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UNIONS, PENSION WEALTH, AND AGE-COMPENSATION PROFILES

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ABSTRACT

This paper examines the effect of unions on both the magnitude and distribution of pension benefits. Our empirical results show that beneficiaries in collectively bargained plans receive larger benefits when they retire, receive larger increases in their benefits after they retire, and retire at an earlier age than beneficiaries in other pension plans. As a result, the pension wealth of union beneficiaries is 50 to 109 percent greater than that of nonunion beneficiaries.

Just as wage differentials within and across establishments are smaller among union workers, benefit differentials within and across cohorts of retirees are smaller among union beneficiaries. This results from the smaller weight given to salary average in determining initial benefits and the larger percentage increases given to those who have been retired the longest under post-retirement increases. The more compressed benefit structure under unionism causes the union-nonunion compensation (wages plus pension contributions) differential to decline more quickly than the union-nonunion wage differential over the life cycle.

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I. Introduction

Studies of the rationale for pensions and analyses of their economic effects have been one of the growth sectors in the field of labor economics over the last five years. Reasons for this surge of interest are fairly obvious--pensions are now a major component of employer expenditures for compensation and have important effects on work effort and labor mobility. Much of the initial impetus toward the growth of pension plans in the 1940s and 1950s came from organized labor. Although pension plans have also become increasingly widespread in the nonunion sector, union membership remains one of the dominant factors in determining pension coverage. According to Kotlikoff and Smith (1983), 76 percent of private wage and salary workers belonging to unions or employee associations are covered by pension plans compared to only 35 percent of nonunion workers.

This paper examines the effect of unions on both the magnitude and distribution of pension benefits. Previous studies have dealt with the former in two different ways. First, Freeman (1983) has estimated the difference in employer contributions to pension plans between union and nonunion establishments over two different periods. Across all establishments, Freeman found contributions were four cents per payroll hour higher between 1967 and 1972 and eight cents higher between 1973 and 1977 when 50 percent or more of the employees were covered by a collective bargaining agreement. Almost all of the difference can be attributed to the greater likelihood of union employees participating in a pension plan. Among only those establishments making contributions to pension plans, contributions were estimated to be only 0.3 cents higher between 1967 and 1972 and 0.2 cents higher between 1973 and 1977 in unionized establishments. Neither of these latter two point estimates was larger than its standard error.

On the surface, Freeman's results suggest there should be little difference in pension benefits between union and nonunion beneficiaries. Such a conclusion depends, however, on whether there are any differences in (1) hours worked and the probability of receiving pension benefits between union and nonunion workers and (2) rates of return and funding ratios between plans administered for union and nonunion establishments. Since there are good reasons to doubt there are no union-nonunion differences in these factors, the effect of unions on pension benefits actually received cannot be deduced from the finding of no difference in hourly contribution rates. This is not to suggest differences in employer contributions are not an important issue; for analyses of labor demand and the effect of unions on cost and profitability, it is the issue. But to understand the sorting of workers between union and nonunion jobs and the role of unions as "voice" institutions, a second approach is required--examination of actual benefit data.

Leigh (1981) attempted to do this by analyzing responses in the National Longitudinal Survey for older men to the question "How much income per month will you get from your pension plan?" Leigh's OLS results showed that, among workers expecting to receive benefits, the expected pension benefit for union workers is $72 per month lower than that of nonunion workers. Correcting for selectivity bias in provision of pension plans, he found that expected monthly benefits are $52 lower for union workers. Interpretation of these results hinges on whether expectations of pension benefits are equally accurate for union and nonunion workers.[1] The "voice" model predicts union workers will have more accurate expectations which is consistent with Leigh's findings of (1) a larger proportion of nonunion workers who said they did not know how much their pension benefits would be and (2) a larger coefficient of variation of expected monthly pension income for nonunion workers who thought they knew the magnitude of their benefits. Although this does not necessarily imply that nonunion workers either over- or under-estimate their benefits, data on actual benefits clearly are free of any such bias. In addition, Leigh did not take into account possible increases in benefits after retirement. As we have shown elsewhere (Allen, Clark, and Sumner, 1984b), these increases are much larger in plans covered by collective bargaining agreements.

The issue of how unions affect the distribution of benefits among retirees has not been addressed in previous studies. One of the "stylized facts" about the effect of unions on wages is that they reduce wage differentials within and across establishments. This paper examines whether they have a similar compressing effect on the distribution of pension benefits.

This question has important implications for our understanding of the union-nonunion compensation differential over the life cycle. Union wage policies raise the intercept and flatten the slope of wage functions, which implies that the compensation differential falls with experience. Freeman has shown how the greater pension coverage of union workers works in the opposite direction by causing the compensation differential to widen with experience, a consequence of the greater rate of accrual of benefits in the later stages of the life cycle. A complicating factor ignored by Freeman but addressed here is whether, among those eligible for benefits, the increment in pension wealth from an additional year of experience varies by union status.

This issue also has direct implications for retirement incentives. If the growth of pension wealth with experience is, say, smaller in union plans, union workers will have a smaller incentive to continue working. The effect of unionism on retirement age also depends on the union-nonunion differential in the stock of pension wealth.

The ways in which unionism is likely to affect pension benefits are examined in the next section. After briefly describing the data set, we report our main empirical results. They show beneficiaries in collectively bargained plans receive larger initial benefits, retire earlier, and get larger post-retirement increases in benefits. The regression results also point out important differences in the determinants of initial benefits and post-retirement increases by union status. The paper concludes by calculating and interpreting the effect of unions on pension wealth and examining the implications of the results for life cycle compensation profiles.

II. How Unions Affect Pension Benefits

The likely effect of unions on pension benefits is discussed in Freeman (1981 and 1983) and Leigh (1981), so we will provide only a brief summary of their arguments. Since our data set is restricted to a sample of beneficiaries, we focus on arguments dealing with the magnitude of benefits rather than on those that focus only on the question of which firms have pension plans. The standard monopoly model predicts both higher pensions and higher wages under unionism, but makes no prediction about how the share of pensions in total compensation differs between union and nonunion plans. Freeman (1981) argues the share of all fringe benefits will be higher under collective bargaining because preferences of older workers receive greater weight and older workers have greater demand for benefits. Older workers are especially likely to have greater demand for pensions. Also, when there is a divergence between what workers want and what management thinks they want with respect to the magnitude of pension benefits, this information will flow more rapidly to management under collective bargaining because of the "voice" aspects of union behavior. The market "exit" mechanism in nonunion settings will be less effective because the younger and more mobile marginal worker will put much less weight on pensions than will the older, inframarginal worker. Thus, because of greater demand for pensions and the mechanisms for revealing information under unionism, we expect retirees from collectively bargained plans to receive larger benefits.

