TEMPORAL CROSS-SECTION SPECIFICATIONS OF THE DEMAND FOR DEMAND DEPOSITS

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OVER A DECADE has passed since Gurley and Shaw [5] alerted the economics profession to the importance of considering the implications of the growth of non-bank financial liabilities for monetary theory and policy. The empirical question at the heart of the issue concerning the effects of non-bank financial intermediaries focused on the degree of substitutability and the stability of the substitution relationship between demand deposits and other liquid assets. A large number of empirical studies were subsequently undertaken to examine the substitution relationships and the purpose of this paper is to inquire whether some consensus emerges from the empirical literature which examines demand functions for demand deposits utilizing pooled temporal cross-section data.

Section I reviews and evaluates the existing cross-section evidence and suggests that while a fairly broad consensus emerges on the substitution relationships between demand deposits and other liquid assets, the differential use of dummy variables in the model specifications appears to be an important factor in explaining substantive variations between studies. Section II considers a covariance model of the demand for demand deposits which permit testing of alternative dummy specifications. Section III reports results based on a random coefficient model which is more restrictive than the covariance model with respect to temporal dummy variables but allows a richer specification of variation in behavior across states.

I. A REVIEW OF THE EXISTING STUDIES OF THE DEMAND FOR DEMAND DEPOSITS BASED ON TEMPORAL CROSS-SECTION DATA

The theoretical contributions of Gurley and Shaw [5]; the empirical postwar reversal of the well established secular decline in velocity[1]; and the dramatic growth in non-bank financial intermediaries reinforced the prevalent notion that demand deposits and other liquid assets were highly competitive. Substantive empirical work was expected to establish the extent to which substitution relationships were stable and predictable.

Given this intellectual climate, it was surprising to discover that systematic efforts to estimate the elasticities of substitution between demand deposits and other intermediary liabilities could not confirm the importance of the substitution relationships. The tension between the apparently overwhelming casual evidence of high substitutability and the systematic evidence of the absence of strong substitution led Laidler [10] in his excellent survey of the theoretical and empirical issues concerning the definition of money to cautiously conclude,

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1. Selden [15].

With the evidence we have at the moment, it is possible to come to only tentative conclusions about what set of assets the monetary authorities should attempt to manipulate in carrying out monetary policy . . . . However, there does seem to have been some change over time, particularly since the second world war, in the assets to whose rate of return the demand for money however defined is most sensitive. The apparent importance of savings and loan associations in recent years is particularly noteworthy. These conclusions must, however, be treated with care, since they are contradicted by Feige's important study.

Since the process of "disintermediation" and "reintermediation" appears to be as dramatic in the 1970's[2] as it was in the earlier decades, it is difficult not to share Laidler's scepticism concerning the failure of several attempts at finding economically significant elasticities of substitution between key liquid assets. The evidence from temporal cross-section studies of the demand deposits indicates that the debate concerning the degree and stability of substitution relationships has not come to rest. At the one extreme are the findings of Feige [3], Hartley [6], and Kichline [9] which conclude,

demand deposits do not appear to be close substitutes for other liquid assets.[3]

demand deposits and time deposits in commercial banks are weak substitutes . . . demand deposits and savings and loan shares are strong complements.[4]

the limited substitutability of demand deposits for income yielding depository type claims does not support Gurley and Shaw's contention that such deposits exhibit a high degree of substitutability.[5]

At the other extreme is the temporal cross-section evidence of Lee [11], who concludes that,

savings and loan shares are close substitutes for money. . . . The empirical evidence from the various data investigated unequivocally substantiates the substitution hypothesis of Gurley and Shaw.[6]

In order to attempt to reconcile these diametrically opposed viewpoints, I wish to consider the evidence in some detail, and deal with several of the critiques which have been raised concerning the reliability of the temporal cross-section studies. The studies considered in this paper share many common elements. Almost all of these studies estimate a demand function for demand deposits using as arguments income and the rates of return on demand deposits, commercial bank time deposits, savings and loan deposits and mutual savings deposits. With the exception of Lee's [11] study dealing with individual household survey data, all the studies use as the basic unit of observation area-wide holdings of demand deposits of individuals, corporations and partnerships. The data span the period 1949-1964.

Laidler [10] has leveled several criticisms against Feige's original study which deserve explicit attention since the issues raised are common to all of the studies reported in this paper.

