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Taxes, Failure Costs, and Optimal Industry Capital Structure: An Empirical Test

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SINCE 1958 WHEN FRANCO Modigliani and Merton Miller (M-M) provided the first proof, invariance propositions have been central to the theory of corporate finance. M-M stated their proposition in two equivalent forms: that the "value of a firm is independent of its capital structure" and that "the average cost of capital to any firm is completely independent of its capital structure."[1] The rationale is well known. Given the expected return on assets for a firm, individual arbitrage[2] will ensure that the value of the firm is simply this expected return capitalized at the rate appropriate to firms in the same risk class. This rate is the average cost of capital and does not depend upon financing decisions of the firm. Consequently, the value of the firm does not depend upon its financing decisions but only on its risk class and its investment decisions.

If taxes, in particular the preferential treatment of interest payments, and failure costs are included in the analysis, the two forms of the original M-M proposition may no longer be equivalent. Even if the financial decisions of the firm do not affect its cost of capital, they still may change the firm's expected earnings by altering its tax liability and probability of failing.[3] These changes in expected earnings would affect the value of the firm. It is thus possible for the second form of the original M-M proposition to be true even if the first form is not. Additionally, by altering characteristics of the firm's income stream besides its expected value, changes in capital structure may affect its average cost of capital.[4] Then neither form of the M-M proposition would be true.

The empirical validity of either M-M proposition then depends (at least partially) upon the empirical relevance of taxes and failure costs. Little testing has been done. Recently Miller [12] has suggested that the size of the net tax subsidy to debt over all market participants and the direct costs of bankruptcy may be small, but little has been done to relate these factors to variations in capital structure. That is what we attempt here. The form of the M-M proposition that we test in this study is that although taxes and costs of failure do not affect the average cost of capital, they do affect expected income and so imply an optimal capital structure. More precisely, measures are constructed of the tax advantage to debt and the costs of failure for 38 major industries, and an attempt

[1] Modigliani and Miller [14], pp. 268–69.

[2] Where firms in the same risk class exist, arbitrage by individual investors results in this invariance. If similar firms do not exist, competition among intermediaries to buy all the financial instruments of a firm and then issue their own stock and debt will accomplish the same result. See Stiglitz [17].

[3] In their original paper, Modigliani and Miller [14] note this when they consider the effect of corporate taxes (see p. 272).

[4] Modigliani and Miller [13].

is made to relate these variables to cross-sectional and temporal variation in industry capital structure for the period 1957–1972.

Section I describes the model to be estimated. Section II addresses and attempts to resolve some rather serious problems in deriving the required data. Section III describes the results of the statistical estimation. A final section summarizes this work and offers tentative conclusions.

I. The Model

Several previous authors have rigorously shown that a tax advantage to corporate debt combined with failure costs may imply an optimal capital structure. The papers of both Kraus and Litzenberger [11] and Scott [15] offer such theorems. The framework in which these theorems were derived is one in which the real investments of the firm are independent of the financial decisions, and there are no clientele effects (changes in financial policy do not attract shareholders with differing characteristics). In the subsequent empirical work of this paper, we also adopt both of these assumptions.[5] However, since the industry is our unit of analysis, it is not necessary that they apply at the level of the firm, only at the level of the industry.

To discuss the model that will be tested, we employ the "typical firm" device. That is, we treat the industry variables as pertaining to characteristics of a (fictitious) firm that is a microcosm of its industry. Let y represent the annual expected operating earnings of the firm (ignoring failure costs) before interest and taxes (EBIT). Interest on debt, i, is a fixed and prior claim on earnings. The residual $(y - i)$ is the property of equity holders that they realize either as dividends or capital gains.

Increases in interest, given EBIT, increase the likelihood of default on debt. We assume that creditors, to induce managers to act with appropriate aversion to default, stipulate in their debt contracts that default implies "failure" of the firm.[6] "Failure" is broadly descriptive of a range of possible settlements between insolvent firms and their creditors. Examples include extension or other negotiated arrangements and court-declared bankruptcy with forced reorganization. Each of these possibilities is costly to the firm. Importantly we presume that failure never results in liquidation of the firm's assets. That is, failure, though costly, does not terminate the earnings stream of the firm.[7] The cost of failure, C, include the administrative costs of an adjudication or negotiation process such as

[5] One interpretation is that these are among the hypotheses that we seek to test jointly. Additionally, there exists empirical support for the second assumption. Crockett and Friend [4], and Blume, Crockett and Friend [2], using large samples of federal income tax returns, determined that with the exception of utilities and the telephone-communications industry, the income profile of owners was nearly the same across major industries. The vastly differing capital structures of these industries apparently has not precipitated measurable differences among their shareholders' incomes. The evidence is consistent with our assumed absence of clientele effects at the industry level.

[6] That is, alternative arrangements that preclude failure altogether (such as preferred stock in place of bonds), are not adopted because the added agency costs would outweigh the eliminated failure costs (cf., Jensen and Meckling [10], p. 334).

[7] Scott [15], p. 48 uses the word "reorganization" in the sense that "failure" is meant here. He defines "bankruptcy" as always implying liquidation of assets and termination of the firm's earnings. Scott's model becomes similar to ours if "bankruptcy" in his sense has a zero probability of occurring but "reorganization" does not.

attorney or court fees, as well as indirect costs due to disruption of ongoing commercial relations or attenuated stockholder control during receivership.

More formally, letting Π represent the annual probability of failure and defining R as operating risk,

$$\Pi = f\left(\frac{i}{y}, R\right), \tag{1}$$

where

$$\frac{\partial \Pi}{\partial \frac{i}{y}} > 0, \qquad \frac{\partial \Pi}{\partial R} > 0.$$

Notice that including $\frac{i}{y}$ rather than (i, y) as an argument reflects a judgment that probability of failure is homogeneous of degree zero in i and y. We define $\frac{i}{y}$ as the capital structure of the firm. To simplify notation, let $D \equiv \frac{i}{y}$. It is assumed that Π is convex in both i and R.