Another consequence of this model is that the sorting of workers with heterogeneous tastes between union and nonunion establishments results in the marginal union worker having a higher supply price of fringes. Young workers who have little use for pensions will avoid jobs in the union sector because they know their preferences will be dominated by those of older workers. If the marginal worker expects longer tenure in a union setting, this will, by increasing the likelihood he will receive a pension, make him even more willing to sacrifice wages in return for future benefits before he is vested. Finally, the ability of unions to monitor the firm and the pension fund is an additional factor leading to higher benefits under unionism. The complexity of benefit formulas and pension plan provisions for such factors as participation, vesting, and portability makes it very difficult for workers to evaluate their plan or compare it to plans at other establishments. Unions hire experts to examine pension plan provisions and pension fund performance, giving workers both more information about their plan and, in all likelihood if expert opinion is translated into bargaining objectives, a more desirable plan. If this additional information results in greater worker confidence in the plan, it will increase demand for pension benefits.

The effect of unionism on pension wealth cannot be estimated accurately in a simple cross section benefit equation for two reasons. First, union participants are more likely to be eligible for early retirement and disability benefits. Kotlikoff and Smith (1983) report 86 percent of the participants of union plans are covered by early retirement provisions and 79 percent are covered by disability provisions. The corresponding figures for nonunion plans are 79 and 70 percent. Of course, greater eligibility for early retirement need not translate into an earlier average retirement age, depending on how the benefit formula is adjusted for early retirees. Nonetheless, we expect earlier retirement in union plans because if union participants receive larger benefits, the resulting wealth effect leads them to spend part of those benefits on more leisure.[2] If union participants retire at an earlier age, the union coefficient in the benefit equation underestimates the true impact of unionism on benefits. To adjust for this, we estimate the impact of union status on age of retirement in a separate equation. (There also may be longevity differences between union and nonunion beneficiaries, but we have no information on the direction of this potential source of bias.)[3]

The second complicating factor is post-retirement increases in benefits. Such increases can be provided only if participants accept lower wages or lower initial benefits or if they result in more efficient separation decisions. In practice, there is usually no formal pension plan provision for post-retirement increases, leaving the nonunion firm a tremendous incentive to renege on any implicit contract to provide such increases. This also makes such implicit contracts less likely among nonunion plans.[4]

Potentially, unions can act as enforcement agents to prevent cheating by the plan and increase the likelihood of post-retirement adjustments. Even though such adjustments are not a mandatory bargaining topic, unions have ample means to pressure employers to discuss the matter. Whether this is in the union's interest is an empirical question. Many workers have the same incentive as the firm to violate arrangements made with retirees because they can use the strike threat to obtain a share of the capital gains. On the other hand, other workers will not want to forfeit the option of using post-retirement adjustments as a risk-sharing device when they retire, especially if they already have accepted lower wages in expectation of receiving future benefit increases. This is especially likely to be true for older workers. The median voter model predicts and Freeman's (1983) findings on pension plan provisions imply that preferences of older workers receive much more weight in forming union objectives. This makes it more likely that unions will act in their interest.

Another factor encouraging unions to act in this fashion is the activity of retirees in union political affairs. In some unions retirees can vote for officers and attend conventions. In the United Mine Workers they even vote on contract ratification. This means distributing to retirees a portion of any rents obtained in negotiations can yield a political payoff to union officers. In contrast, retiree preferences receive zero weight in a nonunion setting, making an intergenerational transfer from workers to retirees unlikely.

We have shown elsewhere (Allen, Clark, and Sumner, 1984b) that union beneficiaries are much more likely to receive post-retirement increases, controlling for other factors such as plan size, salary average, initial benefits, and years of service. At the sample means, the average union participant received a 33 percent increase between 1973 and 1979; the average nonunion participant, a 17 percent increase. Empirically, this calls for a distinction to be made between initial benefits and the rate of change in benefits after retirement to get an accurate reading on the impact of unionism.

In addition to changing the size of the pension package, unions also are likely to change the manner in which benefits are allocated. Many studies have shown that unions compress occupational and life cycle wage differentials. One reason frequently offered for this practice is that it promotes solidarity among union members. A wide dispersion can lead to the creation of dissident groups at either tail or in the middle of the wage distribution. By collapsing the distribution toward the middle, the odds of being able to achieve further redistribution for any particular group become quite small. To examine whether this argument has equal validity when applied to the distribution of pension benefits, we will estimate separate initial benefit and post-retirement increase equations for union and nonunion beneficiaries.

III. Data Description

This analysis uses data from the Pension Benefit Master File (PBMF), made available to the authors through a contract with the Department of Labor. This is a survey based on a stratified random sample of pension plans filing series 5500 and 5500C forms in 1975. The PBMF contained firm-reported information on individuals receiving benefits in December 1978. Plan sponsors were asked to indicate the size and method of all post-retirement increases in benefits between 1973 and 1978. Using this information we were able to construct a benefit series that indicated the annual pension benefits from 1973 to 1979 for all persons retired prior to 1973. In addition, we were able to determine the benefit at retirement for persons retiring between 1973 and 1978 and any post-retirement increases for these more recent retirees between their year of retirement and 1979.

Although the PBMF included defined contribution plans as well as defined benefit plans, this analysis concentrates exclusively on the defined benefit plans. The defined contribution plans were excluded because of limitations in the survey that make it impossible to determine benefit increases between 1973 or the year of retirement and 1979. As a result, we could not calculate initial benefits or post-retirement adjustments for individuals covered by the defined contribution plans. Weights provided by the Department of Labor enabled us to construct weighted samples of individuals and plans representative of the set of defined benefit plans that existed in 1975. For a more detailed description of this sample and our data construction techniques, see Clark, Allen, and Sumner (1983) and Allen, Clark and Sumner (1984a, b).

Some plans did not report collective bargaining status on the PBMF. To reduce the number of cases for which collective bargaining status was unknown, we matched the PBMF with a file of the EBS-1 reports obtained from the National Bureau of Economic Research. For the few cases in which these sources differed on collective bargaining status, the data on the EBS-1 reports were, on the advice of the Department of Labor, assumed to be correct. Even with the addition of the EBS-1 data, the collective bargaining status of some plans remained unspecified. Rather than deleting these observations entirely, we use two union status variables in the pooled samples. The first indicates whether the plan was collectively bargained; the second, whether collective bargaining status was unreported. This allows the coefficient of the first variable to be interpreted as a union-nonunion difference. When separate equations are estimated on the basis of union status, we do not examine the plans that have union status unreported.