2. See Economic Report of the President [2].

3. Feige [3], p. 43.

4. Hartley [6], p. 80.

5. Kichline [9], p. 29.

6. Lee [11], pp. 452 and 455.

Laidler [10] correctly suggests that to the extent that different ownership classes exhibit different portfolio behavior, demand deposits data which combine the holdings of individuals, partnerships and corporations may represent too gross a measure. Unfortunately, no data are directly available by ownership classification, and attempts to correct this problem by use of ancillary ownership survey data have shown the estimates of substitution elasticities to be quite insensitive to ownership adjustments.[7]

The second problem raised by Laidler [10] is that the deposit data by states reflects the ownership of deposits by residents of all states. Thus, even if out of state ownership of deposits represented only a small fraction of the total, these deposits might be quite sensitive to rates of return in all states and thus represent a specification error in the temporal cross-section models. A casual test of this hypothesis can be carried out by examining the pattern of residuals between adjoining states where this type of cross-over is likely to be prevalent. My own investigation of residual patterns displayed no systematic cross-over effect. Moreover, to the extent that such an effect exists, it is likely to be captured in dummy variables which take account of centers of financial activity.

The final critique by Laidler [10] deals with the common usage of measuring the negative rate of return on demand deposits as the ratio of service charges to average deposit balances. Laidler argues,

> the ratio of bank charges to deposits may not remain a constant in the absence of changes in the interest rate on demand deposits. For a given volume of payments this ratio might be expected to rise as the quantity of demand deposits falls, for it should be related to the rate of deposit turnover. Bank charges are primarily bookkeeping charges and are related to the flow of payments made by check rather than to the stock of demand deposits in existence. If the volume of check payments is roughly related to the level of permanent income, one would expect a negative relationship to hold between the ratio of bank charges to deposits and the level of demand deposits held, given the level of permanent income, a relationship such as Feige finds. It would arise from variations in the rate of turnover implied by their being held in different ratios to permanent income. This relationship though would reflect the nature of bank charges and not the influence of the rate of return on demand deposits.[8]

Laidler concludes that the demand deposit rate proxy measures the true return very inaccurately, but he can not access the effects of this inaccuracy on the elasticity estimates. Since the negative return on demand deposits is used in all the studies under consideration here, and is statistically significant for all studies, the rationale for its use requires further explication.

Demand deposits yield a flow of non-pecuniary services which are assumed to be roughly proportional to the stock of deposits held. Commercial banks charge the deposit holder for the services provided by means of a service charge. The service charge, taken as a percentage of the average deposit balance, is viewed as the negative pecuniary rate of return on demand deposits. An increase in the cost of holding cash balances can be viewed in two ways.

7. See Feige [3], and Hartley [6]

8. Laidler [10], p. 522.

An increase in costs can be seen as compensation to the bank for the provisions of additional services per dollar per unit of time. Alternatively, it may simply represent an increase in the per unit cost of services. In the former instance, the increased charge is simply an equalizing differential which represents the marginal cost of additional services per dollar of deposit. The marginal holder of deposits would presumably be indifferent between paying the extra cost of the incremental flow of services per dollar of deposit or retaining the original flow of services per dollar and avoiding the incremental cost. Thus, if increased service charges merely reflect an increase in the flow of non-pecuniary services per dollar of deposit, as suggested by Laidler, this increase is unlikely to affect the quantity of demand deposits demanded. Alternatively, if the increase in the cost of services is not offset by an increase in the flow of services per dollar of deposit, this represents to the holder of demand deposits a real increase in the cost of his deposits and will lead him to reduce his holdings of demand deposits and increase his holdings of some other income yielding asset. Since the non-pecuniary services per dollar of deposit are not directly measurable, one must determine empirically the extent to which an increase in service charges is offset by an increase in service flow per dollar by examining the relationship between service charges and the demand for demand deposits. If changes in service charges have no effect on the desired level of deposits, one can conclude that such changes merely represent equalizing differentials which are offset by an increase flow of services per dollar of deposit. Alternatively, if increased service charges are associated with a reduction in the demand for the stock of demand deposits, one can infer that the negative rate of return (or cost) of demand deposits has increased and thus induced individuals to shift out of demand deposits and into other assets. In the absence of direct information on the change in the flow of services per dollar, the estimated elasticity based on service charges must be regarded as an underestimate of the true elasticity.

The Models and the Evidence

The temporal cross-section models reviewed in this section can be characterized by the following model specification. Given N states and T time periods, y_{nt} represents an observation of the per capita level of demand deposits for the n^{th} state and the t^{th} time period. Let x_{jnt} represent an observation on the j^{th} independent variable (of which there are K) for the n^{th} state and the t^{th} time period and z_{lnt} represent the l^{th} dummy variable (of which there are D) for the n^{th} state and t^{th} time period.