The measure of operating risk we employ is one minus the coefficient of determination $(1 - R^2)$ of EBIT on a log linear time trend. This accords with the conventional view that operating risk ought to be invariant to the scale of the firm and closely related to the annual variability of EBIT.[8]

The specific form of the probability-of-failure function that we adopt in the empirical portion of the paper is the logistic (logit) form.

$$\log \frac{\Pi}{1 - \Pi} = \beta_0 + \beta_1 \log \frac{D}{1 - D} + \beta_2 \log R, \tag{2}$$

where

$$0 < \Pi < 1 \quad \text{and} \quad 0 < D < 1.$$

A useful feature of this functional form is that it allows the calculation of $\frac{\partial \Pi}{\partial i}$ up to a scalar multiple (β_1), which is the same for all industries.

$$\frac{\partial \Pi}{\partial D} = \beta_1 \cdot \frac{\Pi(1 - \Pi)}{D(1 - D)}, \tag{3}$$

which implies that[9]

[8] Scott [15], pp. 47–48 notes that these characteristics of operating risk can only be established under special assumptions. Consequently, our measure of R may not always be appropriate. However, other measures of R such as variance, coefficient of variation and the trend coefficient for industry EBIT were also tried. These other measures yielded less satisfactory results.

[9] Note:

$$D = \frac{i \cdot n}{Y}$$

$$\frac{\partial D}{\partial i} = \frac{n}{Y} = \frac{D}{i},$$

where: i = interest per firm, n = number for firms, and Y = industry operating income.

$$\frac{\partial \Pi}{\partial i} = \beta_1 \cdot \frac{\Pi(1 - \Pi)}{(1 - D)i}. \tag{4}$$

The product of Π and C is the annual expected failure cost. Failure costs, if they occur, will be borne by equity holders as a capital loss. Consequently, expected failure costs must be subtracted from EBIT along with interest in order to calculate the expected annual residual earnings by equity holders.

The earnings of the firm are subject to the corporate income tax and also to the personal income tax since these earnings accrue to security holders. Interest payments are deductible for corporate tax purposes, so the annual after-tax value of interest is $(1 - \tilde{\tau}_{pb})i$ where $\tilde{\tau}_{pb}$ is the average personal tax rate on bond interest income. (Where tax rate symbols appear without tildes, they denote marginal tax rates.) The taxation of the residual income of equity holders $(y - i - \Pi \cdot C)$ is less simple. The residual is first taxed at the corporate tax rate $\tilde{\tau}_c$, and then taxed again as it is received as income by equity holders. Income received by individuals as dividends is taxed at the same rate as ordinary income,[10] but income received as capital gains is taxed at a lower rate. Let $\tilde{\tau}_d$ be the personal tax rate on dividends, $\tilde{\tau}_{cg}$ be the personal tax rate on capital gains, and α be the dividend payout ratio (the fraction of $y - i$ that is received by stockholders as dividends). The annual (expected) after-tax value of the residual, then, can be written

$$(1 - \tau_c)(\alpha\tau_d + (1 - \alpha)\tau_{cg})(y - i) - \Pi(1 - \tau_c)(1 - \tau_{cg})C.$$

This formulation can be simplified by noting that during the period we examine, personal taxes on capital gains were avoided entirely if not realized within the stockholder's lifetime. That is, heirs were not liable for capital gains that accrued prior to the death of their benefactors. A study by Bailey [1] suggests that a very large fraction of capital gains escaped being taxed in this fashion (perhaps more than 80 percent). To simplify, we assume that $\tilde{\tau}_{cg}$ is zero. Now it is possible to write the annual after-tax value of the residual as[11]

$$(1 - \tilde{\tau}_c)(1 - \alpha\tilde{\tau}_d)(y - i) - \Pi(1 - \tilde{\tau}_c)C.$$

The preceding discussion of failure costs and taxes suggests the following formal model of optimal capital structure. The value of the firm, V, is equal to the summed values of its securities: debt, B, and equity, E.

$$V \equiv B + E. \tag{5}$$

Let r_b and r_e represent the after-tax capitalization rates on debt and equity, respectively. Also, define r as the overall after-tax capitalization rate for the firm which, as previously stated, is assumed to be exogenous (independent of capital structure decisions). Then,

$$V = \frac{r_b B + r_e E}{r}. \tag{6}$$

[10] The \$100 dividend exclusion is ignored.

[11] Our formulation assumes that the tax advantage to interest is a sure one. This is not strictly true in that a firm may have accrued losses for tax purposes (a tax credit carryover that would reduce taxes on future earnings) at the time that it permanently suspends operation. In this case, the tax advantage to interest could not be realized. See Brennan and Schwartz [3] on this point. Note that suspension of operations is entirely distinct from "failure" as we define it.

The terms in the numerator are simply the annual expected after-tax values of interest and residual earnings that were described in the previous discussion. Substitution into the numerator yields

$$V = \frac{(1 - \tilde{\tau}_{pb})i + (1 - \tilde{\tau}_c)(1 - \alpha\tilde{\tau}_d)(y - i) - \Pi(1 - \tilde{\tau}_c)C}{r}. \tag{7}$$

The optimal interest payment for the firm, i^* (and so the optimal capital structure of the firm), is simply that which maximizes V. A necessary condition is that i^* solves the equation[12]

$$\frac{\partial V}{\partial i} = \frac{(1 - \tau_{pb}) - (1 - \tau_c)(1 - \alpha\tau_d) - \dfrac{\partial \Pi}{\partial i}(1 - \tau_c)C}{r} = 0. \tag{8}$$

A sufficient condition that equation (8) solves for a maximum value of V is that

$$\left.\frac{\partial^2 V}{\partial i^2}\right|_{i=i^*} = \frac{\partial^2 \Pi}{\partial i^2}\frac{(1 - \tau_c)}{r}C \equiv A < 0. \tag{9}$$

The previous assumption that the annual probability-of-failure function is convex in i, implies that the sufficient condition is satisfied.

