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CHAPTER 11

An Economic Analysis of Pension Coverage and Plan Characteristics

In this chapter, characteristics of the pension market are studied, including pension plan coverage, pension amounts, type of plan coverage (including primary defined contribution coverage and profit sharing and stock bonus secondary coverage), funding levels, and plan terminations. Identification of the economic determinants of these phenomena helps explain some of the pension growth characteristics discussed in earlier chapters. In addition, the data provide an opportunity to test some of the predictions of the underfunding theory discussed in the preceding chapter. In fact, all of the empirical work is arranged around this theory: this provides a focal point around which the various statistical analyses can be evaluated. Put another way, recognizing that each set of statistical results may be interesting in themselves, the central question answered in this chapter is: Do the pension data support the underfunding theory presented in Chapter 10?

EMERGENCE OF PENSIONS

It is natural first to consider whether a union bonding theory of underfunding is consistent with historical evidence. Unfortunately, identification of the role of unions in the creation of pensions over time is difficult. Primarily in response to the National Labor Relations Act enacted in 1935, virtually all the growth in union membership took place between 1935 and 1945. In 1934, only 11.9 percent of the nonagricultural work force was unionized; by 1945, union coverage had

reached its historical peak of 35.5 percent.[1] The greatest growth in pension coverage began during this same period.

In 1940, approximately 4 percent of the work force was covered by pensions; by 1950, 25 percent was covered; by 1960, 40 percent.[2] But federal individual and corporate income tax policy also changed dramatically during the same period. In 1940, only 13 percent of workers paid income taxes; the median marginal tax rate was 4.4 percent. By 1945, 65 percent of workers paid income taxes; the median marginal tax rate was 23 percent. The corporate tax rate for large firms increased from 14 percent in 1935 to 40 percent in 1945.[3]

To the extent that firms experimented with appropriate safeguards to offset the effects of union growth during the post-1935 period, the tax preferred pension option would have provided a natural candidate in the new tax environment. But the tax changes alone would have been expected to produce rapid growth of pension plan coverage. It is difficult to disentangle these effects using time series data. It is noteworthy, however, that most pension plans took the defined benefit form, thereby awarding firms discretionary funding authority. They also awarded past service credits; that is, most emerging pension plans awarded pension credit for service rendered prior to the introduction of the pension plan.[4] This "gift" from stockholders to workers–a long-time puzzle in the literature–could be rationalized in the context of a bond theory; past service credits create an immediate incentive for workers to identify with the long-term viability of the firm.[5]

EXTENT OF BORROWING FROM WORKERS THROUGH UNDERFUNDED PENSIONS

Given the paucity of time series data, it is more fruitful to examine cross-sectional evidence to test more clearly the predictions of the bond theory. Since cross-sectional data are available only post ERISA, care

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[1]Union participation data is found in the Bureau of Labor Statistics, *Handbook of Labor Statistics, 1984.*

[2]The best estimates of pension coverage are found in Skolnik but his data begin only in 1950. See A. M. Skolnik, "Private Pension Plans 1950–1974," *Social Security Bulletin* 39 (June 1976), pp. 3–17. An alternative to this series can be constructed by evaluating the plan creation dates of a cross section of plans that file annual pension reports with the U.S. Department of Labor. Using 1978 plan participants as weights, relative pension growth over time can be calculated. This method, which is reflected in the numbers given in the text, provides similar data to Skolnik's during the years in which the two data series overlap.

[3]See Chapter 2, Table 2–4.

[4]The basic characteristics of major pension plans formed during the post–World War II period can be found in contemporary editions of Banker's Trust, *Corporate Pension Plan Study.*

[5]In the context of the model used above, new rounds of investment are made only after the union takes a significant part of its compensation in the form of an underfunded pension.

must be taken to consider the implications of the dynamics that characterized the market over time.

Prior to the episode of high inflation rates and the enactment of ERISA, the union bond was equal to the product of the underfunding ratio times the real pension liability $(1 - F)P$ (equation 10–2). If we rule out corner solutions, a combination of lower funding ratios and larger pensions provide the desired level of the bond. Consider first, the pension portion of the bond, P.

While inflation and ERISA lowered the desired level of underfunding, the importance of the pension benefit itself persists even after ERISA. That is, like the underfunding bond, the inflation bond is proportional to the level of the pension P (see equation 10–6). We therefore have our first unambiguous prediction for post-ERISA data. If pensions are used as a tool for bonding then, holding other things (especially wages) constant, unionized workers should receive a larger portion of their compensation in the form of pensions. They should more often be covered by pensions and, among pension participants, should implicitly save a larger fraction of their compensation in the form of pensions.[8]

The Importance of Pensions in Compensation

Probability of pension coverage. Consider the first proposition: unionized workers should more often be covered by pensions than nonunion members. To test this proposition, a logistic equation was estimated explaining the probability of pension coverage in terms of unionization and other worker and firm attributes. Data were taken from the May supplement to the 1979 Consumer Population Survey. The results are presented in Table 11–1. The results confirm some well-known characteristics of pension coverage. Pensions are much more frequently associated with older and higher-wage workers, especially in large firms.