The model we are concerned with is

$$y_{nt} = \alpha + \sum_{j=1}^{K} \beta_j x_{jnt} + \sum_{l=1}^{D} \gamma_l z_{lnt} + \epsilon_{nt} \tag{1.1}$$

where ϵ_{nt} is a random error term. In matrix notation the model can be written as

$$Y = \left[i \;\vdots\; X \;\vdots\; Z \right] \begin{pmatrix} \alpha \\ \cdots \\ \beta \\ \cdots \\ \gamma \end{pmatrix} + \epsilon \tag{1.2}$$

where

$$Y = \begin{bmatrix} y_{11} \\ \cdot \\ \cdot \\ \cdot \\ y_{1T} \\ \cdot \\ \cdot \\ \cdot \\ y_{N1} \\ \cdot \\ \cdot \\ \cdot \\ y_{NT} \end{bmatrix}_{(NT \times 1)}, X = \begin{bmatrix} x_{111} & \cdots & x_{K11} \\ \cdot & & \cdot \\ \cdot & & \cdot \\ \cdot & & \cdot \\ x_{11T} & \cdots & x_{K1T} \\ \cdot & & \cdot \\ \cdot & & \cdot \\ \cdot & & \cdot \\ x_{1NT} & \cdots & x_{KNT} \\ \cdot & & \cdot \\ \cdot & & \cdot \\ \cdot & & \cdot \\ x_{1NT} & \cdots & x_{KN1} \end{bmatrix}_{(NT \times K)}, Z =$$

$$\begin{bmatrix} z_{111} & \cdots & z_{D11} \\ \cdot & & \cdot \\ \cdot & & \cdot \\ \cdot & & \cdot \\ z_{11T} & \cdots & z_{D1T} \\ \cdot & & \cdot \\ \cdot & & \cdot \\ \cdot & & \cdot \\ z_{1NT} & \cdots & z_{DN1} \\ \cdot & & \cdot \\ \cdot & & \cdot \\ \cdot & & \cdot \\ z_{1NT} & \cdots & z_{DNT} \end{bmatrix}_{(NT \times D)}, \epsilon = \begin{bmatrix} \epsilon_{11} \\ \cdot \\ \cdot \\ \cdot \\ \epsilon_{1T} \\ \cdot \\ \cdot \\ \cdot \\ \epsilon_{N1} \\ \cdot \\ \cdot \\ \cdot \\ \epsilon_{NT} \end{bmatrix}_{(NT \times 1)} \quad (1.3)$$

and i is a $NT \times 1$ vector of ones, α is a scaler, β is a $K \times 1$ vector of behavior coefficients and γ is a $D \times 1$ vector of dummy variable coefficients. The X matrix contains observations on income and the rates of return on various liquid assets and the Z matrix includes regional dummy variables. Under some specifications considered below the dummy variables are entirely suppressed, and this specification is characterized by constraining the γ vector $= 0$.

In order to gauge the degree of consensus between different temporal cross-section estimates of the demand functions for demand deposits, Table 1 summarizes the critical elasticity estimates from different studies. Table 1 also indicates the estimation procedure employed, the time period under study, the functional form utilized, the unit of observation and the type of dummy variable specification used. A more complete specification of the specific dummy variables used in each of the studies is available from the author upon request.

The first cross-sectional study which revealed weak substitutability between demand deposits and other liquid assets is Feige [T1-1]. After estimating separate demand functions for each year, I tested and could not reject the hypothesis that the coefficients were stable over the entire period 1949-59. The estimates presented as [T1-1] reflect the pooled temporal cross-section estimates. Hartley [T1-2] successfully replicated my results. The first empirical challenge to these surprising results appeared in Lee's [T1-3] recalculation of my estimates. Lee argued that "the strange results in Feige's study stem from an excessive use of dummy variables."[9] Lee asserted that since the regional dummy variables were highly correlated with the included interest

9. Lee [11].

TABLE 1

SUMMARY OF EXISTING POOLED TEMPORAL CROSS-SECTION EVIDENCE FROM DEMAND FUNCTIONS FOR DEMAND DEPOSITS[a]

Reference Number	Author	Estimation Procedure[b]	Time Period	Functional Form	Unit of Observation	Dummy Variable Specification[c]	Elasticity Estimates[d]				
							$\eta_{r_{dd}}$	$\eta_{r_{dt}}$	$\eta_{r_{ds}}$	$\eta_{r_{dm}}$	η_y
[T1-1]	Feige	OLS	1949-59	Linear	State	RD	.31*	−.10*	.30*	.04	.92*
[T1-2]	Hartley (replication)	OLS	1949-59	Linear	State	RD	.31*	−.10*	.30*	.04	.92*
[T1-3]	Lee	OLS	1949-59	Linear	State	ND	.37*	.02	−.31*	−.05*	1.36*
[T1-4]	Hartley (replication)	OLS	1949-59	Linear	State	ND	.39*	−.06	.06	−.05*	1.43*
[T1-5]	Hartley	OLS	1949-64	Linear	State	ND	.34*	−.19*	.11	−.06*	1.46*
[T1-6]	Hartley	OLS	1949-64	Linear	State	RD	.27*	−.18*	.32*	−.02	.97*
[T1-7]	Feige	REE	1949-53	Linear	State	RD	.32*	−.21*	.01	.13*	.98*
[T1-8]	Feige	REE	1954-58	Linear	State	RD	.38*	.06	−.19*	.17*	.94*
[T1-9]	Lee	OLS	1956-59	Linear	Household	RD	.49*	.40	−1.61*	—	N.A.
[T1-10]	Kichline	OLS	1965-66	Linear	SMSA	RD	N.A.	.001	−.006	−.005*	.80*
[T1-11]	Cohen	OLS	1950-60	Log-Linear	State	RD	.22*	−.34*	—	—	1.12*

[a] All elasticities are computed at the mean for linear functions.