Equation (8) can be simplified by defining

$$G \equiv (1 - \tau_{pb}) - (1 - \tau_c)(1 - \alpha\tau_d). \tag{10}$$

G is the marginal annual tax advantage to interest, given EBIT. Substituting G into equation (8) yields

$$\frac{\partial \Pi}{\partial i}\frac{(1 - \tau_c)}{r}C = \frac{G}{r}. \tag{11}$$

The interpretation is straightforward. The marginal expected failure cost of interest is just offset by its marginal tax advantage.[13]

Presuming that equation (11) characterizes capital structure, we use this condition to impute failure costs. Rearranging terms solves for C:

$$C = \frac{G}{(1 - \tau_c)\dfrac{\partial \Pi}{\partial i}}. \tag{12}$$

Recalling the specific form of the probability of failure function that we have adopted (equation (2)), it is possible to substitute the implied value of $\dfrac{\partial \Pi}{\partial i}$ so that

$$C = \frac{1}{(1 - \tau_c)\beta_1}\frac{G(1 - D)i}{\Pi(1 - \Pi)}. \tag{13}$$

[12] The assumption that there do not exist clientele effects at the industry level, which is equivalent to:

$$\frac{\partial \tau_{ps}}{\partial\left(\dfrac{i}{y}\right)} = \frac{\partial \tau_{pb}}{\partial\left(\dfrac{i}{y}\right)} = 0,$$

is used in deriving equation [8].

[13] This is directly analogous to the discussion by Scott [15], p. 43, of conditions for an optimal capital structure within his assumed framework.

This is our measure of failure costs.

Warner (18) has performed the only other study to date that attempts to measure failure costs. He calculated the direct bankruptcy costs for 11 major railroad failures between 1930 and 1955. We believe our approach is superior for two reasons. First, our measure includes not only the direct costs such as lawyer and court fees that Warner measures, but also the indirect costs such as disruption of ongoing commercial relations. Second, we allow for a broader definition of failure that includes possibilities other than just those that result in court proceedings (bankruptcies).

To determine the effects of changes in parameters on optimal capital structure, we first take the total differential of equation (8).

$$Adi + \frac{1}{r}dG - \frac{\partial^2\Pi}{\partial i\partial R}\cdot\frac{(1-\tau_c)C}{r}dR - \frac{\partial\Pi}{\partial i}\frac{(1-\tau_c)}{r}dC - \frac{\partial^2\Pi}{\partial i\partial y}\frac{(1-\tau_c)C}{r}dy - \frac{G - \dfrac{\partial\Pi}{\partial i}(1-\tau_c)C}{r^2}dr = 0. \tag{14}$$

Solving this equation for separate changes in C, G, R, y and r provides the following comparative statics results.

$$\frac{\partial i^*}{\partial C} = \frac{\dfrac{\partial\Pi}{\partial i}\cdot\dfrac{(1-\tau_c)}{r}}{A} < 0 \tag{15}$$

$$\frac{\partial i^*}{\partial G} = -\frac{1}{Ar} > 0 \tag{16}$$

$$\frac{\partial i^*}{\partial R} = \frac{\dfrac{\partial^2\Pi}{\partial i\partial R}\cdot\dfrac{(1-\tau_c)C}{r}}{A} < 0 \tag{17}$$

$$\frac{\partial i^*}{\partial y} = \frac{\dfrac{\partial^2\Pi}{\partial i\partial y}\cdot\dfrac{(1-\tau_c)C}{r}}{A} \gtrless 0 \tag{18}$$

$$\frac{\partial i^*}{\partial r} = \frac{G - \dfrac{\partial\Pi}{\partial i}\cdot\dfrac{(1-\tau_c)C}{r^2}}{A} = 0. \tag{19}$$

Notice that equation (19) implies that if two firms differ in no respect other than their after-tax capitalization rates, then they should have identical capital structures.[14] Another property of the model also deserves special emphasis.

[14] Interfirm differences in capitalization rates, r, are based on differences in the market's evaluation of the riskiness of the firms' net earnings streams. In fact, such risk differences would likely also entail differing probability of failure functions as well . . . but this is not necessarily so. For example, suppose two firms have similar dispersions in EBIT over time but for one firm these dispersions closely track changes in the income from a market portfolio, whereas in the other they do not. The two firms may have identical probability of failure functions, but the former is likely to have a greater capitalization rate.

Recalling that the assumed probability-of-failure function Π is homogeneous of degree zero in i and y, it is evident from inspection of equation (8) that $i^* = f(C, y, G, R)$ is homogeneous of degree one in C and y.[15] Thus, we can write

$$D^* = \left(\frac{i^*}{y}\right) = f\left(\frac{C}{y}, G, R\right). \tag{20}$$

This is the general form of the equation that ought to explain cross-sectional and temporal variation in $\frac{i}{y}$. We will refer to this as the "optimal capital structure equation." Estimation of this equation for major industries is a principal object of this paper. Hypotheses regarding the signs of the derivatives of equation (20) can be inferred from the previous comparative statics results.[16]

$$\frac{\partial D^*}{\partial\left(\frac{C}{y}\right)} = \frac{\partial i^*}{\partial C} < 0 \tag{21}$$

$$\frac{\partial D^*}{\partial G} = \frac{1}{y}\frac{\partial i^*}{\partial G} > 0 \tag{22}$$

$$\frac{\partial D^*}{\partial R} = \frac{1}{y}\frac{\partial i^*}{\partial R} < 0. \tag{23}$$

We next specify the data that were used to calculate our measures of D, G, C, and R.