The coefficient on the union variable is also positive, large, and highly significant. Holding wage, firm size, and other factors constant, union members are characterized by a pension coverage probability that is 38.5 percentage points higher than that for nonunion workers. Union members comprise 35 percent of the sample. The overall pension coverage rate in the sample is 54 percent. The results therefore suggest that, other things constant, nonunion workers are characterized by a 38 percent chance of pension coverage; union workers, by a 77 percent chance.

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[8]In this sense, the bond theory rationalizes the so-called preference of union members for pensions. See Richard Freeman, "Unions, Pensions and Union Pension Funds," Working Paper 1226, National Bureau of Economic Research, 1983.

TABLE 11–1 Determinants of Pension Coverage

Independent Variables	*Mean*	*Coefficient*	*Point Estimate*[a]
Intercept		−4.39	.012
		(11.86)	
UNION (= 1 if worker is in a union job)	.35	1.54	.385
		(21.56)	
Wage (hourly)	5.19	1.10	.091[b]
		(10.90)	
Wage^2	—	−.071	—
		(7.67)	
Age	37.74	.023	.005
		(10.08)	
Years of education	12.72	.046	.011
		(3.28)	
Female	.46	−.185	−.046
		(2.47)	
Nonwhite	.11	.046	.011
		(.48)	
Small firm (fewer than 100 employees)	.54	−1.66	−.415
		(15.51)	
Medium firm (100–1,000 employees)	.31	−.533	−.133
		(4.77)	
Occupational variables[c]	*X*		
Log of likelihood	7167.6		
Number of observations	7509		
Percent positive observations	.54		

The estimates reported in this table are based on a logistic equation. The dependent variable equals unity if the worker has pension coverage, zero otherwise. Numbers in parentheses are t-statistics. Data are from the May supplement to the 1979 Consumer Population Survey.

[a]The point estimate (PE) is the change in the probability of pension coverage generated by an increase in the dependent variable by one unit. It is easy to show that by differentiating the logistic equation, the point estimate for the jth variable equals $PE_j = b_j P(1 - P)$ where b_j in the estimated coefficient reported in the table and P is the relative frequency of positive values of the dependent variable; in this case, 54 percent of the observations are positives (see last line in table).

[b]The point estimate on wage accounts for the wage-squared term.

[c]Nine occupational dummy variables were included in the regression. Five were significant at the 95 percent level of confidence.

Generosity of pension plan. Now consider the second proposition: among pension-covered workers, unionized workers are more likely to receive higher portions of their compensation in the form of pensions. To examine this proposition, a small data base held by the Department of Labor was used. The data contain information about 1,670 workers from 61 pension plans (in 61 firms) who retired during the period 1975–1978. For purposes of this calculation, annual pension benefits were indexed to 1978 prices and multiplied by the number of years of expected pension collection. The expected years of collection was set equal to age 78 (expected age of death) minus age of retirement. The

TABLE 11–2 Variables Affecting Pension Share of Wage

Independent Variables[a]	*Coefficient*
Intercept	−3.08
	(1.16)
UNION (= 1 if plan participants are unionized)	.417
	(9.98)
Service (log)	.946
	(25.90)
Wage (log)	.804
	(16.86)
Firm size (employees):	
<250	.19
	(.76)
250–999	.03
	(.50)
1,000–5,000	.05
	(.85)
10,000–25,000	.40
	(5.69)
>25,000	.16
	(2.51)
Industry dummy variables[b]	*X*
R-squared	.603
Number of observations	1,670
Number of pension plans	61
Number of union plans	37

The estimates reported in this table are based on an ordinary least squares regression. The dependent variable is the log of the annual pension times age 78 minus age of retirement. Numbers in parentheses are t-statistics.

Data are taken from a U.S. Department of Labor sample of retirees from 61 pension plans who retired between 1975 and 1978.

[a]A dummy variable indexing whether the worker took joint and survivor protection was also included (which ordinarily carries an actuarial discount).