[b] OLS denotes ordinary least squares; REE denotes restricted efficient estimation.

[c] RD denotes regional dummies, ND denotes no dummies.

[d] * denotes significance at the 5% level.

rate variables, "such high multicollinearity clearly indicates the spuriousness of Feige's regression estimates." As Friedman and Schwartz [4] have subsequently pointed out,

> Feige uses these variables to allow for special circumstances of particular states or regions, such as states that permit establishment of mutual savings banks and those that do not and states that contain the main financial centers of each of four regions and those that do not. As evidence that the use of dummy variables is excessive, Lee cites multicollinearity between them and other independent variables. This does suggest a real problem with the statistical stability of Feige's estimates. However, it certainly does not mean that if the special features Feige seeks to control are present, as they clearly are, correct results can be obtained by neglecting them, as Lee does in a regression using Feige's data (though not in one using Consumer Survey data).

The major effect of Lee's omission of the dummy variables was to produce an estimate of the elasticity of substitution of demand deposits for savings and loan shares of −.31, which by conventional standards would not be regarded as "strong substitutability." More problematic, however, was the failure of Hartley's study [T1-4] to replicate Lee's results. Hartley, using the same data but omitting the dummy variables found a nonsignificant cross elasticity of .06. When Hartley [T1-5] extended the data to include the period 1960-64, omitting dummy variables, he again found a nonsignificant positive cross elasticity for demand deposits and savings and loan shares. His final recalculation of my original model extended over the period 1949-64, again confirmed a complementarity relationship between demand deposits and savings and loan shares [T1-6].

The other elasticity estimates appear to be remarkably robust given different specifications for dummy variables and different time periods. The major exception appears in the higher income elasticity estimated when the dummy variables are suppressed.

The fact that Lee's critique collapses under both theoretical scrutiny and econometric attempts at replication leaves the original evidence of weak substitution largely intact. Lee's critique of the use of regional dummy variables does give rise to the suggestion that rather than an "excessive use" of dummy variables, the prior studies might be criticized on the grounds of using too few dummy variables. If individual states, for example, had fixed but different intercepts, a richer specification would eliminate the implicit constraint that all these intercepts be equal. This richer specification is attempted in Section II.

I also attempted to estimate the demand function for demand deposits by taking account of the Slutsky condition which imposes a constraint on the substitution coefficients across demand equations. The restricted efficient estimates are reported for two subperiods as [T1-7] and [T1-8], and suggest at best a weak substitution relationship between demand deposits, time deposits, and savings and loan shares.

The most dramatic evidence of a strong substitution relationship is found in Lee's study [T1-9] of individual households from household survey data. Lee estimates a cross elasticity of −1.6 between demand deposits and savings and loan shares. While a higher cross elasticity is to be expected for household data, Lee's results seem highly implausible. They suggest that a one percentage point rise in the cost of demand deposits will reduce per capita demand deposits by $660 and raise the total of time deposits and savings and loan deposits by $951.[10] One problem with Lee's study may arise from the fact that he assigns regional interest rates to individual households. He claims that the regional rates for any single cross section "do not exhibit variations which are large enough to yield meaningful estimates for the purpose of statistical inference."[11] Thus the large estimated effect may result from temporal variations in the rates. In this case his evidence might better be considered along with other time series studies. Unfortunately a review of the extensive time series literature is beyond the scope of this study.

Kichline [9] has provided us with independent evidence [T1-10] which uses Standard Metropolitan Statistical Areas (SMSA) as the basic unit of analysis. His study uses announced interest rates rather than effective rates and he includes convenience cost variables which do not appear to have important effects. Kichline's estimates do not lend support to the substitution hypothesis.

The final study by Cohen [1] uses aggregate data rather then per capita data and does not include the rate of return on savings and loan association shares [T1-11]. Although his functional form is log-linear and his demand deposit holdings are reputedly for individuals, his evidence is broadly consistent with earlier studies, particularly when one notes that the rate of return on time deposits may also reflect some substitution between demand deposits and savings and loan shares.