Although the size of the samples used in the empirical work reported below is quite large in terms of numbers of individuals (about 100,000), only about 200 pension plans are represented. As a result, the precision of the union coefficients will be overstated to some extent. The magnitude of this overstatement cannot be determined, as most of the other variables in the model vary across individuals rather than pension plans. Aggregating the observations by plan does not solve the problem because this eliminates most of the variation between union status and the other independent variables, thus biasing the union coefficient. Even though our approach is not completely satisfactory, keep in mind that the precision of union coefficients in wage equations is overstated in exactly the same way. A relatively small number of union contracts and large nonunion companies will account for a disproportionate share of the observations in most data files commonly used by labor economists today.

The PBMF contains data sufficient to examine the relationship among different types of compensation that include salary average, initial pension benefits, and post-retirement adjustments. This paper focuses on the effect of unions on each of these forms of compensation. First, we examine a sample of persons who retired between 1973 and 1977. Using this sample, we examine the effect of unions on benefit levels and union-nonunion differences in factors that determine benefit levels at retirement. Post-retirement adjustments are also estimated for these retirees. Second, we examine a sample of persons who retired between 1950 and 1972 for differences in post-retirement adjustments attributable to union status.

IV. Empirical Results

In this section, we report the findings from a statistical analysis of union effects on pension benefits. The results indicate that union retirees receive higher initial benefits and greater post-retirement adjustments than nonunion retirees.

Initial Benefits, 1973-77 Retirees.

Benefits at retirement are determined by plan formulas that generally are of three types: (1) a flat dollar amount for all eligible beneficiaries, (2) a flat dollar amount times years of service, or (3) a percentage of final salary average times years of service. Kotlikoff and Smith report the largest (47 percent) proportion of participants in collectively bargained plans have their benefits determined by the second type of formula, with 12 and 21 percent covered by the first and third types of formulas. In nonunion plans 71 percent have their benefits determined by the third type of formula, with 0.2 and 20 percent covered by the first and second types of formulas. Because the formulas are so different between union and nonunion plans, it is impossible to determine the impact of unions on benefits simply by comparing the formulas. Also, formula comparisons require arbitrary assumptions about union-nonunion differences in years of service, earnings, and age of retirement. Accordingly, we estimate initial benefit equations as a function of years of service, age at retirement, and average earnings over the last five years of service.[5] In addition to these formula variables, we include in our specification dichotomous variables for race, sex, year of retirement, industry, and union status. A plan size variable that represents the number of beneficiaries in 1979 also is included in the specification to control for positive correlation between size and union status. After experimenting with a number of functional forms, we decided to specify both initial benefits and salary average in logarithmic form.

Columns 1 and 2 of Table 1 show results of the initial benefit equations when the logarithm of average salary during the last five years of work is excluded (column 1) and included (column 2). Initial benefits clearly are influenced by earnings. First, there is the direct positive relationship noted above in which benefits are calculated as a percentage of final earnings. Second, there is a positive relationship that stems from high-wage workers desiring larger pensions because the after-tax price in the form of wage reductions will be lower. Also, there is the simple positive correlation that results if fringe benefits have a positive income or total compensation elasticity. Finally, there is an inverse relationship stemming from a compensating differential, i.e., higher pension benefits are paid for by lower wages. These simultaneous relationships are not directly addressed by our empirical model.[6]

When the log of average earnings is included as an independent variable, the estimated coefficients indicate that a 10 percent increase in average final earnings raises initial benefits by approximately 3 percent. As for the other coefficients, a 10,000-person increase in the number of 1979 beneficiaries is associated with a 6 percent rise in initial benefits. An extra year of service adds 5 percent to the benefit level. A somewhat puzzling result is the finding that delaying retirement by one year reduces benefits by 1 percent. This may be attributable to the fact that more generous pension plans usually allow earlier retirement.

The key result in Table 1 is that initial benefits are larger for union beneficiaries.[7] Without controlling for earnings in column 1, we find that union beneficiaries receive 4 percent larger initial benefits. The difference widens to 6 percent when the earnings average variable is added to the model in column 2. The coefficients imply that a white male beneficiary previously employed in manufacturing who retired in 1977 with the sample mean values of earnings average, plan size, age at retirement, and years of service would receive an initial benefit of $2411 from a nonunion plan and $2568 from a union plan.

There are two puzzling aspects about these results. First, why is the effect of unions on pension benefits so much smaller than almost all estimates of their effect on wages? Second, why doesn't the union-nonunion difference in benefits shrink in column 2 when earnings are added to the model? Assuming positive correlations between (1) unionism and wages and (2) wages and benefits, the union coefficient in column 1 should have been larger than that in column 2. These apparently contradictory results could be attributed to the sample selection rule from which they are generated. This is a sample of workers who received pensions from defined benefit plans; workers who were not covered by pension plans, workers covered by defined contribution plans, and workers who were covered by defined benefit plans but did not qualify for benefits are omitted.

To get a better feeling for how to interpret our results, we estimated an earnings average equation. The results in column 3 of Table 1 show beneficiaries in collectively bargained plans had a 7 percent lower earnings average in their last five years of employment. At first glance, this seems to contradict years of research on union wage effects. Several factors relating to our data set

Table 1. Benefit at retirement, earnings average and age at retirement equations, 1973-77 retirees

Dependent variable		ln(Benefit at retirement)	ln(Benefit at retirement)	ln(Earnings average over last five years before retirement)	Age at retirement
Independent variable	Mean	(1)	(2)	(3)	(4)
Union	.680	.040 (.006)	.063 (.006)	-.067 (.004)	-.840 (.034)
ln(earnings average)	8.971		.345 (.005)		.410 (.027)
ln(benefit at retirement)	7.677				-.306 (.018)
1979 recipients/10^5	.138	.627 (.013)	.576 (.012)	.148 (.008)	-2.910 (.069)
Years of service/10^2	.245	5.547 (.030)	5.230 (.030)	.920 (.020)	4.724 (.188)
Age at retirement/10^2	.622	-.835 (.058)	-.986 (.057)	.438 (.038)	
White	.918	.008 (.009)	-.022 (.009)	.087 (.006)	.182 (.048)
Male	.779	.323 (.006)	.195 (.006)	.370 (.004)	-.149 (.036)
R^2		.408	.438	.170	.094

Note: Each equation also contains an intercept and dummy variables indicating industry (6), year of retirement (4), whether union status is unreported, and whether sex is unreported. The sample size is 99680 in each equation. Since large plans were intentionally overrepresented in the survey, each observation is weighted by plan weights representing the incidence of similar plans in the pension universe.

account for this apparent contradiction. First, the sample selection criteria exclude many low-wage nonunion workers. Second, the PBMF does not contain any human capital variables except years of service. Thus, we are unable to account for differences in earnings due to education or occupational differences. The sample selection criteria and lack of human capital measures suggest that we are comparing blue collar union workers with white collar nonunion workers who have more years of schooling. These factors are less important in the benefit equations because of the inclusion of earnings as an independent variable. The union-nonunion wage differential also may be understated because of using only the last five years of earnings to derive the earnings average variable. Other studies have shown that the wage differential narrows with age, so we are estimating the union effect at a point in the life cycle at which it is expected to be relatively small.