[b]Six industry dummy variables were included in the regression; three were significant at the .05 level of confidence.

calculation set the discount rate equal to zero.[7]

Normally, pension benefits are expressed in proportion to workers' final wage and service levels, suggesting a log-linear functional form. Dummy variables reflecting unionization, firm size, and industry were also included. The results are presented in Table 11–2. The coefficients on wage and service levels are close to unity, as they should be. In

[7]The calculation assumes that retirement benefits are indexed to inflation. In reality, pension benefits are indexed by only 50 percent of the inflation rate. See Robert Clark, Steven Allen, and Daniel Sumner, "Inflation Adjustments in Pension Plans," Report submitted to the U.S. Department of Labor, 1983. But since the study also shows that union pensions tend to be somewhat better indexed than nonunion plans, the assumption of perfect indexing for all plans tends to bias the union coefficient in a downward direction.

addition, the data show that large firms tend to offer larger pensions relative to wages compared to small firms. Most importantly, however, union plans in the sample are characterized by pension payments that are almost 50 percent (exp .417) higher than plans covering nonunion workers. Since pensions account for approximately 15 percent of compensation for pension-covered workers, the results suggest that in nonunion firms, they account for 12.5 percent of compensation; in union plans, 18.5 percent.[8]

The results from the coverage and the pension amount analyses suggest that pensions comprise a much higher percent of union workers' compensation compared to nonunion workers; unionization is affiliated with a doubling of the coverage probability and a 50 percent increase in pension amounts.

Funding Levels

We can now consider the second component of the underfunding bond: the underfunding ratio itself, $1 - F$. Two propositions from the bond theory can be explored. First, unionized workers should not be covered solely or primarily by fully funded, diversified defined contribution plans. Second, funding ratios in defined benefit plans should be lower when participants are unionized.

Defined contribution plans. Consider the first proposition: union workers should not be covered solely or primarily by defined contribution plans. Approximately 18 percent of pension plan participants are covered solely or primarily by defined contribution plans.[9] Recall that in defined contribution plans, the firm contributes a portion of the wage (or profits) into employee accounts each year. After a short vesting period (usually three to five years) workers, in effect, own the account. These plans are, by definition, fully funded. Virtually all primary defined contribution plans are diversified outside the plan sponsor's stock, and hence can be considered fully-funded accounts, giving workers virtually no exposure in their pension plan to bankruptcy risks in the firm.[10] Before or after ERISA, defined contribution plans have no bond value.

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[8]The coverage data used in Table 11–2 suggest that 40 percent of pension-covered workers are unionized. Pensions as a portion of lifetime compensation can be calculated from data found in the Social Security Retirement History Survey, 1970.

[9]See Chapter 6.

[10]For purposes of these estimates, a defined contribution plan was considered to be primarily invested in the plan sponsor's stock if more than 50 percent of the market value of the portfolio is comprised of the plan sponsor's stock. None of the results in the table changed when the 50 percent criteria was changed to 25 percent or 75 percent. Approximately 15 percent of all defined contribution plans are invested primarily in the firm sponsor's stock. But these plans are virtually always relatively small secondary plans. An empirical analysis showed that the probability of observing an undiversified account in the defined contribution population was not significantly related to union status of participants.

TABLE 11–3 Determinants of Defined Contribution Plan Coverage

Independent Variables	*Coefficient*	*Point Estimate*[a]
Intercept	−2.09	.11
	(3.37)	
Year plan created	.013	.001
	(1.44)	
Number of participants (1,000)	−.0004	0
	(.01)	
UNION (=1 if plan participants are unionized)	−2.86	−.30
	(5.78)	
Industry dummy[b] variables	*X*	
Log of likelihood	730	
Observations	1153	
Mean of dependent variable	.133	

The estimates in this table are based on a logistic equation. The dependent variable equals unity if the pension plan is a defined contribution type and is the sole or primary plan covering participants in the plan. Numbers in parentheses are t-statistics.

[a]See notes to Table 11–1.

[b]Nine industry dummy variables representing nine two-digit SIC industries were included in the equation. Two of these variables were significantly different from zero at the .05 level of significance.

SOURCE: Data is from the 5500 Annual Pension Plan Reports, 1978.

To address this issue, a logistic equation relating the probability of sole or primary coverage by a defined contribution plan against union status of participants was run. The results are presented in Table 11–3.

The results show that defined contribution plans are much less likely to be found when participants are unionized.[11] The point estimate on the union coefficient shows that unionized participants have a 30 percentage point lower probability of being covered solely or primarily by a defined contribution plan compared to nonunionized participants. Translating this result to absolute probabilities, it turns out that while participants in a nonunionized plan have a 25 percent chance of being covered solely by a defined contribution plan, participants in unionized plans have virtually no chance of being found with such coverage.[12]

[11]Union status of pension plan participants is found by merging the Form 5500 Annual Report data with the so-called EBS–1 tape which was filed by law in 1975 with the U.S. Department of Labor by all pension plans. Participants are assumed to be unionized if their plan was collectively bargained as reported on the EBS–1.