When all the evidence presented in Table 1 is considered, one can conclude that, with one notable exception [T1-9], the evidence appears to be remarkably consistent. The greatest consistency is found in the estimate of the own rate elasticity for demand deposits. Considering all of the studies, the estimated range for the elasticity is .22 to .49. When Lee's household survey study [T1-9] is eliminated, the range is reduced to .22 to .39. In all studies the rate is highly significant and obviously robust over different specifications.

The evidence for substitutability between demand deposits and time deposits is also remarkably consistent. If one excludes the household survey [T1-9] study and the non-replicated study by Lee [T1-3], there are six studies reporting statistically significant but low elasticities ranging from —.10 to —.34. Since Cohen's study may also reflect substitution from savings and loan shares, the range may be narrowed to —.10 to —.21. The remaining three studies show statistically non-significant elasticities ranging from —.06 to .06. The evidence thus supports the contention of a low elasticity of substitution between demand deposits and time deposits.

The evidence for savings and loan shares is by far the most ambiguous.

10. See Friedman-Schwartz [4]. Another study by Mueller and Osborne [13] based on the same survey data concludes that "savers at large are not ready to be diverted from their present financial policies by moderate changes in yield patterns." Only one family in twelve responded affirmatively to a question asking whether changes in saving account interest rates influence what they do with their savings.

11. Lee [11], p. 449.

Putting aside Lee's household estimates and his non-replicated study, the range is from —.19 to +.32. Four studies indicate low and statistically non-significant elasticities while three studies show significant complementarity and one significant substitutability. With the exception of Lee's household study, however, the weight of the evidence is counterintuitive, suggesting little evidence for the substitution hypothesis and some confirming evidence for the surprising complementarity relationship.

The demand deposit substitution relationship with mutual savings bank deposits appears to be generally consistent with a weak substitution hypothesis. The complementarity result yielded by the restricted efficient estimates may stem from weight given to evidence from the demand functions for other liquid assets.

Finally, the income elasticities range from .80 to 1.46, but it appears that the suppression of dummy variables may be responsible for the higher income elasticity estimates. In order to examine this specific hypothesis and to get a clear test of the effects of a richer dummy variable specification, we turn to a covariance or error components model which allows for separate state and time effects.

II. Covariance and Error Components Specifications of the Demand for Demand Deposits

The review in the preceding section suggests the possibility that the relevant elasticity estimates may be sensitive to changes in the specification of the temporal cross-section model. In particular, it has become commonplace in temporal cross-section studies to incorporate dummy variables to account for time and state effects which are not readily attributable to identifiable causal variables.[12] For the case of estimating demand functions for demand deposits, it has already been argued that states which include major financial centers may attract a higher level of demand deposits than could be explained by the states' income or rates of returns. Similarly, state laws concerning mutual savings banking, branch banking, bank holding companies, etc., may have an independent effect on the quantity of demand deposits held in a particular state. Analogously, there may be temporal changes which have symmetrical effects on all geographic regions but differ over time. Thus, changes in federal regulations or policy and changes in general expectations may account for shifting levels of demand deposits over time. To the extent that such effects are present, a model which does not take account of them entails a specification error which can result in biased estimates of the remaining behavioral parameters. In order to test this hypothesis, we first estimate a demand function for demand deposits which suppresses all dummy effects.

The model estimated is similar to Hartley's study [T1-5]. The data base was updated to cover the period 1949-1965 and the functional form was assumed to be log-linear. The log-linear form which assumes constant elasticities rather than constant coefficients was chosen after noting the robustness of the elasticity estimates over different time period. Zarembka's independent

12. See Wallace and Hussain [18], Hoch [8], and Mundlak [14].

investigation of the functional form of the demand function for money also supports the log-linear formulation.[13] All other variables are analogous to those used in the earlier studies with the sole exception that the time deposit interest rate is taken as a weighted average of the rates on commercial bank time deposits and mutual savings bank deposits. The estimated elasticities are presented in Table 2 [T2-1]. Not surprisingly, the elasticity estimates are roughly consistent with those obtained in Section I for models excluding dummy variables. The cross elasticities are low and non-significant and the own elasticity is significant and falls within the expected range. As before, the income elasticity exceeds unity.

