Managerial Objectives in Regulated Industries: Expense-Preference Behavior in Banking

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Recent work on the theory of the firm under regulation suggests that managers of regulated firms may be utility maximizers rather than profit maximizers. There is, however, very little empirical evidence on managerial behavior in regulated industries. This article examines one kind of utility-maximizing behavior that seems particularly applicable to regulated firms: expense-preference behavior. Specifically, I develop a test capable of discriminating between expense-preference and profit-maximizing behavior and apply it to the banking industry, a highly regulated industry. My findings indicate that an expense-preference theoretical framework better explains the behavior of regulated firms than does a profit-maximization framework.

I. Introduction

Within recent years economists have begun to view firms as entities which maximize utility rather than profits (Becker 1957; Williamson 1963). One approach which has received considerable attention is the "expense-preference theory." This theory envisages the firm as maximizing utility through the pursuit of non-profit-maximizing policies. In particular, managers increase (beyond the profit-maximizing point) "staff expenditures, managerial emoluments, and discretionary profit," for which they have a positive preference (Williamson 1963; Rees 1974).[1] In its

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[1] In a thorough review of the expense-preference theory, Rees suggests that "the main innovation of the . . . theory lies in its specification of the objectives of the firm." He says, "The precise way in which this concept is made operational by the definition of staff expenditures, managerial emoluments and discretionary profit as types of expenses for which managers have positive preferences, permits a richer specification of the firm's utility function than those made by, say Baumol or Marris, . . ." (Rees 1974, pp. 295–96).

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narrow operational form, therefore, the expense-preference theory posits that firms will hire more staff and/or pay higher managerial wages than will profit-maximizing firms, everything else equal. This theory, like most postclassical theories of the firm, is premised on assumptions that permit firms to pursue non-profit-maximizing policies: separation of ownership from control and imperfections in either the goods or capital markets or both (Williamson 1963; Olsen 1973). Under these conditions, the "transaction costs" that stockholders must incur to exercise a significant degree of control over managers are too great to justify such efforts in all but the most extreme cases of managerial "misbehavior."

It is also widely recognized that a "discretionary" view of managerial behavior is in principle particularly appropriate to regulated industries (Alchian and Kessel 1962; Williamson 1963), where the abridgment of property rights is especially acute. There is, however, no empirical evidence of a rigorous nature to support such a view. The primary purpose of this paper is to fill this void by empirically examining managerial behavior in a highly regulated industry—banking.

Banking is an ideal industry in which to examine the issue. First, there are obvious imperfections in banking markets: monopoly and oligopoly are prevalent (Edwards 1964; Phillips 1967; Jacobs 1971); key interest rates are fixed by law (regulation Q); entry is strictly controlled (Peltzman 1965; Edwards and Edwards 1974); and there are substantial regulatory obstacles to branching, merging, etc. Second, separation of ownership from control is common among large banks (Vernon 1970, 1971). Thus, the market conditions necessary to sustain non-profit-maximizing managerial behavior are clearly present.

Past studies have also unearthed a number of findings which indirectly (and inadvertently) suggest that expense-preference behavior may be common in banking. To begin with, these studies find a small, positive relationship between monopoly power (market concentration) and bank prices or interest rates (Edwards 1964; Phillips 1967; Jacobs 1971) but find no systematic relationship between monopoly power and bank profitability. There are, of course, a number of conceivable explanations for this finding, such as that reported profits are not measured in a consistent and accurate fashion. An explanation that has not been advanced, however, is that the managements of (monopolistic) banks appropriate part of the profits through expense-preference behavior: by paying themselves higher salaries, hiring excessive staff, or being lax in their personnel supervision, all of which may enhance their utility (and increase costs) while lowering reported profitability.

In addition, other studies find that increased competition does not diminish (reported) bank profitability; rather, competition improves bank efficiency or lowers operating costs (Motter and Carson 1964; Kohn and Carlo 1969; Fraser and Rose 1972; Federal Reserve Bank of New York

1973). For example, in a recent study of new entry into one-, two-, and three-bank towns, Fraser and Rose found that new bank formation is not related to reported profitability; rather, new banks are formed because existing banks are inefficient and/or too conservative (Fraser and Rose 1972, pp. 69–70). Banks in markets where new banks entered tended to have higher staff operating costs (or higher ratios of salaries and wages-to-total operating earnings) than the norm, and these costs returned to normal subsequent to entry (Fraser and Rose 1972, pp. 72–75). In still another study, the Federal Reserve Bank of New York found that an enhanced competitive environment resulted in "relatively slower growth in such (non-interest) expense categories as salaries and wages, pensions, and other (non-interest) expenses" (Federal Reserve Bank of New York 1973, pp. 70–71). Again, increased competition had no effect on reported profitability. Thus, by inadvertently discovering a relationship between the degree of competition and bank operating efficiency, past studies of bank performance suggest the presence of expense-preference behavior (as I demonstrate in Section III).