[12]The probability of prime or sole defined contribution coverage is approximately 18 percent. (See Chapter 6.) Unions are found in approximately 40 percent of pension plans. Using the point estimate on the union coefficient in Table 11–3, it turns out the probability of such coverage for unionized participants is negative .05.

Funding ratios in defined benefit plans. Consider now the much larger worker population covered solely or primarily by defined benefit plans. This subject was discussed briefly in Chapter 4 but now can be considered in more detail in the context of the underfunding propositions being examined here. In contrast to the pension benefit component, care must be taken to consider the dynamic aspects of the bonding mechanism on the underfunding ratio itself. After ERISA was enacted, underfunding suddenly became a costly pension attribute with no offsetting bonding benefits. If funding levels could feasibly and costlessly be altered instantaneously, post-ERISA data would not be useful for testing the funding level hypothesis. But IRS rules (and presumably financial constraints at the firm level) ensure that adjustments to shocks occur slowly.[13] Thus, by examining funding ratios on January 1, 1978 (the earliest date in which a funding snapshot is available), a large portion of pre-ERISA underfunding should be observable.

If underfunding was used prior to ERISA as a bonding device to control unions, then, using beginning-year data for 1978, it is expected that funding ratios for pension plans that cover unionized participants would still be systematically lower compared to pension plans that cover nonunionized participants.[14] To evaluate this proposition, an ordinary least-squares regression was run. The log of the funding ratio was the dependent variable. Union status of participants, the year the plan was created, and the size of the pension plan (number of vested participants) were included as independent variables. The results are reported in column 1 in Table 11–4.

The results confirm the bonding hypothesis in a striking way. The coefficient on union status of pension participants is negative, large, and highly significant: union plans have funding ratios that are more than 30 percent lower than nonunion plans. In the second column of the same table, results are presented for similar estimates, including variables that represent industry, occupational, and demographic characteristics of workers in the industry. While the coefficient on the union

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[13]Loss of bond value is not a legitimate reason to alter funding schemes in the eyes of the IRS. Instead, if the firm wishes to improve its funding position, the firm may simply change an actuarial assumption, say the assumed interest rate in the calculation. As a practical matter, the firm will be required to change assumptions moderately at any given time, and further, the actuarial shortfall attributable to the claim of "actuarial error" must be amortized over a 15-year period. While few would disagree that IRS rules award firms great latitude in choosing its funding position over the long run, sudden changes will not be accommodated by the IRS over the short run.

[14]The 1978 annual pension report year was the first year for which all data elements required to calculate funding ratios were keypunched. As many as 6,138 pension plans (with at least 100 participants) had the necessary data keypunched for that year. It is noted that to put all funding ratios on a comparable basis, conventional techniques were used to convert all reported liabilities to a common 2 percent discount rate assumption. See Chapter 4.

TABLE 11–4 Analysis of Funding Levels

Independent Variables	*(1)*	*(2)* Coefficients	*(3)*
Intercept	−.203	.66	.412
	(5.38)	(.80)	(.35)
Year plan created	.0023	−.0017	−.0018
	(3.86)	(2.68)	(2.52)
Plan participants (millions)	−.013	−.35	−.245
	(.01)	(1.17)	(.66)
Firm employees (millions)		.021	−.031
		(1.89)	(.24)
Plan growth[a]			.154
			(3.22)
UNION (=1 if plan participants are unionized)	−.306	−.238	−.262
	(23.70)	(15.91)	(13.57)
Right-to-work state			.091
			(4.11)
UNION · percent employees unionized			−.322
			(6.38)
UNION · industry pension level* (1,000)			.083
			(2.67)
UNION · industry tenure level*			.007
			(1.57)
UNION · industry labor share*			−.538
			(1.81)
UNION · industry growth[b] 1958–1972*			.073
			(2.47)
UNION · industry growth 1972–1981*			.092
			(1.14)
Industry characteristics (three-digit)			
Wage (1,000)		−.08	.24
		(.10)	(.20)
Percent union			−.153
			(2.04)
Demographic variables[c]		X	X
Occupational distribution variables[d]		X	X
Industry dummy[e] variables		X	X
R-squared	.085	.155	.189
Number of observations	6138	5375	4238

The estimates reported in the table are based on ordinary least squares regressions. The dependent variable is the log of the funding ratio. The funding ratio for all plans was evaluated using a common 2 percent interest rate assumption. Numbers in parentheses are t-statistics.

An asterisk (*) signifies that the variable itself was entered separately in the regression as an independent variable.

[a]Plan growth is measured as the percentage change in active participants in the plans from 1978 to 1981.