II. Data

The units of analysis in this study are major industries as classified by the Internal Revenue Service. This corresponds roughly to the three-digit SIC level of aggregation. As many industries are included as the data permit—38 in all. For each industry, data are based on two eight-year time periods: 1957–1964 and 1965–1972. Each period is intended to span a time sufficient to measure accurately expected or target values of the variables, but short enough to pertain to a relatively stable tax and risk environment.

Our definition of capital structure, D, is interest payments divided by annual expected operating income (EBIT). These data are drawn from *IRS Statistics of Income, Corporate Income Tax Returns.* To approximate expected values, both interest and operating income are averaged over the two eight-year periods.

To calculate the annual tax advantage to interest (holding EBIT constant), or $G = (1 - \tau_{pb}) - (1 - \tau_c)(1 - \alpha\tau_d)$, estimates must be made of the marginal tax rates on corporate income, stockholder (dividend) income, and bondholder income. The marginal corporate income tax rate, τ_c, was simply assumed equal to 0.48 in both time periods. Because the personal income tax is progressive, the

[15] Cf. Scott [15], p. 47.

[16] The fact that $\frac{i^*}{y} = f\left(\frac{C}{y}, G, R\right)$ is homogeneous of degree zero in C and y is used in deriving equations (21), (22), and (23).

marginal tax rate on dividend income, τ_d, depends upon the income profile of stockholders. To calculate τ_d by industry then, it is necessary first to develop an income profile of the owners of each industry. Fortunately, other researchers have already addressed this problem. Crockett and Friend [4], and Blume, Crockett and Friend [2], using a large sample of federal personal income tax returns, determined that, with the exception of utilities and the telephone-communications industry, the income profile of owners was nearly the same across all industries.[17] Therefore, we assume the marginal personal income tax rate on dividends to be the same for all industries.

The next step is the actual calculation of this tax rate. Here again the Crockett and Friend, and Blume, Crockett and Friend studies provide the necessary information. These two studies detail the distribution of dividends among income classes (AGI classes from the personal income tax forms) and institutions for the years 1960 and 1971.[18] Multiplying the fraction of dividends received by each group by that group's marginal tax rate and then summing over all groups results in the overall (mean) marginal tax rate on dividends, or τ_d. This calculation is described in detail in Appendix A.

To calculate the marginal personal tax rate on interest payments, it is also necessary to know the income profile of corporate bondholders. Lacking such a profile, we made a simplifying assumption: τ_{pb} is equal to τ_d.

The dividend payout rate α is calculated by dividing the sum of cash dividends paid by an industry over the period by the sum of net income (EBIT less interest) in this industry over the same period. That is, $\hat{\alpha}$ is the average dividend payout rate over the period. (The symbol "ˆ" is used to distinguish our empirical measures from their theoretical counterparts.)

Based upon this set of calculations, a value of the annual tax advantage to interest, $\hat{G}$, was constructed for each industry (variation among industries depends on differences in industry dividend payout rates) for the years 1960 and 1971.

Our measure of failure costs was earlier defined as

$$C = \frac{1}{(1-\tau_c)\beta_1} \frac{G(1-D)i}{\Pi(1-\Pi)}.$$

The sources and calculation of τ_c, G, D, and i already have been described. Constructing a feasible measure for the annual probability of failure, Π, is a difficult task. Dun & Bradstreet, Inc. reports the number of failures by industry in its weekly and quarterly *Failure Reports*. These numbers are based on court records of Ch. X and Ch. XI bankruptcy cases, and on information solicited from creditors. To convert these data into estimated failure rates, an appropriate base must be selected. The population from which these data are drawn is "active" firms in the *Dun & Bradstreet Reference Book*. (In the jargon of Dun & Bradstreet, "active" means "in the sample population," not "in business.") The

[17] Low-income classes held a larger percentage of the stock of utilities and the telephone-communications industry than they did in other industries. Note that in our empirical section, a regulation dummy variable is included which, among other things, adjusts for ignoring the "clientele effect" for utilities and the telephone-communications industry (both of which are regulated).

[18] The figures are contained in Table 1.3 at p. 152 of the Crockett and Friend study [14] and in Table G at p. 40 of the Blume, Crockett and Friend study [2].

Table I

Failure Cost Regressions: Dependent Variable $= \log \hat{C}$

Coefficients on explanatory variables, with t-statistics					
Years	Intercept	Log $\hat{y}$	*REG*	*F*	R^2
1957-1964	.216 (1.196)	.960 (9.856)	1.409 (2.348)	63.118	.783
1965-1972	1.527 (11.836)	.973 (12.719)	1.015 (2.214)	95.458	.845

Note: Each regression is based on 38 observations. See text for explanation of variables.

Reference Book (in 1978) lists roughly three million businesses, incorporated and unincorporated, of which about two million are "active." The number of active firms by industry is unavailable from Dun & Bradstreet for any year. The number of firms by industry in the *Reference Book* is available indirectly but only for selected industries, through the annual *Failure Report*. Faced with undesirable alternatives, we chose as a base for constructing industry failure rates the number of corporations filing incomê tax returns by industry according to the *IRS Statistics of Income*. This choice involves strong assumptions that the IRS population of corporations measures the Dun & Bradstreet population of firms, and that the Dun & Bradstreet sampling procedures for deriving number of failures are not biased. We do not estimate β_1. β_1, however, will be the same for all industries, so our empirical measure of failure costs, $\hat{C}$, is to be interpreted only as a relative measure of failure costs. For the estimations in the following section, $\frac{1}{(1 - \tau_c)\beta_1}$ is (arbitrarily) set equal to one, and means of annual observations are used for $\hat{D}$, $\hat{y}$, and $\hat{\Pi}$.[19]

We measure operating risk, R, as one minus the coefficient of determination of log-linear trend regressions on industry EBIT. The source of the data for EBIT is the *IRS Statistics of Income*.