In sum, both the institutional characteristics of banking markets and the above unexplained phenomena of bank performance imply that expense preference may be a useful theoretical framework within which to analyze managerial behavior in banking. This theory, in contrast to profit maximization, implies that a reduction in the degree of competition will increase firm costs, since managers can be expected to indulge their preferences for greater staff and higher salaries as "true" profits increase (Williamson 1963, pp. 1041–43).[2] In this paper I investigate the relationship between a bank's labor expenses and its monopoly power in order to ascertain whether banks are profit-maximizing or expense-preference firms. Both excessive staff expenditures and managerial emoluments will be reflected in a firm's total labor expenses.

II. Alternative Models of Expense-Preference Behavior

Economists have approached the theory of expense preference in two ways. First, it has been hypothesized that managers treat as distinct decisions the maximization of "true" profits and the maximization of managerial utility (hereafter referred to as the "severable" decision model). Second, managers have been viewed as treating the profit and utility decisions as a single, joint decision to maximize utility (hereafter referred to as the "joint" decision model).[3] The severable decision model

[2] Using concentration ratios as indexes of monopoly power, Williamson (1963) argues that "an increase in the concentration ratio will tend to widen the opportunities for managerial discretion" (p. 1043).

[3] The first approach I would associate with Alchian and Kessel (1962) and the second with Williamson (1963) and Rees (1974).

envisages the firm as first making all the usual choices necessary to maximize profits and then internally allocating the resulting profits between reported profits and utility-enhancing expenses, such as excess staff expenditures and managerial emoluments (Alchian and Kessel 1962, pp. 162–66). Here the firm's output, prices, advertising, and nonlabor inputs are all at their profit-maximizing level, while labor expenditures are above the profit-maximizing level. This behavior results in "pure waste," since the firm is viewed as operating inside its production-efficiency frontier.

Alternatively, the joint decision model envisages the firm as choosing an entirely different production strategy: expanding output beyond the profit-maximization level and utilizing a ratio of labor to other inputs greater than that implied by profit-maximization behavior. In this case the firm operates on its production-efficiency frontier (rather than inside it), but at the "wrong" place (Rees 1974, pp. 298–302).[4] Both theories, however, predict that an expense-preference firm's labor expenditures will exceed what they would be if the firm were solely a profit maximizer.

III. A Test of Expense Preference: The Demand for Labor by an Expense-Preference versus a Profit-maximizing Firm

An empirical test of expense-preference behavior based upon a firm's demand for labor can be developed in the following way: First, given a pure profit-maximizing monopoly firm with a two-factor Cobb-Douglas production function and facing a constant elasticity industry (or market) demand schedule, it is a simple matter to show that the firm's derived demand for labor will be (natural logs = ln)

$$\ln L = b_0 + b_2 \ln W + b_3 \ln r + b_4 \ln Y + \ln \hat{\eta}, \qquad (1)$$

where L is the number of employees; W is the wage rate; r is the cost of capital; Y is a shift variable representing all nonprice factors that affect the demand for the final product; and b_0, b_1, b_2, b_3, b_4, and $\ln \hat{\eta}$ are the usual constant parameters.[5] In particular, $\ln \hat{\eta} = (\ln \eta)/(1 - b_1)$, where $\eta = 1 + (1/e)$, and where e is the elasticity of the firm's final demand schedule (or, in this case, the industry elasticity), and $\ln \eta < 0$, since a monopolist will operate in the elastic portion of the demand curve.

Second, a similar but not identical aggregate demand-for-labor function can be derived for firms in a competitive industry if two additional

[4] Rees treats this case as one in which the firm minimizes costs but with respect to the "wrong" relative factor prices, or where the shadow price for staff is below the actual price. In terms of an isoquant diagram, the firm is on an isoquant but off the expansion path. In the first case, the firm is inside the efficiency-frontier isoquant.

[5] For a detailed derivation of this equation and the others which follow, see the Appendix.

assumptions are made: that firms have constant returns to scale,[6] and that the size of the firm is determined by factors exogenous to the firm.[7] The resulting aggregate labor-demand equation will be

$$\ln L = b_0 + b_2 \ln W + b_3 \ln r + b_4 \ln Y. \tag{2}$$

This equation is identical to equation (1) above except for the term $\ln \hat{\eta}$, which appears in the monopolistic industry's demand equation (1) but not in the competitive industry's demand equation (2).

