the $m_i(t)$ which we estimated would not be invariant to changes in the yields on these assets. In order to test for this type of relationship, we included an inflationary expectations proxy in equation (6). The proxy variable is that of Riddell and Smith [1982] who generate forecasts from an ARIMA model estimated from a moving sample of 384 monthly observations beginning in 1921. The method of moving sample insures that agents form expectations on previous experience only rather than on the basis of experience over the entire sample.[13] Again the estimation yielded an estimate of rho not significantly different from that of equation (6). We conclude that, to the extent that our proxy accurately measures inflationary expectations, the assumption of separability of physical capital from financial assets in the transactions technology is not the cause of the specification error.[14]

Accordingly, equation (6) was reestimated using the Cochrane-Orcutt method of correction for autocorrelation. Prior to the presentation of the regression results let us review the a priori restrictions on the coefficients. First, since the sum of the c coefficients is the elasticity of substitution between M1 and time deposits, it should be positive. Second, given our knowledge of the trend rates of growth of M1 and time deposit velocities, the sum of the b coefficients is expected to be negative Third, $\tilde{\beta}_0$ is the initial estimate of the technical coefficient of time deposits relative to M1 money in the transactions technology and as such we expect it to lie in the unit interval. Since the elasticity of substitution is positive, the restriction on $\tilde{\beta}_0$ can be restated as a restriction that the constant is nonnegative.

Consider the first row of Table 5 which presents the results of the regression of equation (6) over the sample period 1962:4-1975:4. The constant is positive and highly significant. The coefficients on current and lagged permanent income are significant and their sum is negative as expected but the first two coefficients are positive. This result tends to support

Table 5

Regression Results of Equation (6)

	Coefficients[1]													Summary Statistics[2]			Calculated Parameters			
Period	a	b_0	b_1	b_2	b_3	b_4	b_5	b_6	c_0	c_1	c_2	c_3	δ	$\bar{R}^2$	S.E.E.	RHO	σ	$\tilde{\beta}_0$	$\sum b$	α
1962:4	3.94	1.06	0.61	-0.13	-0.61	-0.83	-0.83	-0.54	1.09	0.87	0.62	0.33	0.05	0.99	0.0167	0.75	2.91	0.26	-0.70	0.24
1975:4	(8.50)	(2.43)	(2.14)	(10.01)	(3.31)	(3.01)	(2.91)	(2.87)	(3.06)	(4.30)	(2.60)	(1.72)	(4.26)			(8.13)				
1977:3	-2.27	4.80	2.15	-0.29	-1.39	-2.10	-2.11	-1.41	0.67	1.33	1.44	1.00		0.98	0.0229	0.05	4.44	1.67	0.16	-0.04
1982:3	(0.42)	(4.16)	(3.71)	(0.21)	(5.97)	(5.33)	(5.11)	(5.00)	(1.93)	(5.90)	(6.68)	(6.26)				(0.20)				
1962:4																				
1975:4	3.63	3.34	1.36	-0.12	-1.10	-1.57	-1.55	-1.03	0.93	1.24	1.18	0.77	0.05	0.99	0.0216	0.83	4.11	0.41	-0.66	0.16
1972:3	(5.46)	(5.27)	(4.90)	(6.83)	(6.46)	(6.05)	(5.92)	(5.85)	(3.89)	(9.01)	(7.43)	(6.02)	(3.66)			(12.67)				
1982:3																				

[1] All numbers in parentheses are absolute valus of t-statistics.

[2] Rho is the Cochrane and Orcutt estimate of AR1 serial correlation.

the view that individuals view M1 money, among other things, as a temporary abode of purchasing power until it can be allocated in an optimal fashion among all assets.

The coefficents on current and lagged interest rates are significant and possess the correct sign. The sum of these coefficients which represents the elasticity of substitution between M1 and time deposits is 2.91. This compares to an elasticity of 6.09 between M1 and chartered bank personal savings deposits obtained by Short and Villanueva [1977] using annual Canadian data over the period 1951-1973. The calculated initial value of $\tilde{\beta}$ is 0.26 which satisfies the a priori restiction.

Since rho is insignificant from one at the 5 percent level, the equation was reestimated in first difference form. The individual lag coefficients and the mean lags are statistically identical as is the elasticity of substitution. We conclude that equation (6) is a robust specification of the relative asset demand of M1 and M3-M1 over the sample period in spite of the restrictive assumptions used in its derivation.

Row two of Table 5 presents the results of the regression over the post-policy-change sample. The results indicate significant structural instability. The constant is of the wrong sign and together with the elasticity of substitution of 4.44 implies an initial value of $\tilde{\beta}$ of 1.67. Also, the sum of the b coefficients is positive counter to a prior expectations. The calculated F statistic of 11.01 exceeds the one percent critical value of 4.11, a result which supports a rejection of the null hypothesis of structural stability.