In order to allow for a richer specification which includes temporal and spacial effects, we turn to a covariance specification of the temporal cross-section model. Given N states and T time periods, the relationship connecting the observations on the dependent variable y_{nt}, with the K independent variables $x_{1nt} \ldots x_{Knt}$ is

$$y_{nt} = \alpha + U_n + V_t + \sum_{j=1}^{K} \beta_j X_{jnt} + \epsilon_{nt} \qquad (n = 1 \ldots N;\ t = 1 \ldots T) \tag{2.1}$$

where ϵ_{nt} is a random error with mean zero and variance σ^2. Letting

$$Y = \begin{bmatrix} y_{11} \\ \vdots \\ y_{1T} \\ \vdots \\ y_{N1} \\ \vdots \\ y_{NT} \end{bmatrix}, \alpha = \begin{bmatrix} \alpha \\ \vdots \\ \alpha \\ \vdots \\ \alpha \\ \vdots \\ \alpha \end{bmatrix}, U + V = \begin{bmatrix} U_1 + V_1 \\ \vdots \\ U_1 + V_T \\ \vdots \\ U_N + V_1 \\ \vdots \\ U_N + V_T \end{bmatrix},$$

$$X = \begin{bmatrix} X_{111} & \cdots & X_{K11} \\ \vdots & & \vdots \\ X_{11T} & \cdots & X_{K1T} \\ \vdots & & \vdots \\ X_{1N1} & \cdots & X_{KN1} \\ \vdots & & \vdots \\ X_{1NT} & \cdots & X_{KNT} \end{bmatrix}, \beta = \begin{bmatrix} \beta_1 \\ \vdots \\ \beta_K \end{bmatrix}, \epsilon = \begin{bmatrix} \epsilon_{11} \\ \vdots \\ \epsilon_{1T} \\ \vdots \\ \epsilon_{N1} \\ \vdots \\ \epsilon_{NT} \end{bmatrix}$$

be a $TN \times 1$ vector, $TN \times 1$ vector, $TN \times 2$ matrix, $TN \times K$ matrix, $K \times 1$ vector and $TN \times 1$ vector respectively, we have,

$$Y = \alpha + U + V + X\beta + \epsilon. \tag{2.2}$$

13. See Zarembka [19].

TABLE 2

Covariance Model Estimates of the Demand for Demand Deposits

Reference Number	Author	Estimation Procedure	Time Period	Functional Form	Unit of Observation	Dummy Variable Specification	Elasticity Estimates			
							$\eta_{r_{dd}}$	$\eta_{r_{dtm}}$	$\eta_{r_{ds}}$	η_y
[T2-1]	Feige	OLS	1949-65	Log-linear	State	ND	.36*	−.04	−.01	1.20*
[T2-2]	Feige	OLS	1949-65	Log-linear	State	T + SD	.20*	.03*	−.18*	.71*

The first model estimated in this section is simply (2.2) with the restriction that $U = V = 0$. The richer unconstrained model represented by (2.2) can be viewed as a covariance model which requires estimation of $N + T$ state and time intercepts. Alternately, if U and V are regarded as random errors, (2.2) can be viewed as an error components model requiring the estimation of the means and variances of the time and state error components.[14] Since the major focus of our interest is to estimate the β coefficients, it is computationally possible to "sweep out" the state and time intercept terms by means of a covariance transformation. Letting I_T, I_{NT}, P_T, and P_{NT} be $T \times T$ and $NT \times NT$ identity matrices and $T \times T$ and $NT \times NT$ matrices of ones, respectively, and defining

$$A = \begin{bmatrix} I_T^{11} & \cdots & I_T^{1N} \\ \vdots & \ddots & \vdots \\ I_T^{N1} & \cdots & I_T^{NN} \end{bmatrix}, \qquad B = \begin{bmatrix} P_T^{11} & & 0 \\ & \ddots & \\ 0 & & P_T^{NN} \end{bmatrix},$$

the transformation operator,

$$Q = \left[I_{NT} - \frac{1}{N} A - \frac{1}{T} B + \frac{1}{NT} P_{NT} \right]$$

can be applied to (2.2) to "sweep out" the intercept terms since

$$QY = Q\alpha + Q(U + V) + QX\beta + Q\epsilon$$

reduces to

$$QY = QX\beta + Q\epsilon. \tag{2.3}$$

The effect of the covariance transformation operator is to transform the variables of the analysis into deviations from state and time means plus the grand mean. Ordinary least squares estimators applied to the transformed model (2.3) are unbiased and efficient. If (2.2) is viewed as an error components model, Wallace and Hussain [18] and Henderson [7] have shown that ordinary least squares estimators are unbiased and asymptotically as efficient as the generalized least squares estimators when the state and time effects dominate the residual error component. Since we have no a priori reason for choosing between the fixed intercept covariance model and the error components interpretation, we assume the covariance specification and estimate the appropriately transformed equation system by ordinary least squares.

The results of the covariance model estimates are presented in Table 2 as [T2-2]. As expected, the income elasticity is lower than that found in the more restrictive prior studies. The own elasticity is significant but lower than in previous estimates and the cross elasticity for time and mutual deposits is low and insignificant. The cross elasticity for savings and loan associations is −.18 and significant.