The term $\ln \hat{\eta}$, therefore, captures the distinction between the (profit-maximizing) competitive and monopolistic industry labor demands. Specifically, this term equals $(\ln \eta)/(1 - b_1)$ and is negative in value for a monopolistic firm or industry, since $\eta < 1$ and $(1 - b_1) > 0$ (assuming constant returns to scale), while for a competitive industry the term equals zero, since $\eta = 1$. Thus, monopolistic industries will demand less labor, everything else equal. Alternatively, in any empirical estimation of a labor-demand function across both competitive and monopolistic industries or markets, the only parameter that will vary because of differences in industry structure is the constant term $(b_0 + \ln \hat{\eta})$, which should be smaller for a monopolistic than for a competitive industry.

How will the demand-for-labor equation change if the firm is an expense-preference rather than a profit-maximizing firm? The severable decision model is the easiest to handle. This model envisions the firm as first computing the profit-maximizing labor force and then simply hiring $(1 + z)$ times that amount, where z is a constant and >0. Thus, if actual employment is $L^* = (1 + z)L$, then $\ln L^* = \ln (1 + z) + \ln L$, and equation (1) becomes

$$\ln L^* = \ln (1 + z) + b_0 + b_2 \ln W + b_3 \ln r + b_4 \ln Y + \ln \hat{\eta}. \tag{3}$$

The difference between the expense-preference and profit-maximization firm, therefore, is the term $\ln (1 + z)$. In a competitive industry, $z = 0$ and $\eta = 1$, so the intercept of equation (3) is simply b_0, as in equation (2). For a monopoly expense-preference firm, on the other hand, the intercept equals $[b_0 + \ln (1 + z) + \ln \hat{\eta}]$, which is different from the intercept of equation (1).

Further, since $\ln \hat{\eta} < 0$ but $\ln (1 + z) > 0$, this intercept may be greater or less than b_0, the intercept for a competitive industry. If

[6] The assumption of constant returns to scale appears reasonable for many industries, and for banking in particular. Studies of economies of scale in banking have found only very small scale economies, and then only over small-bank size classes (at most, up to an asset size of \$50 million). In addition, the empirical work in this paper utilizes data from banks in metropolitan areas, where banks are usually large enough to have exhausted whatever economies exist. For evidence on economies of scale in banking, see Bell and Murphy (1968, chap. 4).

[7] Alternatively, the equivalent assumption that the number of firms in the industry is exogenous can be assumed. See Appendix.

$[b_0 + \ln(1+z) + \ln \hat{\eta}] > b_0$, it follows that $\ln(1+z) > 0$ and the firm must be an expense-preference firm. However $[b_0 + \ln(1+z) + \ln \hat{\eta}] \leq b_0$ does not rule out expense-preference behavior, since the negative $\ln \hat{\eta}$ effect may offset or outweigh the positive $\ln(1+z)$ effect. Thus, an intercept greater than b_0 is consistent only with expense-preference behavior, while an intercept equal to or less than b_0 is consistent with either expense-preference or profit-maximization behavior.

The more complex view of expense-preference behavior—the joint decision model—yields, surprisingly, a derived demand for labor identical in all essential respects to equation (3). In this model, which envisions the firm as maximizing the utility function $U = U(\pi, E)$, where π is the firm's profits and E is its total labor expenses, the expense-preference firm's demand-for-labor equation becomes

$$\ln L = b_0 + b_2 \ln(1 - U_E/U_\pi) + b_2 \ln W + b_3 \ln r + b_4 \ln Y + \ln \hat{\eta}, \quad (4)$$

where U_E is the marginal utility of a dollar of additional labor expenses and U_π is the marginal utility of another dollar of reported profits.[8] The only difference between equations (4) and (3) is that the term $[b_2 \ln(1 - U_E/U_\pi)]$ in (4) replaces $\ln(1+z)$ in (3). In addition, as before, for a competitive industry the intercept term in equation (4) is equal to b_0, since $U_E = 0$ and $\eta = 1$. For a monopoly firm, however, the intercept is now $[b_0 + b_2 \ln(1 - U_E/U_\pi) + \ln \hat{\eta}]$. Since $b_2 < 0$ and $\ln(1 - U_E/U_\pi) < 0$, $[b_2 \ln(1 - U_E/U_\pi)] > 0$. Thus, this intercept term may be $\gtreqless b_0$, depending upon whether $|b_2 \ln(1 - U_E/U_\pi)| \gtreqless |\ln \hat{\eta}|$. Once again, therefore, if the intercept of equation (4) exceeds b_0, it can only be because $[b_2 \ln(1 - U_E/U_\pi)] > 0$, which clearly implies expense-preference behavior. Profit-maximizing behavior would never yield an intercept greater than b_0.[9] If the intercept $\leq b_0$, expense-preference behavior may or may not exist.