An extra year of service raises final earnings by 0.9 percent, indicating that earnings rise until retirement in this sample. Consistent with this response is the finding that later retirement increases final earnings. Delaying retirement by one year increases final earnings by 0.4 percent. Increases in plan size raise final earnings. This conforms to expectations if plan size is a proxy for the size of the firm. The point estimate indicates that an additional 10,000 beneficiaries increases final earnings by 1.5 percent.

To estimate the effect of unions on pension wealth, we must also determine whether unionism has any effect on age of retirement. To get a rough estimate of this effect, we regressed age at retirement on union status, earnings average, initial benefits (a proxy for pension wealth), and the other independent variables used in the benefit and earnings average equation. We find that union beneficiaries retire almost one year earlier than nonunion beneficiaries. At the sample means of all independent variables, the average union beneficiary retires at age 61.9; the nonunion beneficiary, at age 62.7.

Although our model in column 4 is much less econometrically elaborate than most retirement models, results for the earnings average and pension benefit coefficients are comparable to previous results. We find persons with higher earnings in their last years of work and smaller initial benefits more likely to work longer. However, the magnitude of both coefficients is rather small. A one-unit increase in log earnings average (equivalent to thousands of dollars) is associated with only a 0.4-year delay in retirement; a one-unit decrease in log initial benefits corresponds to a 0.3-year delay.

Differences in benefit formulas used in collectively bargained plans suggest that determinants of pension benefits vary greatly by union status. Table 2 reports the results of benefit at retirement equations when the sample is sorted by union status. Recalling differences in the benefit formulas, we expect that increases in earnings will have a larger effect in nonunion plans. This prediction is confirmed by the results shown in Table 2. In addition, the plan size effect is twice as large for the nonunion sample. The only qualitative difference between the two equations concerns the age of retirement variable. Nonunion plans have the expected relationship of higher benefits for delayed retirement, whereas members of union plans have lower benefits with delayed retirement.

A final relationship is shown below when type of retirement (normal, early, postponed or disability) variables are included in the regression equations. This does not produce any major changes in the other independent variables (including age at retirement), so they are not repeated here. Union retirees who retire early receive 10 percent more in benefits than those retiring at the normal age whereas nonunion retirees receive 10 percent less. Union members taking a disability retirement receive 6 percent more than those retiring at the normal age, but nonunion disability retirees receive 18 percent less. Postponed retirement for both groups leads to lower benefits.

Table 2. Benefit at retirement equations, all 1973-77 retirees by union status

Sample Independent variable	Union Mean	Union Coefficient (Standard error)	Nonunion Mean	Nonunion Coefficient (Standard error)
Intercept		4.308 (.046)		.272 (.192)
ln(earnings average)	8.989	.280 (.004)	8.930	.546 (.014)
1979 recipients/10^5	.191	.567 (.009)	.028	1.072 (.131)
Years of service/10^2	.254	4.550 (.027)	.227	6.018 (.080)
Age at retirement/10^2	.616	-1.338 (.046)	.633	.579 (.204)
White	.901	.053 (.007)	.952	-.122 (.033)
Male	.853	.152 (.006)	.613	.284 (.016)
R^2		.543		.420
N		81828		17725
Mean of dependent variable		7.794		7.460

Note: The dependent variable is in logarithmic form. Each equation also contains an intercept and dummy variables indicating industry (6), year of retirement (4), and whether sex is unreported. The same weighting procedure is used here as in Table 1.

Type of retirement	Estimated coefficient in logarithm of benefits at retirement equation	
	Union	Nonunion
Early	.098 (.005)	-.103 (.016)
Postponed	-.430 (.021)	-.053 (.028)
Disability	.062 (.009)	-.181 (.029)

Post-Retirement Increases

Until recently, it was widely believed that private pension benefits were fixed in nominal terms. Allen, Clark, and Sumner (1984a,b) present evidence that this has not been true. Instead there were sizable increases in benefits after retirement during the mid-1970s. Column 1 of Table 3 reproduces a post-retirement adjustment equation (Allen, Clark, and Sumner (1984b)) for persons retiring between 1950 and 1972. The second column shows the results from a similar equation for persons retiring between 1973 and 1977. Because there are fewer years for potential increases for the second sample, and because increases tend to be larger in percentage terms for those retired the longest, we expect the coefficients to be smaller for this sample of more recent retirees.

The union coefficient indicates that the 1950-72 retirees in collectively bargained plans received increases which were 16.9 percentage points larger than their nonunion counterparts. For the 1973-77 retirees, the union differential is only 5 percentage points. A large plan size effect is found for both samples, with a 10,000 beneficiary increase resulting in a 4 percentage point increase for the older retirees and a 1 percentage point increase for the more recent retirees. For the older sample, additional years of service increase

Table 3. Post-retirement increase equations, 1950-72 and 1973-77 retirees

Sample Independent variable	1950-72 retirees Mean	1950-72 retirees Coefficients (Standard errors) (1)	1973-77 retirees Mean	1973-77 retirees Coefficients (Standard errors) (2)
Union	.724	.169 (.006)	.681	.050 (.002)
1979 recipients/10^5	.170	.439 (.011)	.136	.121 (.004)
Years of service/10^2	.247	.360 (.026)	.245	-.038 (.011)
Age at retirement/10^2	.625	-.023 (.050)	.622	.221 (.021)
White	.930	-.054 (.008)	.919	.007 (.003)
Male	.756	-.211 (.005)	.764	-.049 (.002)
R^2		.069		.052
N		137038		103579
Mean of dependent variable		.308		.075

Note: The dependent variable is the ratio of the change in benefits between 1973 and 1979 to 1973 benefit for 1950-72 retirees; the ratio of the change of benefits between the year of retirement and 1979 to benefits at retirement for 1973-77 retirees. Each equation also contains intercept and dummy variables indicating industry (6), year of retirement (22 in column 1, 4 in column 2), whether union status is unreported, and whether sex is unreported. The same weighting procedure is used here as in Table 1.

post-retirement adjustments, whereas for those that retire at younger ages, larger increases are provided. The relationships for years of service and age at retirement did not hold for the more recent retirees.

How were benefit increases distributed to retirees by union and nonunion plans? The frequency distribution of benefit increase formulas is reported in Table 4. A few plans did not use the same type of benefit increase formula each time they gave an increase. For instance, a firm may have given the same percentage increase to all retirees in 1974, but larger percentage increases to those retired the longest in 1977. As a result, the columns in Table 4 are non-additive. There is only one entry per plan per type of benefit increase formula in the table. A plan will be represented more than once only if it uses more than one type of benefit increase formula.