[b]Industry growth statistics are calculated at the four-digit SIC level. Growth is measured in terms of employment levels. Most industry data were available from 1958 through 1981 but approximately 20 percent of the SIC categories reported employment beginning in 1972; a dummy variable was also included (not reported) to denote these industries.

[c]The average years of education, average age of work force, average age of retirement, and percent of work force female by three-digit SIC industry classification were included. None of these variables was significant.

[d]Occupation variables describing the percent of the work force in each of five occupation classes were included as indicated. None of the coefficients was significant.

[e]Eight industry dummy variables were included as indicated. The omitted industry is manufacturing. Two of these coefficients are significantly different from zero in the second equation.

SOURCE: Data taken from the Form 5500 Annual Pension Plan Reports, 1978.

variable is reduced somewhat to approximately negative 25 percent, its size and significance essentially remain unaffected.

Additional tests that can further distinguish the bonding theory can be made. The optimal funding ratio is expected to be higher: the longer the natural length of tenure in the firm (see the role of R in equation (10–10) in Appendix 10–1),[15] the smaller the portion of unionized labor in the firm's work force (see expression 10–1),[16] and the smaller the share of labor relative to capital in the firm (see expression 10–1).[17] Moreover, recall that the underfunding bond is equal to the product of the underfunding ratio and the pension amount ($B = (1 - F)P$); hence the pension amount P is a natural substitute for the underfunding ratio of $1 - F$ at any given bond level B (see equation 10–2). As such, the higher the exogenously measured pension amount P, the higher the optimal funding ratio.[18]

Finally, two additional variables suggested by the theory but not explicitly considered above were also included in the regression.

First, the theory suggests that a negative correlation should exist in union plans between funding ratios and firm growth rates. Prior to ERISA, if unions are observed in declining firms and if a causal relation is assumed (see below), funding ratios should be observed to fall relatively rapidly. That is, once held up, the firm should respond within IRS constraints to reduce contributions to the plan to ensure a large

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[15]In brief, as long as wage-service profiles are upsloping, long-tenure workers tend to absorb higher opportunity wage penalties from firm failure (except those that will be retired at firm failure). In this sense, long-tenure jobs are natural substitutes for low funding levels. Tenure levels (for workers 50 years of age and older) are taken from the May 1979 Consumer Population Survey. By including only age 50 and over workers, a measure of long-term tenure in the three-digit industry is obtained; recall that industry wage and percent unionization are included separately in the regression.

[16]Recall that the firm can always fire unionized workers. This will be a relatively easy task as long as union workers comprise a relatively small portion of the work force. The easier this task, the lower the underfunding bond required to prevent a holdup. Usually, there is only one large union in the firm in the data base but sometimes there are several. It is simply assumed that if there is more than one union organization in the firm then they will act in concert to maximize their potential impact on the firm.

[17]This is just another measure of union workers' relative importance in the firm. The presumption is that if labor's share in relation to capital is small, presumably it is easier for the firm to continue production without the union. Hence, small labor-capital ratios reduce the ultimate power of the union and hence reduce the size of the bond. Labor share of value added is taken from census data for the 1975 year on a three-digit basis.

[18]This variable is measured by the industry pension level. The calculation is made by cumulating all pension payments from the annual pension reports, and sorting by three-digit industry. Dividing the sum by the number of retirees collecting pensions (also sorted by three-digit industry), a measure of the level of industry benefits was derived. To purge the measure of potential bias, average industry wage and percent of industry unionized were also included in the regression. These variables are taken from the May 1979 Consumer Population Survey.

bond loss when failure occurs. Post-ERISA, this special union relationship should evaporate.[19]

Second, a separate dummy variable was included to denote if the firm was located in a right-to-work state. Potential unionization is expected to be more difficult in a right-to-work state.[20] Similarly, if already unionized, it may also be more difficult to extract rents from stockholders because it may be easier to replace an exploitive union. Hence, bonding should be less valuable in a right-to-work state.

The results of this extended regression are shown in the third column of Table 11–4. To facilitate interpretation, all variables interacted with UNION were subtracted from their mean value; thus, the union dummy variable measures the impact of unionization on funding levels in the case where all interaction variables are equal to their mean values. The coefficient on the union dummy variable is −.262 virtually identical to its value in the second regression. Moreover, all of the signs on the coefficients on the union interaction terms are of the expected sign, and five out of the six coefficients expected to be different from zero are significant at least at the 95 percent level of significance (one-tail test). The insignificant post-ERISA growth-union interaction coefficient is expected to be zero. In short, the results reported in Table 11–4, taken as a whole, are strongly supportive of the bonding hypothesis.