III. Results

Before estimating an optimal capital structure equation, we first explore the relation between failure costs and EBIT. Recall from the discussion of the model in Section I that proportionate variation in failure costs and EBIT ought not to have any impact on capital structure.

The relation between our estimated failure costs, $\hat{C}$ (with unknown scalar multiple $\frac{1}{(1 - \tau_c)\beta_1}$ set equal to one) and EBIT per firm $\hat{y}$ is indicated by the regressions in Table 1. Failure costs are estimated in the way described in Section

[19] The resulting estimated costs of failure (up to a scalar multiple) are reported in a data appendix available from the authors.

II. A dummy variable, REG, is included that assumes the value one for each of the three regulated industries and zero for all others.

The close fits in the equations are encouraging. Also, except for the intercepts, the two regressions are remarkably similar. Because C is a relative measure only, the upward adjustment of the intercept in the second period is subject to ambiguous interpretation. If the (unknown) scale coefficient $\frac{1}{(1-\tau_c)\beta_1}$ is the same for both periods, the change of intercept has the following meaning. Within the logic of the formal model, industries behaved in the later period as though failure costs had risen as a proportion of EBIT; in fact, as though $\frac{C}{y}$ had risen by about 370 percent! An alternative explanation for the higher intercept in the later equation is an upward adjustment in $\frac{1}{(1-\tau_c)\beta_1}$, the scale coefficient on $\hat{C}$. Recall from equations (2) and (12) that β_1 is a parameter in the probability-of-failure function. A decrease in β_1 means that, holding operating risk, R, and capital structure, D, constant, failure is less likely (on average and on the margin). A look at the data showed that just such an adjustment apparently occurred between the two periods. Though $\hat{D}$ rose on average and $\hat{R}$ rose (earnings streams showed greater deviation from trend), failure probabilities markedly diminished.[20] Thus, little should be inferred from the differences in intercepts of the two Table 1 regressions.

Turning attention to the other Table 1 point estimates, inter-industry influences of $\hat{y}$ and REG on $\hat{C}$ are essentially the same in both periods. The elasticity between $\hat{y}$ and failure cost is approximately unitary, suggesting that the ratio $\frac{C}{y}$ does not vary across industries. This is a highly significant finding because it implies that failure costs ought not to explain any of the inter-industry variation in capital structure. In contrast to our estimates, Warner's data ([18], pp. 341–42) imply an elasticity of direct failure cost with respect to market value of the firm (at time of filing of the bankruptcy petition) equaling 0.423.[21] A comparison of that result with ours suggests that indirect failure costs (reflected in our estimate but not Warner's) are substantial and highly correlated with the size of the firm.

The positive sign on REG is difficult to interpret. There is a possible measurement error that confounds the issue. No account was taken in calculating $\hat{G}$ of the slightly lower incomes of shareholders of public utilities found by Crockett and Friend [4] and Blume, Crockett and Friend [2]. This suggests that $\hat{G}$ is too

[20] A possible explanation for the lower failure probabilities in the later period can be found in the increased merger activity in that period. Merger may be an attractive alternative to failure. A more inclusive and useful definition of "failure" should include at least some mergers. Accordingly, the lower estimated failure probabilities in the later period may reflect "measurement error."

[21] The regression based on Warner's data (t-statistics in parentheses) is given here:

$$\log C = \underset{(-3.589)}{-0.979} + \underset{(5.806)}{0.423} \log V \qquad R^2 = .789, \qquad F = 33.708$$

where C = estimated direct failure costs, and V = market value of firm at filing of the bankruptcy petition.

small for these industries and consequently, so is $\hat{C}$.[22] Given this caveat, however, there exists the interesting possibility that some feature of the regulatory process leads firms to select capital structures as though they perceived higher costs of failure (not probabilities of failure) than unregulated, but otherwise similar firms. This may be due to failure costs that are, in fact, greater for regulated firms or it may be due to the incompleteness of our model for regulated firms. That is, there may be an additional disadvantage to debt that arises because of the regulatory process. With regard to this last point, Herendeen [7] and Sherman [16] find that rate-setting based on the guidelines of the *Hope* decision may, in fact, create a bias toward equity in regulated firms. Our result here supports this finding.

We now turn to estimation of equation (20), the optimal capital structure equation. Our capital structure measure, $\hat{D}$, is constrained to lie between zero and one. The usual assumption of a normally distributed error term limits our choice of dependent variable to functions of $\hat{D}$ that map from the unit interval to the real line. Although any inverse cumulative distribution function has this characteristic, the inverse of the logistic c.d.f. (already used in deriving $\hat{C}$) permits linear estimation, simplifies calculations, and at the same time is somewhat general. It therefore is a logical choice.

Accordingly, the capital structure equations we estimate are of the form

$$\log\left(\frac{\hat{D}}{1-\hat{D}}\right) = \gamma_0 + \gamma_1 \hat{G} + \gamma_2 \hat{R} + \gamma_3 \,\mathrm{REG} + \gamma_4\left(\frac{\hat{C}}{\hat{y}}\right) + \xi. \tag{24}$$

Notice that although $\frac{\hat{C}}{\hat{y}}$ is included as an argument of equation (24), its coefficient could not be estimated with any degree of confidence. This follows from the absence of significant variation in $\frac{\hat{C}}{\hat{y}}$, which was detailed in the previous discussion. Accordingly, $\frac{\hat{C}}{\hat{y}}$ is not included in the regressions described in the text below. See footnote 24 for a discussion of regression results that include $\frac{\hat{C}}{\hat{y}}$.