Only 16.9 percent of the plans gave straight percentage increases. Both union and nonunion plans tended to favor other approaches. Half the nonunion plans increased benefits by a percentage, with the percentage increasing with the amount of time the person had been retired. Most union plans gave either the same dollar amount increase to all retirees or gave a fixed dollar amount per year of service. The former method is parallel to compressed occupational wage differentials in the union sector. The latter reflects the political dominance of senior workers in union decision making. These two approaches also produce larger percentage increases for those who have been retired the longest, as long as their benefits are lower than those of more recent retirees.

The net effect of union-nonunion differences in benefit increase formulas can be gauged by estimating separate benefit increase equations by union status over 1950-72 retirees receiving at least one increase between 1973 and 1979. These are reported in the appendix. The most interesting difference between the union and nonunion benefit equations is the pattern of the year

Table 4. Benefit increase formulas, by union status

Formula	Total plans	Union plans	Nonunion plans	Percent of total plans	Percent of union plans	Percent of nonunion plans
Percent increase	13	8	5	16.9	17.8	15.6
Percent increase, percentage increasing with time retired	27	11	16	35.1	24.4	50.0
Percent increase, percentage increasing with CPI	6	2	4	7.8	4.4	12.5
Flat dollar increase	14	13	1	18.2	28.9	3.1
Flat dollar increase per year of service	22	16	6	28.6	35.6	18.8
Flat dollar increase, amount increasing with time retired	3	2	1	3.9	4.4	3.1
Flat dollar increase per year of service, amount increasing with time retired	3	2	1	3.9	4,4	3,1
Other	4	3	1	5.2	6.7	3.1
Total	77	5	32	100.0	100.0	100.0

Note: Some plans used more than one benefit increase technique over this period, making the columns non-additive. Only plans with persons retired before 1973 are included in the sample.

of retirement dummies, presented in Figure 1. These dummies indicate that the increases awarded to longer term retirees are relatively much greater for union than for nonunion beneficiaries, a possible response to union efforts to compress the compensation and/or benefit distribution. Union beneficiaries retiring in the 1950s received benefit increases 47 percentage points larger than those received by union beneficiaries retiring in 1972. This difference was only 18 percentage points for nonunion beneficiaries. A similar pattern holds for those retiring in the 1960s, as union beneficiaries received increases 27 percentage points larger than those of 1972 retirees, whereas nonunion beneficiaries who retired in the 1960s received benefit increases only 9 percentage points larger. In contrast, 1970 and 1971 retirees in both union and nonunion plans received increases only slightly larger than those of 1972 retirees. The smoother time pattern of increases in benefits seems attributable to a larger and more even distribution of beneficiaries across the years of retirement in the union sample.

V. Unions and Pension Wealth

Empirical results in the last section show union beneficiaries retire earlier and receive both larger initial benefits and post-retirement increases in benefits. To compare pension wealth for union and nonunion retirees, two sets of calculations were performed. In the first set, we used mean values from the union (nonunion) sample of years of service and salary average to calculate initial pension benefits for union (nonunion) workers. To isolate the effect of unionism on pension benefit formulas, we performed a second set of calculations using the mean values of years of service and salary average for the entire sample to obtain initial benefits for both union and nonunion beneficiaries. In both sets of calculations, the initial benefits for union (nonunion) beneficiaries are derived from the coefficients of the equation estimated over the

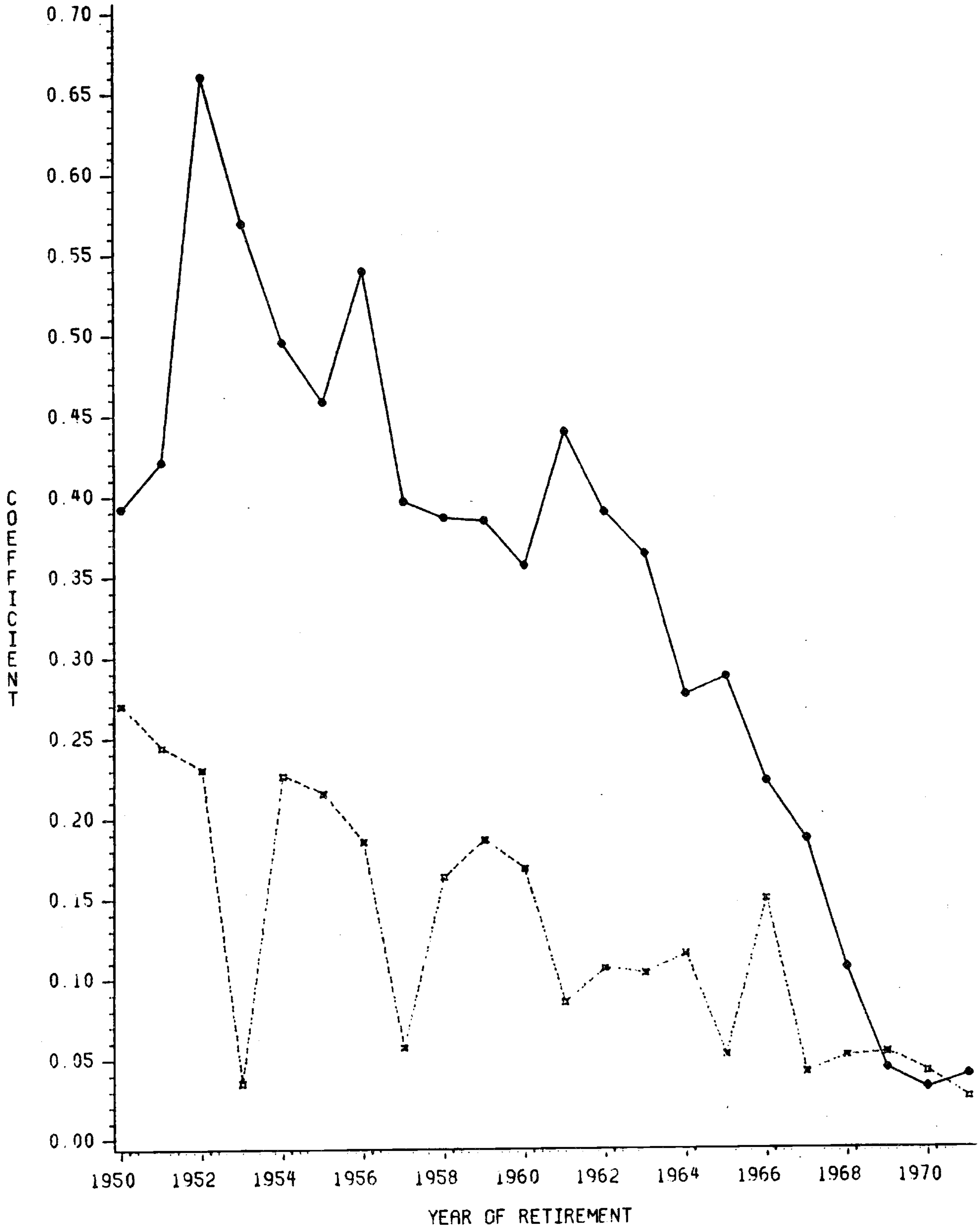

Figure 1. Coefficients of year of retirement dummies in post-retirement increase equations: *solid line*, union beneficiaries; *dashed line*, nonunion beneficiaries