In sum, by estimating the usual reduced-form demand for labor for a sample of both competitive and monopolistic markets, it may be possible to detect the existence of expense-preference behavior. Specifically, if we make the usual assumption that concentration ratios constitute reasonable indexes of monopoly power, and that the concentration-monopoly relationship is discrete rather than continuous (or that some threshold level of concentration must be reached before firms recognize their mutual interdependence and begin to coordinate their activities

[8] See Appendix, equations (17)–(19), for the precise derivation of equation (4).

[9] If the expense-preference firm's behavior is described by some combination of the "severable" and "joint" models, its demand-for-labor function will be a combination of equations (3) and (4), and once again an intercept $> b_0$ will imply expense-preference behavior.

either implicitly or explicitly),[10] equation (4) can be modified by the inclusion of a dummy variable to represent the difference in the constant term due to industry structure:

$$\ln L = b_0 + b_2 \ln W + b_3 \ln r + b_4 \ln Y + hM. \tag{5}$$

The M is a dichotomous variable which takes the value one when concentration is above the required threshold level and zero otherwise, and h is the difference in the constant term resulting from the difference in market structure. Given a pure profit-maximizing behavioral assumption, we would expect the coefficient of M (or h) to be less than zero, reflecting the lower price elasticity of demand which monopolistic firms face. On the other hand, if h is positive, this clearly implies expense-preference behavior. If h were not to differ significantly from zero, this might be interpreted as weak confirmation of the expense-preference hypothesis, since profit maximization necessarily implies $h < 0$. (In other words, the positive expense-preference effect on employment may not be large enough to dominate the negative effect that a lower price elasticity of demand has on the firm's demand for labor.)

Finally, since the expense-preference theory focuses on both excess staff expenditures and managerial emoluments, two versions of equation (5) are estimated: one using total labor employed (L) as the dependent variable, and one using total wage and salary expenses (TW) as the dependent variable. Theoretically, of course, the latter equation can be derived from the first by simply multiplying through by the wage rate.

IV. Tests and Empirical Findings

Equation (5) was estimated for a 3-year (1962, 1964, 1966) sample of 44 metropolitan areas (SMSAs), over which the three-firm concentration ratio varies from 39 to 94 percent.[11] Separate regression equations were estimated for each year and for all years pooled together, although only the pooled results are reported in table 1.[12] The results indicate that the expense-preference hypothesis cannot be rejected.

[10] The use of a discrete concentration-monopoly relationship is common in industrial organization studies. See, for example, Chamberlin (1945, pp. 46–51), Fellner (1949, chap. 4), Bain (1951), and Weiss (1974, p. 371). In addition, I explicitly test the validity of this assumption later in the paper. See n. 17 below.

[11] Specifically, I used aggregate banking data for 43 SMSAs in 1962, 43 SMSAs in 1964, and 44 SMSAs in 1966. These SMSAs were those for which complete data were available for each of the 3 years. The major data sources were: U.S. Bureau of the Census (1962, 1964, 1966, 1968), U.S. Department of Labor (1963, 1965, 1967), and Federal Deposit Insurance Corporation (1966, 1970).

[12] The null hypothesis of structural stability over the 3 years could not be rejected when I estimated separate equations for each year using the optimal "switching level" of M (see n. 17 below). In addition, the coefficient of M ranges from a low of .139 (with a t of 1.63) in 1964 to a high of .294 (with a t of 2.77) in 1962.