In fact, the results suggest that virtually all underfundedness found in defined benefit plans can be attributed to the presence of unions. It has been shown above that a reasonable estimate of the overall long-term funding ratio is approximately 75 percent.[21] The results suggest that other things constant, nonunion plans are approximately 90 percent funded while union plans are on average 60 percent funded.[22]

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[19]Since ERISA, all failing firms have the same incentive to reduce funding once pending failure becomes apparent. In this way, the value of the federal insurance is maximized. Good measures of firm growth are not available. But since unions often are present across four-digit industries, the four-digit industry employment growth rates are a good measure of union impact. These measures were conveniently available for two periods which are approximately pre- and post-ERISA, 1958–1972 and 1972–1981. Interacting these variables with the union variable, we would expect that the pre-ERISA variable would be significantly positive, not the post-ERISA measure.

[20]While it is only possible to observe whether employees in a firm are unionized or not, presumably a possibility exists for nonunion employees to unionize *if* the stockholders provided an opportune commitment to long-term capital investment.

[21]Over the period 1950 to 1981, the average funding level was 68 percent. The target ratio could be somewhat higher, in the vicinity of 70 percent, perhaps as high as 85 percent. Thus 75 percent is a reasonable measure of long-term funding ratios in the United States (see Chapter 4).

[22]The overall funding ratio is defined as $F = (A_u + A_n)/(L_u + L_n)$ where $A_u(A_n)$ are assets held by union (nonunion) plans and $L_u(L_n)$ are liabilities held by union (nonunion) plans. By using the relations, $F_u = A_u/L_u$ and $F_n = dF_u$, where d is estimated in Table 11–5, and known asset weights, it is straightforward to solve for F_u and F_n.

EMERGENCE OF PROFIT SHARING STOCK BONUS PLANS AND TERMINATION OF DEFINED BENEFIT PLANS

At least three additional tests are suggested by the bond theory. First, if ERISA eliminated the incremental bond value of underfunding, firms should react by imposing substitute bonds. Second, if unions engage in opportunistic behavior, occasional errors in calculation of the optimal bond value alone will generate a causal relationship between unionization and firm failure; this relation will be exacerbated after ERISA renders positive underfunding bonds valueless. Third, if the bond theory provides a plausible rationale for the enactment of ERISA, the beneficiaries of ERISA transfers should be comprised predominately of participants in unionized plans.

Creation of Substitute Bonds

Recall the simple model used above. If a firm holds capital that has a remaining life of $t^* - t$ periods when ERISA is enacted, a union holdup may not be optimal. But during the next round of investment, the horizon again returns to its initial value, t^*. If the inflation bond alone is insufficient to prevent opportunistic behavior at that time, a supplementary bonding mechanism will be necessary. For the purposes of testing for this substitution, we will consider profit sharing (PS) and stock bonus (SB) plans as imperfect but plausible substitutes for underfunding (see above). Since the *de facto* underfunding bond conferred by inflation still exists after ERISA, the creation of a supplemental plan tied to firm profits could plausibly replace the incremental bonding created by underfunding prior to ERISA.

To evaluate this hypothesis, a random sample of defined benefit plans that existed prior to ERISA was chosen from the annual pension reports filed with the U.S. Department of Labor. If participants in a plan were covered by a supplemental profit sharing or stock bonus plan by 1981, the creation date of the supplemental plan was recorded on the defined benefit plan record. By using the plan dates available on the records, the history of SB and PS supplemental creation could be constructed. This history was broken into four periods. That is, a pseudo-longitudinal data series was comprised of plan information pertaining to four periods—pre-1960, 1960–1967, 1968–1973, and 1974–1981. A zero-one index was created for each defined benefit plan. The index was converted from its initial value of zero to unity if a PS or SB plan was started during the period; upon attaining a value of unity, the observation was dropped for subsequent periods. Thus, if a PS or SB plan was created to supplement a defined benefit plan in 1966, the value of the index for this plan was set to zero for the pre-1960 period, unity for the 1960–1967 period, and then it was dropped from the

TABLE 11–5 Probabilities of New Profit Sharing or Stock Bonus Plan Coverage over Time

Period	*Non-union*	*Union*	*Union + Nonunion*
Pre-1960	.086	.076	.88
1960–1967	.060	.047	.78
1968–1973	.049	.030	.61
Post-ERISA	.076	.098	1.28

Numbers in this table are probabilities that a PS or SB supplemental plan is created during a period, given that participants are already covered by a defined benefit (but not PS or SB) plan at the beginning of the period. These estimates are derived from the results of a logistic equation estimate of PS or SB plan coverage reported in note 23.

sample. Defined benefit plans that never had a PS or SB supplemental plan created had zero entries in all four periods.