Perhaps the most striking result of this regression is the fact that the coefficient of serial correlation turns negative and is insignificant for the post-75:4 sample. Given the difficulty in eliminating autocorrelation in the sample 1962:4 - 1975:4, this is indeed surprising. Since the total sample still exhibits autocorrelation, one might be tempted to explain the disappearance of autocorrelation with technical arguments such as problems with the number of degrees of freedom. However we are unable to find any other twenty quarter sample which does not possess autocorrelation. Consequently we interpret the disappearance of autocorrelation as evidence of a fundamental change in the actions of private agents resulting from the policy change of the Bank of Canada which occurred in 1975.

If the hypothesized relationship between control of the money supply and innovations in the underlying transactions technology is valid, then it should be possible to model these innovations with an additional variable which reflects the degree of control. In other words, equation (6) is incompletely specified for the post-control sample and we require another variable which measures the degree of central bank control of M1.

The hypothesis suggests that chartered banks will have an incentive to utilize costly resources to effect deposit substitution when they perceive a reduction in the rate of growth of money relative to some average of past rates. There are many reasons for believing that this relationship may be subject to significant lags. Banks will wish to insure that current downturns reflect a permanent reduction in the rate of growth of money before undertaking expenditures for innovations. There may be delays before research and development efforts culminate in innovations. Also, bringing the new technology on line may involve significant lags. Consequently it seems appropriate to model these effects with a variable which provides flexibility in the relationship between the rate of growth of money and innovations.

With the above in mind, we propose a variable defined as the difference between two moving averages of the rate of growth of M1.[15] Formally, let μ denote the quarterly growth rate of M1. We define S(t) by[16]

$$S(t) = \frac{1}{m} \sum_{i=t-m+1}^{i=t} \mu_i - \frac{1}{n} \sum_{j=t-n+1}^{j=t} \mu_j \quad , \; m > n. \tag{7}$$

Since m exceeds n, a positive S reflects a sustained reduction in the rate of growth of M1 which, by hypothesis, should result in deposit substitution from M1 to time deposits.

The evidence in Table 5 indicates that the structural change in the relative asset demand equation is centered primarily about the constant term and by implication $\tilde{\beta}_0$. Accordingly, we propose as a working hypothesis that $\tilde{\beta}_i$ is a function of S as well as permanent income. Formally, we assume that

$$\tilde{\beta}_i(t) = \tilde{\beta}_0 Y^{\alpha_i}(t) e^{\gamma_i S(t)} , \tag{8}$$

where $\gamma_i > 0$.

That is, a rise in S will generate innovations which, by assumption, impact on the relative technology coefficient of time deposits. Substituting equation (8) into equation (4), taking logs, and adding a dummy variable and error term yields

$$\ln m_i(t) = a_i + b_i \ln Y(t) + c_i \ln g_i(t) + \delta DUM1 + d_i S(t) + \varepsilon(t), \quad (9)$$

where $d_i \equiv -\sigma\gamma_i$.

A priori, we expect d to be zero for the period 1962:4 - 1975:4 and negative for the period 1977:3 - 1982:3.

Equation (9) was estimated for the periods in question and the results are presented in Table 6. For the period 1962:4 - 1975:4, the coefficient on S is insignificant as hypothesized. The constant term standard error declines but there is no appreciable change in the lag structure or sum of lags on permanent income. The last two lags on the interest rate turn insignificant. The standard error of the regression shows little change.

For the period 1977:3 - 1982:3, the coefficient on S possess the correct sign and is highly significant. The constant term turns positive and is highly significant. The estimate of $\tilde{\beta}_0$ accordingly declines to 0.31 from 1.67. The standard errors of the b coefficients decline appreciably and the sum of these coefficients is now negative. The standard error of the equation declines by 55 percent. The addition of S also induces negative serial correlation.

Since the results in Table 6 indicate that the innovation proxy performs as hypothesized, we reestimated equation (9) with the restriction that d = 0 for the period 1962:4 through 1977:4. The results of these regressions are presented in Table 7.