14. See Maddala [12].

In order to test the appropriateness of the covariance specification against the more restrictive model estimated as [T2-1], we can test the null hypothesis that $U = V = 0$. The hypothesis can be tested by

$$F = \frac{(e^{*\prime}e^{*} - e'e)/T + N - 2}{e'e/TN - T - N - 3}$$

where $e^{*\prime}e^{*}$ and $e'e$ are respectively the residual sum of squares for the constrained and unconstrained regressions. The F-statistic is distributed as F with $T + N - 2$ and $TN - T - N - 3$ degrees of freedom. The resulting F-statistic of 258.22 is highly significant, rejecting the hypothesis of zero time and state intercepts. The test suggests that Lee's [11] concern over an "excessive use of dummy variables" is entirely unfounded, and that an appropriate temporal cross-section model must allow for separate state and time intercepts. Interestingly enough, however, the results of the covariance model are broadly consistent with earlier estimates.

III. A Random Coefficient Specification of the Demand Function for Demand Deposits

While the covariance model presented in the preceding section is a considerably less restrictive specification than that of earlier models, it still assumes homoscedastic disturbances across states and identical slope coefficients over time and states. Feige [3] tested the stability of slope coefficients over time, but no comparable test has been performed to test stability of the coefficients over states. Updating of the temporal cross-section data to include a larger number of years makes this test possible.

The concern of this section is the relaxation of the assumptions of homoscedastic disturbances and fixed slope and intercept parameters over states.

Specification of heteroscedastic disturbances would be appropriate if different states were differentially affected by random temporal disturbances. For example, changes in monetary policy may result in larger shifts in deposits in states with sophisticated financial markets. In order to take account of this type of effect, consider a model for the n^{th} state as,

$$y_n = X_n\beta_n + \epsilon_n, \tag{3.1}$$

where y_n is a $T \times 1$ vector of dependent observations, X_n is a $T \times (K + 1)$ matrix of observations on the n^{th} state, β_n is a $(K + 1) \times 1$ vector of coefficients and ϵ_n is a $T \times 1$ vector of random disturbances, with each element distributed with mean zero and variance σ_n^2. For simplicity, we shall no longer differentiate between the intercept and coefficient vector since they are treated symmetrically for the remainder of the paper. Thus, the first column of the observation matrix has unit elements, and the first element in the β vector is the state intercept. When the N states are pooled, and we impose the restriction,

$$\beta_n = \beta \qquad \text{for } n = 1 \ldots N, \tag{3.2}$$

the model is identical to the constrained model estimated in the preceding section [T2-1] except that it allows for heteroscedastic disturbances.

This constrained model

$$y_n = X_n\beta + \epsilon_n \quad (3.3)$$

can be estimated by means of the Aitken-Zellner generalized least squares estimator, and the resulting elasticities are presented in Table 3 as [T3-1]. The estimated elasticities correspond closely with the estimates obtained by ordinary least squares for the homoscedastic model [T2-1].

Having allowed for heteroscedastic disturbances, we wish to examine the hypothesis that behavior is consistent across states, that is, whether (3.2) is a valid restriction. Zellner [20] has shown that the estimated parameters ($\hat{\beta}_n$) and error variances (s_n^2), derived from the N state models (3.1), can be combined with the generalized least squares estimates ($\hat{\beta}$) of the pooled model (3.3), so that under hypothesis (3.2), the asymptotic distribution of

$$F = \frac{1}{(N-1)(K+1)} \sum_{n=1}^{N} \left[(\hat{\beta}_n - \beta)' \frac{X'_n X_n}{S_n^2} (\hat{\beta}_n - \beta) \right]$$

is the F distribution with $(N-1)(K+1)$ and $(T-K-1)N$ degrees of freedom. The value of the calculated F statistic was 208.61 which is well above the 5% value of F with 240 and 588 degrees of freedom. We therefore must reject the hypothesis of fixed and equal intercepts and slopes across states.

Rejection of hypothesis (3.2) suggests the specification of a less restrictive model which makes explicit allowance for inter-state variation. Swamy [16] has developed estimation procedures for a class of random coefficient models whose specification permits bounded variation across states. In particular, the random coefficient specification replaces the assumption (3.2) of fixed state intercepts and slopes, with the assumption that the β_n's in (3.1) are independently and identically distributed with

$$\begin{aligned} E\beta_n &= \beta \\ E(\beta_n - \beta)(\beta_m - \beta)' &= \Delta \qquad \text{for } n = m \\ &= 0 \qquad \text{for } n \neq m \end{aligned} \quad (3.4)$$

where Δ is a $(K+1) \times (K+1)$ variance-covariance matrix. Retaining the assumption of heteroscedastic error disturbances, the best linear unbiased estimator of β given the specifications in (3.1) and (3.4) is

$$\hat{\beta} = \sum_{n=1}^{N} W^*_n \hat{\beta}_n \quad (3.5)$$

where

$$W^*_n = \left[\sum_{n=1}^{N} \{\Delta + s_m^2 (X'_m X_m)^{-1}\}^{-1} \right]^{-1} \{\Delta + s_n^2 (X'_n X_n)^{-1}\}^{-1}$$

and $\hat{\beta}_n$ and s_n^2 are the ordinary least squares estimates of the state parameters and error variances. Swamy has demonstrated that an unbiased estimator of Δ is