The dependent variable was therefore the probability that a profit sharing or stock bonus plan was created during the period, conditional on the firm existing and having not previously created such a plan. A zero-one union dummy variable was interacted with each time period. It is expected that relative to nonunion participant groups, unionized participants would experience a significant increase in profit sharing or stock bonus coverage after ERISA.

A logistic equation was run using these data.[23] The results of these estimates are presented in Table 11–5. The results show that in the pre-1960 period, the probability that an SB or PS plan was created to cover nonunion participants was 8.6 percent; the probability for nonunion participants was 7.6 percent, only 88 percent of the nonunion probability. During the subsequent two periods, the probability of new PS or SB coverage fell for both groups, but fell faster for union participants. That is, during the period 1960–1967, the probability of PS or SB plan creation was only 78 percent for unionized participants compared to

[23] The logistic estimate was:

$$\begin{aligned} PSSB_t = &-2.36(24.3)[.086] - .53(4.8)[-.026]\cdot D(1960\text{–}1967) \\ &- .75(6.8)[-.037]\cdot D(1968\text{–}1973) - .21(2.1)[-.01]\cdot D(1974\text{–}1981) \\ &- .195(2.0)[-.010]\cdot UNION - .058(1.3)[-.003]\cdot UNION\cdot D(1960\text{–}1967) \\ &- .173(2.3)[-.009]\cdot UNION\cdot D(1968\text{–}1973) \\ &+ .632(3.6)[.032]\cdot UNION\cdot D(1974\text{–}1981), \end{aligned}$$

where $PSSB_t$ equals unity if a SB or PS plan was created during period t, zero otherwise; UNION equals unity if the defined benefit plan covered union workers, zero otherwise; and $D(t)$ is a dummy variable that equals unity during period t, zero otherwise. The numbers in parentheses are t-statistics and those in brackets are point estimates (PE). There were 22,357 observations; the log of the likelihood was 9198. The numbers in Table 11–5 reflect a summation of appropriate point estimates. For example, the union entry in the post-ERISA period is equal to .086 (the intercept PE) − .010 (the UNION PE) − .01 [the $D(1974\text{–}1981)$PE] + .032 [the UNION·$D(1974\text{–}1981)$PE], or .098.

nonunion participants, only 61 percent as high during the period 1968–1973. But this pattern changed sharply after ERISA. The data show that post-ERISA, the probability of new PS or SB coverage increased for both groups but much more so for union participants. Post-ERISA, the conditional probability of new PS–SB coverage increased to 9.8 percent for unionized workers, 3.26 times the probability for the preceding period; it was 7.6 percent for nonunion workers, only 1.55 times the probability for the preceding period. The relative union/nonunion creation probability ratio increased from 61 to 128 percent. The data are consistent with the notion that ERISA caused more initial PS–SB coverage for union compared to nonunion workers.

Determinants of Firm Failure

Consider now the issue of firm failure. Even pre-ERISA, a bond model would predict a causal relation between firm failures and worker unionization. Even though firms set bonds, as long as there is some distribution of errors in calculation or perception, the union will presumably be observed on occasion forsaking the bond in favor of extorting the firm. ERISA exacerbated the problem by rendering the underfunding bond valueless thereby increasing the probability of a holdup post-ERISA.

To investigate this hypothesis directly, the 1978 pension plan data base was used. Since the enactment of ERISA, all pension plans that terminate must file a termination notice with the federal government; these filings provide a trail of failed firms that previously operated a pension plan. By matching plans in such failed firms since 1978 with the pension plan population in existence in 1978, all plans about to fail during the period 1978 to 1983 can be identified. Of 4,007 plans (that had at least 100 participants) from the stratified sample, 58 were affiliated with about-to-fail firms.

The probability of failure was analyzed from this sample data base. A dummy variable UNION FIRM was set equal to unity if at least one union pension plan was found in the firm; the percent of employees accounted for by union employees in the firm was also included; its value was normalized to the mean union share in unionized firms (equal to .4). Firm size, date of plan creation, and industry growth statistics were also included in the estimates. The results are reported in the first column of Table 11–6.