union (nonunion) sample. We assume union beneficiaries live for 18 years after retirement in converting this initial benefit estimate into a pension wealth estimate. When retirement ages of union and nonunion beneficiaries are held constant, this same assumption is used for nonunion beneficiaries. We use a discount rate of 3 percent. For age of retirement we assume that either (1) it is identical for union and nonunion beneficiaries or (2) it is .84 years later for nonunion beneficiaries, reducing the time in which benefits are received by an equal amount. The latter assumption is based on the age at retirement regression results in Table 1.

For post-retirement increases we examine four cases: (1) none are given; (2) only union beneficiaries receive them; (3) all union and nonunion beneficiaries receive them; and (4) increases are given by union and nonunion plans with an adjustment for the greater proportion of union beneficiaries receiving increases.[8] The magnitudes of the post-retirement increases are derived from the results in Appendix Table 1. For the first seven years after retirement, we use the predicted value for a person retiring in 1972 (.249 for union retirees; .190, nonunion). For the next five- and six-year periods, we use the predicted values for persons retiring in 1967 and 1961, respectively (.435 and .688 for union; .235 and .275, nonunion). The underlying assumption is that both the magnitude of increases (in percentage terms) and the distribution of increases across cohorts of retirees are constant over time. Based on our examination of a sample of union contract histories reported in BLS Wage Chronologies, this does not seem to be an unreasonable assumption for union beneficiaries. Under these assumptions, benefits increase by 202.5 percent over an 18-year period for union beneficiaries receiving increases; 87.4 percent over 18 years for nonunion beneficiaries. This is equivalent to an annual increase of 6.34 percent for union beneficiaries, 3.55 percent for nonunion beneficiaries.

Let's turn now to the pension wealth comparisons in Table 5. The first row examines the case in which there are no post-retirement adjustments.

Pension wealth is 50 percent greater for union beneficiaries when we allow years of service and salary history to vary by union status. If we set years of service and salary history to be equal for union and nonunion beneficiaries, the effect of unionism on pension wealth falls to 8 percent. The latter figure is slightly larger than the amount implied by the union coefficient in the initial benefit equations in Table 1 because it is derived from the specification with complete union status interactions in Table 2.

The effect of unionism on pension wealth becomes much larger when we take into account post-retirement adjustments. The most reliable comparisons of pension wealth are reported in the last row of Table 5. Allowing for post-retirement increases in benefits raises the effect of unionism from 50 to 98 percent, allowing for union-nonunion differences in years of service and salary average; from 8 to 43 percent, controlling for these differences. Many nonunion beneficiaries (42 percent) received no increases. When we compare pension wealth for nonunion beneficiaries who do not receive post-retirement increases with that of union beneficiaries who do, these gaps widen further to 162 percent allowing for union nonunion differences in years of service and salary average; 89 percent, controlling for these differences.

Allowing for differences in retirement age results in a modest increase in our estimates of the effect of unionism on pension wealth. The difference increases by 6 to 17 percentage points when retirement for nonunion beneficiaries is delayed by .84 years.

Considering all three effects simultaneously and adjusting for the greater proportion of union beneficiaries receiving increases, pension wealth is 50 to 109 percent higher for beneficiaries in plans covered by collective bargaining. This range is well above all reasonable estimates of the effect of unions on wages, implying wage differences vastly understate the effect of unions on total compensation among persons eligible for benefits.

Table 5. Estimated pension wealth ratios for union and nonunion retirees, by assumptions about post-retirement benefit increases and age of retirement

Assumption about post-retirement increases		Means of independent variables vary by union status: Assumption about age of retirement		Sample means of independent variables: Assumption about age of retirement	
		No difference by union status	Nonunion retirement .84 years later	No difference by union status	Nonunion retirement .84 years later
No post-retirement increases	Union	\$39706	\$39706	\$36502	\$36502
	Nonunion	26379	24768	33627	31573
	Ratio of union to nonunion	1.50	1.60	1.08	1.16
Post-retirement increases given only by union plans	Ratio of union to nonunion	2.62	2.79	1.89	2.01
Post-retirement increases given by union and nonunion plans without adjustment for greater proportion of union beneficiaries receiving increases	Ratio of union to nonunion	1.91	2.00	1.38	1.44
Post-retirement increases given by union and nonunion plans, with adjustment for greater proportion of union beneficiaries receiving increases	Ratio of union to nonunion	1.98	2.09	1.43	1.50

These results are especially puzzling once we reconsider Freeman's (1983) finding of no union-nonunion difference in employer contributions to pension plans per hour among establishments making such contributions. If union employers contribute no more than nonunion employers, where does the money to finance larger benefits come from? There are four possibilities: (a) larger total contributions in union plans (in contrast to employer contributions per hour), (b) the smaller percentage of employees receiving pensions in union plans; (c) higher rates of return in union plans; and (d) lower funding ratios for union plans.

a. Contributions. There is no evidence that union members work more hours than nonunion workers. Although we know of no data about how the magnitude of employee contributions varies by union status, Freeman has examined whether union or nonunion plans are more likely to allow such contributions. Controlling for a number of plan characteristics, he found union plans are 21 percent less likely to involve voluntary employee contributions. The greater incidence of employee contributions in nonunion plans implies larger benefits in nonunion plans.

b. Percentage receiving pensions. This depends upon turnover rates and vesting requirements in union and nonunion establishemnts with pension plans. Mitchell (1982) found much lower quit probabilities for union workers, even with controls for pension coverage. Freeman (1980) indirectly controls for the effect of all fringe benefits on tenure (which depends on layoff, discharge, and quit rates) by using the omitted variable bias formula. Ignoring fringes, he finds tenure is 1.06 years higher for union workers. The difference narrows to 0.79 years once fringes are controlled for. This suggests contributions per eventual beneficiary are higher in nonunion plans because participants are less likely to stay long enough to collect benefits. Offsetting this are the stricter vesting requirements in union plans. However, almost all persons receiving benefits have much more than the minimum years of service required for vesting. Kotlikoff and Smith report 9.7 years are required for full vesting in union plans; 9.0 years, in nonunion plans. Both figures are well below the mean years of credited service in our sample.

c. Rates of return. There has been a great deal of controversy for some time about whether union plans earn the highest possible rates of return. One source of concern has been corruption among trustees appointed by union officials in Taft-Hartley plans, which are jointly administered by labor and management. Cases in which pension fund assets have been invested in very speculative ventures (frequently associated with organized crime) or used for personal gain by union officials have been widely publicized.