TABLE 1

REGRESSION EQUATIONS FOR ALL YEARS POOLED

Equation No.	Dependent Variable	M	Y	W	B	LD	D	BO	Constant	Corrected R^2	df
All observations:											
1.	TW	0.229	1.117	−0.604	−0.093	. . .	. . .	. . .	1.558	.94	125
		(4.52)	(40.59)	(3.32)	(1.51)				(6.80)		
2.	TW	0.216	1.115	−0.644	−0.092	0.053	−0.036	. . .	1.381	.94	123
		(4.06)	(40.18)	(3.43)	(1.47)	(0.86)	(0.35)		(4.67)		
3.	L	0.199	1.072	−0.803	−0.086	. . .	. . .	. . .	0.552	.94	125
		(4.38)	(43.24)	(4.90)	(1.55)				(2.68)		
4.	L	0.191	1.071	−0.826	−0.087	0.216	−0.032	. . .	0.473	.94	123
		(3.99)	(42.72)	(4.87)	(1.54)	(0.39)	(0.35)		(1.77)		
5.	TW	0.199	1.098	−0.627	−0.186	0.058	−0.027	0.059	1.482	.94	122
		(3.54)	(33.14)	(3.32)	(1.54)	(0.95)	(0.27)	(0.91)	(4.68)		
6.	L	0.175	1.054	−0.810	−0.182	0.027	−0.024	0.060	0.577	.94	122
		(3.44)	(35.24)	(4.75)	(1.62)	(0.49)	(0.26)	(1.02)	(2.02)		
Without unit-banking observations:*											
7.	TW	0.215	1.170	−0.690	0.20	0.052	−0.080	−0.091	0.949	.95	98
		(3.99)	(32.26)	(3.61)	(3.57)	(0.90)	(0.814)	(1.34)	(3.23)		
8.	L	0.189	1.117	−0.854	0.176	0.024	−0.074	−0.074	0.071	.96	98
		(3.94)	(34.63)	(5.00)	(3.52)	(0.46)	(0.85)	(1.23)	(0.27)		

NOTE.—t-ratios are in parentheses beneath the estimated coefficients; in all equations, $M = 1$ if concentration ≥ 76 percent and 0 if concentration ≤ 76 percent.

* Since all observations in this sample are either limited branching or statewide branching cities, variable B is redefined in equations (7) and (8) to take the value of 1 if the city is in a statewide branching state, and 0 otherwise.

Estimation of equation (5) encountered several difficulties. First, and most familiar, it was not always easy to obtain precise measures of the variables. I used as a measure of the demand (or scale) variable (Y), total personal income in the respective SMSA, and I assumed that the cost of capital (r) was identical for all banks, a common assumption in banking studies.[13] The most difficult problem was how to measure the wage rate (W). A bank-specific measure of the wage rate could be "contaminated" by expense-preference behavior, since monopolistic banks might be paying excessive salaries. Further, the only wage data in banking available are for total wages and salaries paid (TW) and for total employees (L), so that the only bank-specific wage rate available is one constructed by dividing TW by L. This variable is unsatisfactory because it confuses hours worked and labor quality with wage rates. Thus, I opted to use the hourly wage rate in manufacturing in the respective SMSA as an exogenous index of bank wage rates. Bank wage rates, uncontaminated by the expense preferences of bank managers, should vary directly with manufacturing wage rates, since within an SMSA we would expect there to be labor mobility among occupations.[14]

The second problem was the homogeneity assumptions implicit in equation (5). In particular, the nature of the banking industry required the inclusion of three additional independent variables in equation (5) in order to account for differences in "branching" and in "output mix." Differences in branching law are important for at least two reasons. First, the organizational technology of operating many offices is different from that of operating only one office—that is, more labor is used by a branch bank of a given size than by a unit bank (Bell and Murphy 1968, pp. 168–69). Second, prior research on banking structure suggests that state branching restrictions may be an important barrier to entry: the more restrictive a state's branching law, the greater are the barriers to entry into banking markets within the state and the less competitive are these markets (Shull and Horvitz 1964; Jacobs 1971). Thus, the expense-preference hypothesis implies that labor expenditures will be lower in branch banking markets, since these are the more competitive markets. Banks located in liberal branching-law states (e.g., statewide branching), therefore, can be expected to employ more labor because of their multi-office technology, but less labor because they operate in a more competitive environment (given expense-preference behavior). For these reasons alone, it is impossible to predict a priori the relationship between

[13] See Bell and Murphy (1968, p. 13). They argue that the market for bank capital is nationwide and not regional. In addition, banking regulation arguably prevents differences in r among banks.

[14] The above notwithstanding, when I constructed an SMSA bank-specific wage rate (by dividing TW by L as defined in table 1) and used it in equation (5), it was not significant. In addition, its use did not change any of the other estimated coefficients. Thus, my results are insensitive to the wage rate used.

the type of branching law and the amount of labor an expense-preference bank will employ.

This ambiguous relationship is further confounded by the possibility that the degree of branching may itself reflect the objective function of bank management. Expense-preference managers may use branching as a strategy to increase utility-enhancing expenditures on staff. In particular, the pervasiveness of branching in a market may be positively related to the degree of concentration in the market, since we would expect to find expense-preference behavior in concentrated markets.[15] Consequently, it is inappropriate to include in equation (5) a variable measuring the actual extent of branching. (Nevertheless, the inclusion of such a variable does not alter my findings. See eqq. [5]–[8] in table 1.)