The dummy variable REG, with a value of one for the three regulated industries, is included in equation (24) for three reasons. First, there may be a measurement error in $\hat{G}$ for regulated industries due to our failure to take account of the slightly lower incomes of shareholders in regulated firms (see footnote 22). Second, the formal model of Section I may be incomplete for regulated industries (there may be an additional disadvantage to debt due to rate-setting practices). Third, the regulatory process may affect capital structure indirectly by altering the probability-of-failure function.

The OLS estimates of equation (24) for each of the two sample periods, separately and pooled, are reported in Table 2. We were, at first, encouraged by an apparently positive coefficient on $\hat{G}$ in the first time period, but inspection of residuals showed that this result depended crucially upon a single outlying

[22] This is somewhat ambiguous. Recall $\hat{G} = (1 - \tau_{pb}) - (1 - \tau_c)(1 - \alpha\tau_d)$. So $\partial\hat{G}/\partial\tau_d = \alpha(1 - \tau_c)$, which is positive. So an overstatement of the personal tax rate on dividends τ_d implies an overstatement of $\hat{G}$. However, if we assume that $\tau_d = \tau_{pb}$ (as we have indeed done), then $\partial\hat{G}/\partial\tau_d = \alpha(1 - \tau_c) - 1$, which is negative, and so the overstatement of τ_d implies an understatement of $\hat{G}$.

observation: motels and other lodging. Consequently, results shown in Tables 2, 3, 4, and 5 are based on samples deleting this industry.

The most striking characteristic of the estimates in Table 2 is the poor performance of $\hat{G}$ in cross sections, but good performance in the pooled equation. Noteworthy is the fact that the time-series variation in $\hat{G}$ is far greater than its cross-sectional variation. Cross-sectional variation in $\hat{G}$ is based on inter-industry differences in dividend payout rates, and the time-series change is mainly due to the decreases in personal tax rates in 1964, the last year of the early period. In Table 3 the quartiles and sample standard deviation of $\hat{G}$ have been recorded for each of the two time periods. It is evident that interperiod variation in $\hat{G}$ far exceeds inter-industry variation. Almost the entire cross-sectional variation in $\hat{G}$ is confined to a difference between mining and petroleum industries (five industries in all) and the rest of the sample. In contrast, every single industry experienced a marked time-series increase in $\hat{G}$. Thus, a credible interpretation of the results in Table 2 is that degrees of freedom are inadequate for relating D and G in each cross section separately, but sufficient when the two sets are pooled.

The pooled regression equation requires further study. OLS estimates of pooled time-series/cross-section regressions are seldom efficient because of the usually different random processes generating temporal and lateral variation. A statistical

Table II

Industry Capital Structure Regressions, OLS: Dependent Variable $= \log \dfrac{\hat{D}}{1 - \hat{D}}$

Data Set	Intercept	$\hat{G}$	$\hat{R}$	*REG*	*F*	R^2
1957–1964	−1.290	−0.256	−0.803	0.623	0.989	0.082
	(−3.108)	(−0.099)	(−0.831)	(1.493)		
1965–1972	−0.294	−2.478	−0.563	0.928	3.394	0.236
	(−0.915)	(−0.661)	(−1.876)	(2.731)		
Pooled	−2.004	3.462	−0.388	0.620	4.625	0.165
	(−6.906)	(2.565)	(−1.190)	(2.309)		

Note: There are 37 observations in each of the cross sections; 74 observations in the pooled sample. See text for definition of variables.

Table III

Quartiles and Sample Standard Deviation of $\hat{G}$, Both Time Periods[a]

Item	1957–1964	1965–1972
Lower quartile	.139	.234
Median	.149	.243
Upper quartile	.163	.256
Sample standard deviation	.045	.025

[a] Each time period contains 37 observations.

Table IV

Pooled Data Set (GLS: Fuller and Battese Method);

Dependent Variable = $\log \frac{\hat{D}}{1-\hat{D}}$

Coefficients on Explanatory Variables with t-statistics				Variance Component Estimates		
Intercept	$\hat{G}$	$\hat{R}$	REG	Cross Section	Time-Series	Error
−2.585 (−10.017)	4.973 (6.024)	0.063 (0.271)	0.554 (1.561)	0.282	0.000	0.080

Note: The error term in the regression (equation (24)) is assumed to have the decomposition

$$\xi_{it} = \nu_i + e_t + \epsilon_{it},$$

with the errors ν_i, e_t and ϵ_{it} independently distributed with zero mean and variances of $\sigma_\nu^2 \geq 0$, $\sigma_e^2 > 0$, $\sigma_\epsilon^2 > 0$, respectively. Estimates of these variances appear in the last three columns of the table. See Fuller and Battese [6] for discussion.

Table V

Estimated Marginal Effects of G and REG on Industry Capital Structures

D	$\frac{\partial D}{\partial G} = D(1-D)\hat{\gamma}_G$	$\frac{\Delta D}{\Delta REG}$ [a]
.10	.488	.062
.15	.637	.085
.20	.796	.103
.25	.935	.117
.30	1.044	.127
.40	1.194	.137
.50	1.243	.135

Note: Estimates are based on Table 4.

[a] $\frac{\Delta D}{\Delta REG} = D' - D^0$, where $D^0 = \frac{1}{1+e^{-\gamma x^0}}$ = first column in table and $D' = \frac{1}{1+e^{-\gamma x^0 - \hat{\gamma} REG}}$.

model incorporating these differences in error structure (GLS) is likely to yield more accurate and reliable estimates than OLS for pooled data. Our pooled equation was therefore re-estimated using the Fuller and Battese method, an Aitken-type transformation of the data.[23] The assumed error structure is

$$\xi_{it} = \nu_i + e_t + \epsilon_{it}, \tag{25}$$

where i indexes cross sections, t indexes time periods, and each of the error components is independently distributed with a zero mean and variances of σ_ν^2, σ_e^2, and σ_ϵ^2, respectively. Each of the variance components is estimated from

[23] Fuller and Battese [6].

the data, and these are used to transform the observations so that estimates will be efficient.