More recently union officials have become interested in using pension funds to advance a variety of social causes, including the organization of union workers. By refusing to invest in certain companies because they do business in South Africa or have violated labor laws, some union plans may be unable to put together the most desirable portfolio of assets. Whether this is actually the case is an empirical question. So far no one, to our knowledge, has produced any evidence that such restrictions on possible portfolios result in lower rates of return.

Some plans knowingly have invested in projects offering lower returns to increase the utilization of their members, and possibly, add new members. This practice generates additional contributions to the fund, but it also may create additional future liabilities, depending on the benefit formula, rules for vesting, and whether the project increases employment or hours of union workers. The net effect of such practices on the amount of funds available in future years thus is unclear.

Despite all of these possible sources of lower returns to collectively bargained plans, available evidence indicates that returns to union and nonunion plans are not very different. Munnell (1983) reports the results of a comparison made by the A. G. Becker Co. of the median rate of return for Taft-Hartley plans to all plans between 1973 and 1982. Over the entire period, a plan earning the median rate of return of the Taft-Hartley plans would have grown by 69 percent, equivalent to an annual return of 5.4 percent. A plan earning the median rate for all plans would have grown by 64 percent, equivalent to a 5.1 percent annual return. Although Taft-Hartley plans fared better on average, the difference is very small and seems to be primarily attributable to smaller holdings of equities in union plan portfolios in 1973 and 1974, two especially disastrous years for the stock market.

d. Funding ratios. If the larger pension wealth for union beneficiaries does not come from greater contributions per eventual beneficiary or from a higher rate of return, we are left with a final possibility--the funding status of collectively bargained plans. Despite the tax advantage of full funding described by Black (1980) and Tepper (1981), many plans are not fully funded.[9] One reason for this, Sharpe (1976) argues, is that fully funded plans lose the option of allowing the Pension Benefit Guaranty Corporation to assume the liability for unfunded benefits not covered by 30 percent of the sponsor's net worth.

A recent study by Richard Ippolito (1983) of the U.S. Department of Labor produces strong evidence that plans covered by collective bargaining agreements have much lower funding ratios than nonunion plans. Using a smaple of 826 defined benefit plans filing reports with the Department of Labor in 1978, Ippolito found the funding level of union plans (with respect to vested labilities) 31 percentage points below that of nonunion plans. Based on Ippolito's estimate of a 60 percent funding ratio in 1978, the results suggest, holding plan size, industry, year created, and employment growth in industry constant, the average nonunion plan is 78 percent funded, whereas the average union plan is only 47 percent funded.

There is good reason to doubt the accuracy of the absolute magnitudes of these figures because they are based on an interest rate assumption of 2 percent. At least among the large firms represented on the COMPUSTAT data files, most plans are overfunded, especially when vested liabilities are evaluated at the same interest rate and that interest rate is at a reasonable level, currently about 6 or 7 percent (see Feldstein and Morck, 1983).

Although we are not currently prepared either to explain or to estimate the average funding ratios of union and nonunion plans, consider two possible scenarios, both of which are consistent with Ippolito's regression results: (1) union plans tend to be fully funded whereas nonunion plans are over funded and (2) union plans are underfunded, whereas nonunion plans are fully funded. Under the first scenario, the seeming contradiction between higher pension wealth for union beneficiaries despite identical or lower contributions per eventual beneficiary is easily resolved--part of the nonunion contributions is being used to maintain the overfunded status of the plan. In other words, nonunion plans are contributing more than necessary to fund current and future benefits, presumably because of the resulting tax advantages. Why aren't managers for collectively bargained plans following the same strategy, especially those in which none of the trustees are appointed by labor? The reason is that unions have the ability at regular intervals to push contract negotiations for higher initial benefits and increases in benefits for those already retired. The strike threat gives the union power to convert an overfunded plan to a fully funded or underfunded one. (Unions cannot bargain directly over funding ratios.) Knowing this, managers of collectively bargained plans put no more money than necessary into the fund. According to this scenario, then, unions are able to constrain the behavior of fund managers so that all contributions eventually end up in the hands of beneficiaries instead of the stockholders or management.

In the second scenario, union plans are underfunded and the burden of higher benefits falls on younger workers. Identical contributions per hour can provide larger benefits for union beneficiaries under this scenario as long as the ratio of participants to beneficiaries is sufficiently high. Younger workers are willing to participate in an unfunded plan as long as (1) they receive rents and

(2) they expect another generation of workers to pay for their benefits when they retire. Such an arrangement is not viable for nonunion plans. The nonunion firm always has an incentive to terminate an underfunded plan. By doing so, it can avoid paying the difference between vested and funded benefits. There is one catch, however, pointed out by Bulow (1982) that keeps nonunion plans fully funded--workers will not stay with a firm when the plan is not fully funded unless the firm offers a compensating wage differential. A questionable aspect of this scenario is whether the participant-beneficiary ratio has been or will remain high enough to guarantee survival of the system. Ippolito accounts for greater underfunding in the union sector via an entirely different mechanism--employers intentionally underfund union plans to discourage the union from threatening the firm's financial health. If union behavior causes the firm to go out of business, employees then lose a substantial proportion of their pensions. However, the threat of job loss seems to be equally credible in this regard. Further empirical work will be needed to establish which of these scenarios and interpretations are consistent with the data.

VI. Unions and Age-Conpensation Profiles

Many studies have shown that tenure-earnings profiles for union workers have higher intercepts and flatter slopes. Do pension accruals imply that the total compensation profile also has this pattern? Freeman shows how the greater coverage of union workers by pension plans and the larger increments in pension wealth in the last years of the life cycle in defined benefit plans widens the wage difference among older workers. This partially but not totally reverses the flattening of the profile. Here we go one step further and compare tenure-earnings profiles and tenure-compensation profiles for union and nonunion workers covered by pension plans where compensation equals earnings plus pension contributions.

We use earnings functions of the form

$$w(t) = w(0)(1+g)^t,$$

where $w(0)$ equals \$10,000 for union workers and \$8,000 for nonunion workers, g equals .03 for union workers and .033 for nonunion workers, and t represents tenure and is allowed to be as long as 30 or 40 years. The pension benefit formulas are derived from the coefficients in Table 2 using the means for the pooled union and nonunion sample of plan size, industry, year of retirement, race, sex, and age at retirement. These formulas are:

$$\log B_u = 4.257 + .0455t + .280 \log (\bar{E})$$

$$\log B_N = 1.429 + .0602t + .546 \log (\bar{E}),$$

where B_u = initial benefit for union workers, B_N = initial benefit for nonunion workers, and $\bar{E}$ = earnings average over the last five years. We assume workers are fully vested in their tenth year and have zero vesting beforehand. Pension wealth estimates are based upon 18 years of retirement and a 3 percent discount rate. The pension contribution for each year equals the increase in pension wealth minus the return (also 3 percent) on last period's pension wealth. The latter figure must be subtracted because it would have accrued even if the participant did not work.