Banks may also have different output mixes, which may account for differences in the amount of labor they utilize. Consequently, two "output homogeneity" variables were included in equation (5): average loan-to-deposit ratios (*LD*) and average demand-to-savings deposit ratios (*D*) in the SMSA. These variables obviously standardize for only the most obvious possibilities. Neither adds to the explanatory power of equation (5).[16]

Finally, in order to estimate equation (5) it was necessary to determine the critical (or threshold) level of concentration, or the level at which banks coordinate their activities, either explicitly or implicitly. Since there is no way of knowing this level a priori, Quandt's "switching of regimes" technique was used to determine it. This procedure shows the critical level to be a three-firm concentration ratio of 76 percent.[17] Consequently, M is set equal to one when concentration equals or exceeds 76 percent, and zero when concentration is less than 76 percent.[18]

[15] In a recent study, Larry White finds a significant negative relationship between branching and concentration, using a profit-maximizing, non-price-competition model. His findings imply that, "Banks in a more competitive environment will appear to have higher costs than other banks in less competitive environments, . . ." (1976, p. 104).

[16] I also tried "average size of bank" to standardize for output mix with the same results. All of these variables may, of course, be contaminated by expense-preference behavior. If so, they should reduce the significance of the coefficient of M in equation (5). We do not observe this result. In addition, it should be noted that the extremely high R^2s in table 1 suggest that we are not omitting important explanatory variables.

[17] See Quandt (1958, 1960), Goldfeld and Quandt (1972, 1973), and White (1974). Quandt shows that the optimal "switching level" for M can be determined by examination of the relevant t-statistics for M. Specifically, estimation of equations (2) and (4) in table 1 for all possible "switching levels" of M shows that the t-statistic for M is highest at the 76 percent level: 4.06 in equation (2) and 3.99 in equation (4). For all concentration levels less than 76 percent, the t-statistic for M never exceeds 2.84, while for all levels greater than 76 percent the t-statistic does not exceed 0.25. Finally, these results also support my earlier assumption of a discrete concentration-monopoly relationship.

[18] The average level of concentration in cities with less than 76 percent is 63.6 percent, and in cities which have ≥ 76 percent the average is 84.3 percent.

In sum, alternative pooled versions of equation (5) are estimated for the years 1962, 1964, and 1966 and reported in table 1. The variables are defined as follows:

TW = natural logarithm of total wages and salaries paid by all banks located in an SMSA, measured as four times the wage payments in the SMSA during the first quarter of the respective year;[19]

L = natural logarithm of the number of employees of all banks located in SMSA, measured as of mid-March of the respective year;

M = dummy variable for three-bank concentration ratio as of 1964, based on total bank deposits in SMSA, taking the value of one when concentration equals or exceeds 76 percent, and zero otherwise;

Y = natural logarithm of total personal income in SMSA in respective years;

W = natural logarithm of manufacturing hourly wage rate in SMSA;

B = unit-branch dummy variable, zero if SMSA is a unit-banking state, and one otherwise;

D = natural logarithm of the ratio of bank demand deposits to bank savings deposits in the SMSA;

LD = natural logarithm of the ratio of bank loans to total bank deposits in the SMSA;

BO = natural logarithm of the average number of branches per bank in the SMSA.

Equations (1)–(4) in table 1 show that the coefficient of M is significantly greater than zero, indicating that monopolistic banks have higher labor expenses and larger staffs, as predicted by the expense-preference theory. In addition, concentration (M) has almost an identical impact on TW and L (the respective regression coefficients in eqq. [2] and [4] are .216 and .191), which suggests that management indulges its taste for expenses primarily by hiring excess staff (rather than by paying higher salaries).[20] Lastly, all equations in table 1 perform quite well: R^2s are high, and the signs and magnitudes of the coefficients of the independent variable W and Y are consistent both with theory and prior empirical

[19] Data for TW and L on an SMSA basis are available only for the first quarter of each year. See U.S. Bureau of the Census (1962, 1964, 1966).