A comparison of rows one and two reveals surprising similarity in the underlying parameters of the transactions services technology. The elasticity of substitution is 2.91 in the first sample and 3.14 in the second sample. The standard errors are 0.68 and 0.41 respectively. The initial values of $\tilde{\beta}$ are of

Table 6

Regression Results of Equation (9)

Period	Coefficients[1] a	b_0	b_1	b_2	b_3	b_4	b_5	b_6	c_0	c_1	c_2	c_3	δ	d	Summary Statistics[2] $\bar{R}^2$	S.E.E.	RHO	Calculated Parameters σ	$\tilde{\beta}_0$	$\sum b$	α
1962:4	4.06	1.50	0.57	-0.13	-0.58	-0.80	-0.77	-0.51	0.98	0.75	0.51	0.26	0.04	-0.61	0.99	0.0171	0.78	2.49	0.20	-0.72	0.18
1975:4	(7.78)	(2.20)	(1.91)	(8.95)	(3.09)	(2.79)	(2.69)	(2.64)	(2.45)	(2.63)	(1.71)	(1.17)	(2.68)	(0.68)			(8.88)				
1977:3	3.66	3.60	1.47	-0.12	-1.18	-1.69	-1.67	-1.01	0.73	0.94	0.89	0.58		-6.25	0.99	0.0104	-0.56	3.14	0.31	-0.68	0.22
1982:3	(2.33)	(9.31)	(7.98)	(3.06)	(14.51)	(12.65)	(12.04)	(11.74)	(6.56)	(12.15)	(9.49)	(7.85)		(8.43)			(3.03)				
1962:4 1975:4	4.02	2.87	1.15	-0.13	-0.97	-1.39	-1.36	-0.90	0.81	1.03	0-97	0.63	0.03	-1.78	0.99	0.0211	0.92	3.43	0.30	-0.72	0.21
1972:3 1982:3	(3.32)	(3.80)	(3.42)	(4.11)	(5.05)	(4.61)	(4.47)	(4.40)	(3.33)	(6.04)	(5.26)	(4.49)	(1.89)	(2.31)			(19.26)				

[1] All numbers in parentheses are absolute values of t-statistics.

[2] Rho is the Cochrane and Orcutt estimate of AR1 serial correlation.

Table 7

Regression Results of Equation (9)[1]

	Coefficients[2]														Summary Statistics[3]			Calculated Parameters			
Period	a	b_0	b_1	b_2	b_3	b_4	b_5	b_6	c_0	c_1	c_2	c_3	δ	d	$\bar{R}^2$	S.E.E.	RHO	σ	$\tilde{\beta}_0$	$\sum b$	α
1962:4	3.94	1.60	0.61	-0.12	-0.61	-0.83	-0.81	-0.53	1.09	0.87	0.62	0.33	0.05	0	0.99	0.0172	0.77	2.91	0.26	-0.70	0.24
1975:4	(8.41)	(2.40)	(2.12)	(9.89)	(3.27)	(2.98)	(2.88)	(2.83)	(3.02)	(4.26)	(2.57)	(1.70)	(4.22)				(8.66)				
1977:3	3.66	3.60	1.47	-0.12	-1.18	-1.69	-1.67	-1.10	0.73	0.94	0.89	0.58		-6.25	0.99	0.0104	-0.56	3.14	0.31	-0.68	0.22
1982:3	(2.33)	(9.31)	(7.98)	(3.05)	(4.51)	(2.65)	(2.04)	(1.74)	(6.55)	(12.15)	(9.49)	(7.85)		(8.43)			(3.03)				
1962:4																					
1975:4	3.90	3.0	1.21	-0.13	-1.01	-1.44	-1.41	-0.93	1.02	1.05	0.88	0.54	0.05	-6.19	0.99	0.0189	0.83	3.49	0.33	-0.71	0.20
1972:3	(6.84)	(5.43)	(5.0)	(8.45)	(6.81)	(6.34)	(6.18)	(6.11)	(4.91)	(8.40)	(5.87)	(4.43)	(4.30)	(4.79)			(12.5)				
1982:3																					

[1] These results are obtained under the constraint that d = 0 for the period 1962:4 - 1975:4.

[2] All number in parentheses are absolute values of t-statistics.

[3] Rho is the Cochrane and Orcutt estimate of AR1 serial correlation.

the correct magnitude and are very close. The sum of the b coefficients should be considered identical.

Given the rather large differences in technology between the two samples implied by Table 5, we consider the conformity achieved by the addition of S to be quite remarkable. Recall that we assumed innovations impact on the relative technology coefficient rather than on the elasticity of substitution. Also, our proxy variable for innovations could be considered crude at best. Yet, inspite of these restrictive assumptions we are able to achieve a high degree of conformity in parameters over the two samples. We conclude that this evidence supports the hypothesis that the rise in the velocity of M1 above trend over the past eight years can be linked causally to the decision of the Bank of Canada to target on the growth rate of M1.

III Conclusions

The experience of Canada with a central bank policy of targeting M1 growth suggests that there may be a fundamental difficulty in controlling the rate of inflation through a policy of targeting on this variables. The evidence presented in this paper indicates that the degree of control of M1 growth does have a significant role to play in explaining the rise in the velocity of M1 and, as a consequence, the inflation rate has taken much longer to respond to restrictive monetary policy than most observers would have predicted prior to the policy inactment.