The Fuller and Battese estimates are reported in Table 4. A comparison with the cross-section estimates in Table 2 shows that $\hat{G}$ does indeed perform much better but that operating risk, $\hat{R}$, performs less well.[24] The more substantial variation in $\hat{G}$ between time periods than across industries likely accounts for its superior performance in the pooled equation. The less satisfactory performance of $\hat{R}$ in the pooled regression may have similar cause. That is, deviation from trend in EBIT may adequately measure difference in risk across industries but fail to measure more important changes in risk between the two time periods. We have no evidence, however, that this is indeed the case.

In Table 5, marginal effects on capital structure of changes in G and REG are reported, based on the Fuller and Battese point estimates of Table 4. Interpretation of the figures should proceed as follows. For $D^0 = .20$, a one-cent increase in the marginal annual tax advantage to interest implies an increase in interest by 0.796 percent of EBIT, to $D' \cong .208$.[25] Also for $D^0 = .20$, regulation would induce an increase in interest by 10.3 percent of EBIT to $D' = .303$. It is inherent in the logit transformation that marginal effects depend upon the value of the dependent variable. This is why Table 5 lists marginal effects for a range of initial values of D.

From Table 5 it is apparent that the effect of regulation is large. This effect cannot be attributed either to lower-than-estimated personal tax rates for shareholders in regulated firms, or to Herendeen-Sherman bias in rate-setting practices, for it is in the contrary direction. An appealing hypothesis is that regulation lessens the probability of failure, and that this insulation from business failure

[24] Regression equations like those of Table 2 (OLS) and Table 4 (GLS) were also estimated with $\frac{\hat{C}}{\hat{y}}$ included. This resulted in greater standard errors for coefficients on $\hat{G}$, $\hat{R}$, and REG, and coefficients on $\frac{\hat{C}}{\hat{y}}$ were not significant. For example, the Fuller and Battese (GLS) estimate with $\frac{\hat{C}}{\hat{y}}$ included is shown below, with t-statistics in parentheses.

$$\log \frac{\hat{D}}{1-\hat{D}} = \underset{(-8.614)}{-2.451} + \underset{(3.629)}{4.203\hat{G}} - \underset{(-0.458)}{0.107\hat{R}} + \underset{(1.147)}{0.436\text{REG}}$$

$$+ \underset{(0.935)}{0.038(\hat{C}_{t1}/\hat{y}_{t1})} + \underset{(1.114)}{0.023(\hat{C}_{t2}/\hat{y}_{t2})}$$

Only the cross-sectional variation in $\frac{\hat{C}}{\hat{y}}$ was included in the equation as its temporal change must largely be due to measurement error (see discussion in text). The coefficient on $\hat{G}$ has roughly the same value with $\frac{\hat{C}}{\hat{y}}$ included as when not; 4.203 compared to 4.973. Moreover, estimated marginal effects of a change in $\frac{\hat{C}}{\hat{y}}$ are not significant. What cross-sectional variation in $\frac{\hat{C}}{\hat{y}}$ does exist is apparently not systematically related to capital structure.

[25] The increase in the tax advantage to interest $\hat{G}$ between the two periods was approximately .10 for most industries. From the estimates in Table 5, this should yield increases in D in the range of .045 to .120 (controlling for operating risk). In contrast, the cross-sectional variation in $\hat{G}$ is of a much smaller order, and based on the same estimates, is unlikely to explain significant inter-industry capital structure differences. This corroborates the explanation given above for the superior performance of $\hat{G}$ in the pooled regressions to that in the cross sections.

leads regulated industries to exploit more fully the tax advantage to interest. Furthermore, as measured, this effect is the dominating influence of regulation on capital structure.

IV. Conclusion

Several theorists have suggested that the validity of the invariance proposition of corporate finance first posited by Franco Modigliani and Merton Miller depends upon the absence of taxes and failure costs. That is, taxes (specifically a tax subsidy to debt) and failure costs imply the existence of an optimal capital structure. In this paper we have attempted to determine empirically if these factors do explain cross-sectional and temporal variation in industry capital structure. In order to do this, though, it was first necessary to measure the size of the tax advantage to debt and the costs of failure.

To calculate the tax subsidy to debt, we considered the effect of changes in industry debt on personal as well as corporate tax liability. The bases for our calculation of personal tax liabilities were the income distributions of shareholders reported in Crockett and Friend [4] and Blume, Crockett and Friend [2]. Taking account of both corporate and personal taxes, we found that on the margin, the annual tax advantage to one dollar of interest generally ranged cross sectionally between $0.14 and $0.16 for the 1957–1964 period and between $0.23 and $0.26 for the 1965–1972 period. The marked increase in this tax advantage between the two periods was due to the decrease in personal tax rates occurring in 1964.

To determine failure costs, we followed an inferential approach that allowed measurement of indirect as well as direct costs. Approximately unitary elasticity was found between failure costs and income (EBIT). This finding is significant because theory implies that interest is homogeneous of degree one in failure costs and income. That is, variation in capital structure (as we defined it) ought not to be related to proportionate variation in failure costs and income.

Using our tax variable as well as a measure of operating risk and a dummy variable for regulation, we estimated cross-sectional and pooled (over our two time periods) regressions with industry capital structure as the dependent variable. Cross-sectional variation in capital structure was best explained by differences in operating risk, including that related to the regulatory process, and not by inter-industry differences in the tax advantage to interest, which were quite small. The general increase in interest as a percentage of EBIT that occurred between the earlier and later periods, however, was strongly related to the very great temporal increase in the tax advantage to interest that occurred between the two periods.

These findings add empirical support to theoretical assertions that taxes and failure costs do imply optimal capital structure, at least for industries.