[20] Otherwise, we would expect the coefficient of M to be substantially greater in equation (2) than in equation (4). In addition, when I estimated regression equations using a bank-specific wage rate as the dependent variable, I was unable to find any relationship between this wage rate and concentration.

work (see, for example, Bell and Murphy 1968, pp. 78–93). Thus, thes results strongly support the expense-preference theory.[21]

To determine the role that branching plays in expense-preferenc behavior, equations (5)–(8) in table 1 were estimated. Equations (5 and (6) add the variable "average number of branch offices per bank' (BO) to equations (2) and (4) and are estimated with the full dat sample, while equations (7) and (8) are also estimated with variable B included but after omitting all unit-banking observations.[22] In all o these equations the coefficient of M remains significantly greater tha zero, although, as expected, its magnitude falls somewhat when th degree of branching is introduced.[23] Thus, in addition to providing stron support for the expense-preference hypothesis, the results in equation (5)–(8) suggest that excessive branching is not one of the primary way that bank managers indulge their expense-preference proclivities.

V. Conclusions and Implications

The finding that wage and salary expenditures in banking increase wit monopoly power indicates that an expense-preference model may be more useful framework for describing and predicting bank behavior tha is the traditional profit-maximization model. More generally, my result suggest that the managerial objective function of regulated firms is no one of profit maximization.

An expense-preference model of bank behavior explains a number o previous findings which have heretofore gone unexplained. For example the finding that prices (or interest rates) are higher in monopolisti banking markets while reported profits are not can be explained by th monopoly-cost relationship implied by expense-preference behavior. I addition, previous findings that increased competition (due to new entry) does not reduce the reported profitability of existing banks can be explained by the fact that expense-preference managements respond t greater competitive pressure by reducing costs.

[21] This conclusion, it should be noted, holds for several other "switching levels" o concentration as well (see n. 17 above). In addition, the results are insensitive to using a three-way dummy variable for branching, splitting B into limited branching and statewide branching states. Our sample consists of eight unit-banking SMSAs, 25 limit-branching SMSAs, and 10 statewide-branching SMSAs.

[22] In my full sample the simple correlation between M and ln BO is .40, while in the sample omitting all unit-banking observations the simple correlation is .15. In the latte sample, banks in low concentration markets have an average of 5.63 branches, and banks in high concentration markets have an average of 6.68 branches.

[23] It makes no difference whether BO or ln BO is used in these equations. In addition, there is no multicollinearity between M and BO in these equations. Using the full sample, the partial correlation coefficient between M and ln BO is $-.07$; and, using the sample omitting unit-banking observations, the partial correlation coefficient between M and ln BO is $-.15$. Neither is significant at the 5 percent level. The negative relationship suggested by these correlations is, incidentally, consistent with White's findings (see n. 15 above).

Finally, the findings of this study suggest that we may be overlooking an important social cost of either monopoly or regulation if we ignore the monopoly-cost relationship implied by expense-preference behavior.

Appendix

Equation (1) in the text is derived by assuming the firm is a monopoly that maximizes the objective function

$$\pi = PQ - WL - rk, \tag{6}$$

where P is the price of the firm's output Q, W is the wage rate, r is the cost of capital, and L and K are the respective amounts of labor and capital employed. The firm is also assumed to have the Cobb-Douglas production function $Q = AK^bL^c$, so that the standard first-order conditions for a maximum are[24]

$$c(Q/L)P\eta - W = 0, \tag{7}$$

$$b(Q/K)P\eta - r = 0, \tag{8}$$

where $\eta = 1 + (1/e)$ and where e is the elasticity of the demand schedule for the firm.[25] Finally, by making the usual assumption that the final demand schedule for the firm's product is of constant elasticity

$$P = j_0 Q^{j_1} Y^{j_2}, \tag{9}$$

where j_1 is the reciprocal of the monopolistic industry demand elasticity (or $1/e$), and Y is a shift variable which represents all nonprice factors, the following demand for labor (or eq. [1] in the text) can be derived:[26]

$$\ln L = b_0 + b_2 \ln W + b_3 \ln r + b_4 \ln Y + \ln \hat{\eta}, \tag{10}$$

where

$$b_0 = \ln \left\{ \left[A \left(\frac{b}{c} \right)^b \right]^{(1+j_1)} cj_0 \right\} / (1 - b_1)$$

$$b_1 = (b + c)(1 + j_1)$$

$$b_2 = \frac{b(1 + j_1) - 1}{1 - b_1}$$

$$b_3 = - \frac{b(1 + j_1)}{1 - b_1}$$

$$b_4 = \frac{j_2}{1 - b_1}$$

$$\ln \hat{\eta} = \frac{\ln \eta}{1 - b_1} .$$

[24] All marginal products are assumed greater than zero. This type of production function has been found to be appropriate for banking. See, for example, Bell and Murphy (1968, chap. 5).