It would appear that the noninterest-bearing characteristic of M1 plays an important role in explaining this causal link since it may, on occasion, provide chartered banks with an incentive to increase the monetary effectiveness of other classes of deposits. However, this interpretation is not necessarily warranted. Recall that Canada is free from restrictions on the payment of interest on any type of deposit. Accordingly, chartered banks could have initiated interest payments on demand deposits in the face of sharply higher market rates of interest. We argued that the differential reserve requirement probably influenced the decision to innovate rather than pay interest. Yet, as long as the option of innovation is open to the private sector, there will always be the possibility of nonprice competition through the enhancement of the transactions efficiency of deposits.

The characteristic which appears to be critical in explaining the rise in M1 velocity is the scope of the targeted aggregate. A narrow definition insures a wide selection for deposit substitution through innovation. Therefore, the case for control of a narrow aggregate would appear to be weakened considerably by the evidence presented in this paper.

Footnotes

1. See Laidler [1980] for a discussion of some of these issues.

2. Kaldor appears to be the first to present this view although it has been proposed as well by Holland [1975] who coined the phrase circumventive regulation. Notes that this interpretation inverts the adage that expansionary monetary policy is akin to "pushing on a string."

3. The annual rate of growth of M1 from 1975:3 to 1982:3 has been 7.3, 8.7, 10.2, 7.9, 4.6, 4.3, and 0.1 percent respectively. This compares to an average annual rate of 13.1 percent over the previous five years.

4. See Landy [1980] and Silber [1982] for a discussion of the innovations in the U.S.

5. On this note, it should be pointed out that in November of 1982 the Bank of Canada publicly abandoned its policy of targeting on M1 growth because of the instability of the demand function and attendant difficulties.

6. This section draws heavily upon Freedman [1982] and Landy [1980]. The interested reader is referred to these articles for a more complete discussion.

7. The reluctance of banks to offer pure interest bearing checking accounts is no doubt due to the differential of 8% in the reserve requirement on these and savings accounts.

8. All variable definitions and data sources can be found in the appendix.

9. It is surprising that the strike of 1981:3 which lasted 43 days does not generate the hypothesized underprediction of money. Could it be that the general public is adapting to the frequent interruptions of postal service in Canada?

10. The regression was run over the period 1972:4 to 1982:3. While the sample contains observations prior to the hypothesized break about 1976:1, it was necessitated by the minimum test requirement of forty observations.

11. For a formal derivation of this proposition the reader is referred to Sargent and Wallace [1981].

12. This interpretation is due to Moroney and Wilbratte [1976]. In the original article Chetty assumed a CES utility function and maximized utility subject to a wealth constraint. The difficulty with this interpretation is that it assumes the only motive for holding assets is to facilitate transactions services.

13. See Friedman [1979] for the theoretical argument for this type of approach.

14. Moroney and Wilbratte also report positive autocorrelation in their estimates of equation (6) for the U.S. using government debt and corporate debt as well as time deposits of commercial banks. Nor are they able to explain the cause of the specification error.

15. In an earlier attempt we constructed a ratchet variable utilizing rates of growth of M1 of a type proposed by Simpson and Porter [1980]. The use of a ratchet variables seems warranted because once the costs of an innovation have been incurred it will be maintained even if the forces leading to its adoption are mitigated. However we found that the use of this type of variable results in insignificant b coefficients. We interpret this result as evidence of multi-collinearity between permanent income and the ratchet variable, as might be expected, since both variables are dominated by trend. Consequently, we followed the approach described in the text.

16. The values of m and n were determined by the data. The best results were obtained with m equal to 20 and n equal to 5.

Appendix

All data in this study were supplied by CANSIM. Interest rate observations and money holdings are quarterly averages of seasonably adjusted monthly data.

M1: Currency and demand deposits - Series B1609.

M3: Currency and all checkable, notice, and personal term deposits plus Canadian dollar non-personal fixed term deposits and bearer term notes - Series B1603.

r: Chartered Bank 90 Day deposit rate - Series B14018.

rp: Prime corporate 90 day paper rate - Series B14017.

rt: 90 Day Treasury bill yield - Series B14007.

Y: The permanent income series was constructed using de Leeuw's [1965] formula:

$$Y(t) = 0.114 \sum_{t} (0.9)^{t} (y)_{-t} \quad , \; t = 0 \ldots 19,$$

where y is GNP deflated by the implicit price deflator.

GNP: Gross national product at market prices - Series D40252.

P: GNP Deflator - Series D40625.

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