The results confirm the union-bonding hypothesis. The probability that the typical plan in the sample would be found in a firm that would fail in the ensuing five years was 1.4 percent. If the firm was characterized by a work force that was 40 percent unionized (the average for union firms in the sample), the failure probability was virtually unaffected by union presence. But if unionized employees represented 80

TABLE 11–6 Determinants of Firm Failure and Plan Termination

Independent Variables	Probability of Firm Failure[a]	Probability of ERISA Transfer[b]
Intercept	−5.35	−4.94
	(6.01)	(6.09)
	[.0047]	[.007]
Date of plan creation	.0076	.0028
	(.58)	(.23)
	[.0001]	[.0001]
Employees in firm (1,000)	−.07	−.097
	(.01)	(.02)
	[−.001]	[−.004]
UNION	—	1.02
		(2.68)
		[.046]
UNION FIRM	.44	.25
	(.32)	(.55)
	[.006]	[.011]
Percent employees unionized	1.88	—
	(3.13)	
	[.026]	
Industry growth 1958–72	.005	−.140
	(.02)	(.33)
	[.00007]	[−.006]
Industry growth 1972–81	−3.27	−3.24
	(3.67)	(3.68)
	[−.045]	[.146]
Log of likelihood	541.4	672.1
Observations	4007	4243
Mean of dependent variable[c]	.014	.045

The estimates reported in the table are based on logistic equations. Numbers in parentheses are asymptotic t-statistics; numbers in brackets are point estimates (see notes to Table 11–1).

[a]Dependent variable: probability that a plan observed in 1978 was in a firm that failed by 1983.

[b]Dependent variable: probability that a plan observed in 1978 was terminated by 1983 by an ongoing or failed plan sponsor and received a net transfer from the Pension Benefit Guaranty Corporation.

[c]Samples are deliberately stratified to overrepresent failed firms and transfer plans.

percent of the firm's employment, the probability of firm failure increased by over 70 percent; that is, the probability of failure is increased by an additional 1.04 percentage points. In short, while the sample of failed firms is small, the data are consistent with the notion that firm failure is affiliated with unionized work forces, especially if unions represent a large share of the firm's employment.

Determinants of ERISA Transfers

Now consider the ERISA insurance transfer experience. It was argued earlier that the bond model naturally leads to the proposition that unionized workers would expect to be the principal beneficiaries of

ERISA insurance; therefore, these workers would have a strong incentive to work for its enactment. It is a matter of public record that unions favored the enactment of ERISA.[24] If the arguments and the model are plausible, we should be able to find verification that unionized workers indeed have been the dominant transfer recipients. Recall that transfers are received when poorly funded plans are terminated as a result of firm dissolution and when the firm decides to transfer its pension liability to the federal insurance corporation in exchange for the assets in the plan plus 30 percent of the firm's net worth.

The same matching process between termination records and the 1978 plan population data that generated a picture of the firm failure experience also produces a record of plans that existed in 1978 but subsequently terminated and received a subsidy from the federal insurance corporation; 191 such cases were found. The probability that a plan in the stratified 1978 sample collected a transfer was analyzed in a way similar to the failure probability. The results are shown in the second column of Table 11–6.

The results are quite dramatic. While a plan in a unionized *firm* is not especially likely to receive an insurance transfer, a unionized *plan* is much more likely to be a transfer recipient. While the overall probability of receiving a transfer in the sample is 4.5 percent, a unionized plan has an *incremental* probability of receipt of 4.6 percentage points.

CONCLUSION

The bond theory not only rationalizes firms' choices to underfund their pension plans, but it also explains many characteristics and recent developments in the pension industry, hitherto unexplained: the creation of "past service credits" when most major pension plans were created; the supergenerosity of union pension plans and their superunderfunded status; and the absence of primary defined contribution plan coverage in the union sector. The theory is also consistent with firms' peculiar acquiescence to ERISA insurance and unions' strong preference for it; in the post-ERISA period, it is consistent with the preponderance of union plans claiming federal insurance transfers, as well as with the abnormally high incidence of profit sharing and stock bonus plan creation in the union sector.

The empirical work reported in this chapter confirms some of the well known relations that characterize the pension market: the role played by wage level in pension coverage and pension share of wages, the role of industry growth on funding levels and plan terminations,

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[24] See, for example, U.S. Senate, *Private Welfare and Pension Plan Legislation*, Hearings before the General Subcommittee on Labor (Washington, D.C.: U.S. Government Printing Office, 1970).

and the like. But some novel results were also reported, notably the strong role played by the union variable in virtually all of the empirical analyses. Unions are strongly related to coverage, pension amounts, funding levels, plan-type coverage, and terminations. Among other things, for projecting the pension market, if union representation of the U.S. work force continues to fall, important changes could take place in some pension aggregates. For example, to the extent that underfunding is explained by the union-bonding theory, the creation of ERISA will ultimately result in higher funding levels in union plans, simply because ERISA erased the bonding characteristics of underfunded plans, though there may be some tendency in the opposite direction.[25] This trend is reinforced by the gradual reduction of union importance in the economy. On the other hand, since unions are affiliated with abnormally high coverage and pensions, the increased funding implied by the demise of unions could be offset by reduction in total pensions paid; hence, the net effect on asset growth is ambiguous.

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[25]Some of these plans may deliberately reduce funding levels to take advantage of the federal government insurance program.