Appendix A

Calculation of Marginal Tax Rate on Dividend Income, 1960 and 1971

This appendix details the method used to calculate the overall (mean) marginal tax rate on dividend income, τ_d. First, the tax rate for each type of dividend recipient was calculated.

For Dividends Received by Individuals

The average taxable income for each AGI class was drawn from *IRS Statistics of Income, Individual Income Tax Returns* for the appropriate years. Because members of an income class could file either joint or individual returns, the average taxable income on both types of return was calculated for each AGI class. These data were then used to construct weighted averages of the marginal tax rates on individual and joint returns appropriate for each AGI class.

For Dividends Received by Institutions

For tax exempt institutions, a marginal tax rate of zero was assigned. For estates and trusts, it was assumed that all dividends accrued indirectly to individuals in exactly the same proportions as dividends directly received by individuals. The marginal tax rate on these dividends is thus a weighted average of the individual tax rates.

For dividends received by corporations, 85 percent of the total is tax exempt, and the remainder is taxed at the rate of 48 percent. A fraction, $\alpha \simeq 0.6$ is then paid out to shareholders (including some corporations) as dividends. That paid to individuals is subject to personal taxes, whereas that paid to corporations goes through still another round in the tax cycle. Accordingly,

$$\tau_j = \frac{(.15)(.48) + 0.6 \sum w_i \tau_i}{1 - w_j},$$

where j = domestic corporations, w_i = percentage of dividends paid to i^{th} recipient type, and τ_i = marginal tax rate on dividend income for the i^{th} recipient type.

The marginal tax rates for the years 1961 and 1970 are listed in the last two columns of the appendix table. The percentages of dividends received by each class are listed in the first two columns. For both years, each tax rate is multiplied by its respective weight and these are summed to yield the overall marginal tax rates on dividend income.

Appendix A, Table I

Marginal Tax Rates and Percentage of Dividends Received by Each Class of Dividend Recipient, 1960 and 1971

Class of Dividend Recipient	w_i = Percentage of All Dividends (Other Than Own Stock) Distributed by Domestic Corporations		τ_i = Estimated Marginal Tax Rate	
	1960	1971	1960	1971
Institutions				
Domestic corporations	18.211	17.143	.32772	.28067
Corporate pension funds	2.678	7.724	0	0
State and local government retirement funds	0	1.036	0	0
Other tax exempt institutions (including those distributed through fiduciaries)	3.214	4.521	0	0
Estates and trusts, or utilized to pay taxes or administrative costs	2.500	5.212	.43051	.3846
Subtotal	26.603	35.636		
Individuals (by adjusted gross income)				
Under \$5,000	9.704	6.142	.20000	.14756
\$5,000–9,999	11.253	6.496	.20875	.18583
10,000–14,999	8.567	6.462	.22710	.19865
15,000–24,999	10.853	9.831	.31455	.25391
25,000–49,999	11.707	11.818	.44697	.36718
50,000–99,999	8.837	9.454	.66342	.53707
100,000–199,999	3.319	6.075	.76862	.62296
200,000–499,999	1.881	4.381	.89329	.70000
500,000–and over	7.276	3.705	.91000	.70000
Subtotal	73.397	64.364		
Total	100.000	100.000		

	1960	1971
Overall marginal tax rate on dividend income, $\tau_d = \sum_i w_i \tau_i$ =	.3864144	.2704001

Source: Constructed from Table 1 and Table G, Marshall E. Blume, Jean Crockett, and Irwin Friend, "Stock Ownership in the United States," *Survey of Current Business* November, 1974 and Table 1.2 and Table 1.3, Jean Crockett and Irwin Friend, "Characteristics of Stock Ownership," *Proceedings of the Business and Economics Statistics Section of the American Statistical Association*, 1963.

Appendix B

Definitions of Variables

Capital Structure

$$\hat{D} = \frac{\hat{I}}{\hat{Y}}$$

where $\hat{I}$ = real value of interest deductions, by industry; mean of annual observations. Source: *Statistics of Income, Corporate Income Tax Returns.*

$\hat{Y}$ = real value of income, before taxes or distributions to creditors or shareholders, by industry; mean of annual observations. Source: *Statistics of Income, Corporate Income Tax Returns.*

Marginal Annual Tax Advantage of Perpetual Interest Payment

$$\hat{G} = (1 - \hat{\tau}_d) - (1 - \hat{\tau}_c)(1 - \hat{\alpha}\hat{\tau}_d)$$

where $\hat{\tau}_d$ = marginal personal tax rate on dividend (and interest) income; see Appendix A for sources and method of calculation.

$\hat{\alpha}$ = industry (cash) dividend payout as a percentage of income less interest; mean of annual observations. Source: *Statistics of Income, Corporate Income Tax Returns.*

$\hat{\tau}_c$ = marginal corporate tax rate, assumed equal to 0.48.

Costs of Failure

$$\hat{C} = \frac{\hat{G}(1 - \hat{D})\hat{i}}{\beta_1(1 - \hat{\tau}_c)\hat{\Pi}(1 - \hat{\Pi})}$$

where β_1 = coefficient on $\log \dfrac{D}{1 - D}$ in equation [2].

$\hat{\Pi}$ = annual number of business failures by industry; Source: Dun & Bradstreet, *Quarterly Failure Report*: divided by number of corporations filing income tax returns; mean of annual observations.

$\hat{i} = \dfrac{\hat{I}}{n}$; where n = number of corporations filing income tax returns by industry; Source: *Statistics of Income, Corporate Income Tax Returns.*

Operating Risk

$\hat{R}$ = one minus the coefficient of determination in log linear trend regression on real value of industry operating income (Y).

Regulation

REG = 1 for regulated industries, 0 for all others. The regulated industries are: transportation; communication, and electric, gas and sanitary services.

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