TABLE 3
RANDOM COEFFICIENT MODEL ESTIMATES

Reference Number	Author	Estimation Procedure	Time Period	Functional Form	Unit of Observation	Dummy Variable Specification	Elasticity Estimates			
							$\eta_{r_{dd}}$	$\eta_{r_{dtm}}$	$\eta_{r_{ds}}$	η_y
[T3-1]	Feige	GLS	1949-65	Log-linear	State	ND	.37*	−.07*	.08*	1.27*
[T3-2]	Feige	WGLS	1949-65	Log-linear	State	Random Coefficients	.09*	−.02	−.21*	.62*

$$\hat{\Delta} = \frac{\sum_{n=1}^{N} \hat{\beta}_n \hat{\beta}'_n - \frac{1}{N} \sum_{n=1}^{N} \hat{\beta}_n \sum_{n=1}^{N} \hat{\beta}'_n}{N-1} - \frac{1}{N} \sum_{n=1}^{N} s_n^2 (X'_n X_n)^{-1}. \tag{3.6}$$

The random coefficient model was estimated as outlined above and the results are reported in Table 3 as [T3-2]. The own elasticity is significant with the expected sign but considerably below the range found in other studies. The cross elasticity for time and mutual savings deposits conforms with that found in earlier studies and the savings and loan cross elasticity indicates a statistically significant elasticity of —.21. The income elasticity is significant but somewhat lower than that found in earlier studies.

The results for the random coefficient model must, however, be approached with some caution. The estimated variance-covariance matrix $\hat{\Delta}$ is formed as the difference between two matrices as shown in (3.6). While one would theoretically expect Δ to be positive semi-definite, Swamy has pointed out that "in some numerical applications, $\hat{\Delta}$ will yield negative estimates for the variances of some coefficients."[15] In our calculations, the variance of the time deposits elasticity was found to be negative. This finding suggests that the data did not entirely meet the specifications and assumptions of the random coefficient model and thus the results can not be regarded as conclusive. The estimates do, however, underline some of the findings based on the earlier covariance model.

Unfortunately, no formal test exists which allows a direct comparison between the appropriateness of the covariance specification vis-à-vis the random coefficient specification. The random coefficient model is less restrictive with respect to state slope specification. The common element of both models is the allowance of state intercept variation. Since the results are qualitatively similar, one might hazard the presumption that allowance for separate state intercepts is the critical specification requirement. When all the studies are reviewed, however, the only systematic effect of the inclusion of separate state intercept terms is to reduce the estimated income elasticity. No such systematic effects are observed for the interest rate elasticities.

IV. Summary and Conclusions

We have reviewed the evidence on the demand function for demand deposits based on temporal cross-section data and have considered a rather rich menu of alternative model specifications. Tests of alternative model specifications led to the conclusion that an appropriate temporal cross-section specification must allow for the inclusion of separate state effects and possibly separate temporal effects.

Considering the differences in data bases, model specifications and estimation procedures reviewed, one can not help but be struck with the overall robustness of the parameter estimates. With one notable and troublesome [T1-9] exception, the studies seem to support the following generalizations.

15. Swamy [16], p. 170.

(1) The rate of return on demand deposits is an important argument of the demand function and the own elasticity appears to fall in the range .20 to .40.

(2) Demand deposits and commercial bank time deposits appear to be weak substitutes with a cross elasticity range of −.20 to 0.

(3) Demand deposits and mutual savings bank deposits appear to be weak substitutes; however, one can not exclude the possibility of a weak complementarity relationship. The range of the cross elasticity estimates is from −.05 to +.15.

(4) The relationship between demand deposits and savings and loan deposits remains somewhat ambiguous. At best the two assets exhibit weak substitutability; however, a complementarity relationship can not be entirely ruled out. The estimated range for the cross elasticity is −.20 to +.30. The evidence from both the covariance models and the random coefficient model are, however, consistent with the weak substitution hypothesis.

(5) The income elasticity of demand deposits for the post war period appears to be less than unity. There is a strong systematic relationship between model specifications with respect to dummy variables and the estimated income elasticity. The more appropriately specified models which include separate spacial effects yielded income elasticities ranging between .65 and .95.

Since the evidence presented is limited to demand deposit estimates from the class of temporal cross-section models, the need remains for a systematic investigation of the consistency of the foregoing results with evidence culled from time series studies and estimates of the demand functions for other liquid assets.

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