[25] For a monopolist, $\eta = 1 + (d \ln P/d \ln Q)$, which is less than one but greater than zero, since a monopolist will operate only in the elastic portion of the demand curve $(-\infty < e < -1)$. For a firm in a perfectly competitive industry $\eta = 1$, since such a firm will see its demand curve as being infinitely elastic.

[26] See Gould (1966, pp. 16–18) and Waud (1968, p. 409). My derivations for both the monopoly and competitive industry are similar to those of Gould.

A similar, but not identical, aggregate demand-for-labor equation can be derived for firms operating in a perfectly competitive market. In this case, of course, the firm's output will not equal the industry output. At each point in time, however, the competitive firm's output will be some proportion of total industry output, and this proportion will be determined by factors exogenous to the firm. This proportion can be defined as $s_i = (Q_i/Q)$, where Q is industry output and Q_i is the firm's output. In this analysis s_i is treated as an exogenous variable, just as price is in the usual theory of the competitive firm.[27]

Thus, the demand equation (9) can be rewritten as

$$P = j_0 \left(\frac{Q_i}{s_i}\right)^{j_1} Y^{j_2}. \tag{11}$$

Further, if the firm is a competitor, equation (7) becomes

$$c(Q_i/L_i)P = W. \tag{12}$$

Substituting (11) for P in (12) and making the same substitution for K_i in the production function as was done in the previous derivation for the monopoly firm, we have the following demand-for-labor equation for a competitive firm:

$$L_i = e^{b_0} W^{b_2} r^{b_3} Y^{b_4} s_i^{b_5}, \tag{13}$$

where $b_5 = (-j_1)/(1 - b_1)$.

Finally, assuming that all firms have identical Cobb-Douglas production functions, an aggregate demand-for-labor function for a competitive industry (or market) can be derived by summing (13) over all firms:

$$L = \sum_{i=1}^{n} L_i = \sum_{i=1}^{n} e^{b_0} W^{b_2} r^{b_3} Y^{b_4} s_i^{b_5}, \tag{14}$$

or

$$L = e^{b_0} W^{b_2} r^{b_3} Y^{b_4} \sum_{i=1}^{n} s_i^{b_5}, \tag{15}$$

since all terms except the last term are the same for all firms in a competitive market. Expression (15) can be further simplified if constant returns to scale are assumed. In this case, $b_5 = 1$ for all firms and the term

$$\sum_{i=1}^{n} s_i^{b_5}$$

becomes 1, since the s_i's sum to one in the aggregate. Under these assumptions the resulting aggregate demand-for-labor function for a competitive industry (or eq. [2] in the text) is

$$\ln L = b_0 + b_2 \ln W + b_3 \ln r + b_4 \ln Y. \tag{16}$$

Lastly, equation (4) of the text is derived by assuming that the firm maximizes the utility function

$$U = U(\pi, E), \tag{17}$$

where π is total profits and E is the firm's total labor expenses.[28] Given our

[27] Alternatively, we could assume that the number of firms in a competitive market is exogenous.

[28] The derivation which follows could also be done without any essential difference if it were assumed that firms maximize utility subject to the constraint that π equal or exceed some level.

previous assumptions and assuming that an interior solution exists, the first-order condition for a maximum is

$$c \frac{Q}{L} P\eta = W(1 - U_E/U_\pi), \tag{18}$$

where U_E is the marginal utility of a dollar of additional labor expenses; U_π is the marginal utility of another dollar of reported profits; and P, Q, L, W, c, and η are as defined above. The ratio U_E/U_π, therefore, can be considered to be an index of the firm's intensity of expense preference. In competitive markets, $U_E/U_\pi = 0$, and in monopoly markets, $U_E > 0$, and $U_E/U_\pi > 0$ but < 1.[29] Further, if we assume that all monopoly firms have identical expense-preference intensities, this ratio can be considered a constant.

Thus, using the first-order condition expressed by equation (18) (instead of eq. [7]) to derive the firm's demand for labor, the expense-preference firm's demand-for-labor equation becomes

$$\ln L = b_0 + b_2 \ln (1 - U_E/U_\pi) + b_2 \ln W + b_3 \ln r + b_4 \ln Y + \ln \hat{\eta}, \tag{19}$$

which is equation (4) in the text.

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[29] This follows from the first-order conditions and the assumption of an interior maximum. Specifically, the above first-order condition together with the capital first-order condition (eq. [8] above) yield $1 - U_E/U_\pi = (r/W)(f_L/f_K) > 0$, since r, W, f_L, and f_K are all positive, given the assumption that firms do not "waste." It follows, therefore, that $U_E/U_\pi < 1$.

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