Results for 40-year life cycles are reported in the first seven columns of Table 6. By construction, the percentage difference in earnings between union and nonunion workers falls from 25 percent in the first year to 12 percent in the fortieth year. This pattern is similar to those of previous studies (summarized in Lewis, Ch. 7). Pension contributions begin in the tenth year. The contribution for union workers is 33 percent larger in that year because the union benefit formula has a larger intercept. This figure is much larger than the union-nonunion earnings diffential in the tenth year (22 percent). The total compensation differential based on both earnings and pension contributions, is 26 percent. This is considerably larger than the total compensation differential (based on earnings only) of 22 percent in the year preceding vesting.

Table 6. Earnings and pension contribution profiles, by union status and length of profile

	40-year profile							30-year profile			
	Earnings			Pension contribution			Earnings and pension contribution	Pension contribution			Earnings and pension contribution
Tenure	Union	Nonunion	Ratio	Union	Nonunion	Ratio	Ratio	Union	Nonunion	Ratio	Ratio
1	10000	8000	1.25				1.25				1.25
5	11255	9109	1.24				1.24				1.24
9	12668	10373	1.22				1.22				1.22
10	13048	10715	1.22	8553	6423	1.33	1.26	11495	8632	1.33	1.27
11	13439	11069	1.21	487	536	.91	1.20	654	720	.91	1.20
15	15126	12604	1.20	679	824	.82	1.18	913	1108	.82	1.17
20	17535	14825	1.18	1030	1411	.73	1.14	1385	1896	.73	1.13
25	20328	17438	1.16	1563	2414	.65	1.10	2100	3245	.65	1.08
30	23566	20512	1.15	2371	4133	.57	1.05	3186	5554	.57	1.03
35	27319	24127	1.13	3596	7073	.51	.99				
40	31670	28379	1.12	5456	12107	.45	.92				

Note: The simulations assume firms fund only vested liabilities. These are fully funded each year and are included as a part of total compensation.

In all following periods, larger contributions are required for nonunion workers. This happens because increases in average earnings and years of service have larger effects on pension benefits for nonunion workers and because earnings grow more rapidly for nonunion workers. Even though union workers have higher earnings, the increment in pension wealth resulting from rising earnings over the life cycle is rather small because of the compressed benefit structure. As a result, after the tenth year, the union-nonunion compensation differential is smaller than the earnings differential. In the 35th year and thereafter, total compensation is greater for nonunion workers. The figures for the 30-year life cycle case in the last four columns tell basically the same story--<u>among those receiving pension benefits, the pension benefit structure under unionism causes the compensation differential to decline more quickly than the earnings differential over the life cycle</u>.

VII. Conclusion

The main finding of this paper is that pension wealth for beneficiaries in plans covered by collective bargaining agreements is significantly larger than pension wealth for other beneficiaries. Three factors contribute to this union-nonunion pension wealth differential: higher initial benefits, larger post-retirement increases in benefits, and earlier receipt of benefits for those in collectively bargained plans.

This finding is somewhat surprising, given the earlier results of Freeman and Leigh on union-nonunion differences in the magnitude of pension contributions and expected pension benefits. The key factor in reconciling Freeman's results with ours seems to be the lower funding ratios in plans covered by collective bargaining agreements. Except in the case of Taft-Hartley multiemployer plans, unions are unable to control directly a plan's financial management. Funding ratios and investment decisions generally are not subject to collective bargaining; unions can influence these decisions only indirectly through bargaining for changes in the benefit formula. The question of whether and why collective bargaining status seems to have an important effect on plan financial decisions should be investigated further.

The contrast between Leigh's findings on expected benefits and our results on benefits actually received suggests that either union workers are more likely to underestimate their benefits or nonunion workers are more likely to overestimate theirs. The best way to determine which of these two explanations is correct would be to compare questionnaire responses of employees to actual pension plan records. Lacking this, some insight can still be obtained by comparing expected benefits in 1971 in Leigh's sample of men between the ages of 50 and 64 with initial benefits for our sample of persons retiring between 1973 and 1977. Controlling for other factors, Leigh found the expected monthly benefit for union beneficiaries to be $306; nonunion, $358. The initial benefit for a union retiree, based on the results in column 2 of Table 1 is $2568; for a nonunion retiree with the same characteristics, $2411. Converting Leigh's figures to annual amounts, we find that both the mean union and nonunion expectations of benefits are well above the respective means of benefits actually received, but the estimates of union workers are closer to the mark. The degree of overestimation amounts to $1104 annually for the union workers, compared to the $1885 overestimate of annual benefits for nonunion workers. This is consistent with the prediction of the "voice" model that unions have important effects on the flow of information within establishments.

The other important result in this paper is our finding that among employees participating in a pension plan, the union-nonunion compensation differential narrows more rapidly over the life cycle than the union-nonunion earnings differential. This results from the more compressed pension benefit structure under unionism. Benefits increase with earnings and tenure at a slower rate in collectively bargained plans. This effect is accentuated by the slower growth of earnings in the union sector.

This finding, combined with Freeman's examination of earnings profiles, shows that pension coverage is an important conditioning factor for understanding how the union-nonunion compensation differential changes over the life cycle. It would be interesting to see whether the observed patterns of union-nonunion earnings differentials are a function of pension coverage.

Implications of our results on the effect of retirement age on initial benefits and the effect of union status on retirement age are less clear because of unresolved biases. The tremendous difference in the retirement age coefficients in the initial benefit equations for union and nonunion workers does suggest that future attempts to model the pension contract or to explain retirement decisions pay more attention to how incentives for separation and retention of older employees vary by union status.

Footnotes

[1]A complicating factor in this regard is the difference in benefit formulas between union and nonunion plans. Most nonunion plans calculate benefits as a percentage of the product of final earnings and year of service. These formulas partially index benefits to earnings growth, reducing the likelihood of frequent adjustments in the formula. This is not the case for union plans, where benefits equal so many dollars per year of service. Thus, to predict benefits accurately, nonunion employees need to predict future earnings, whereas union employees have to predict future revisions of the benefit formula. This assumes, of course, that both sets of employees know what type of formula is actually being used.

[2]The results reported below also show that the increment in pension wealth resulting from working an additional year is smaller in union plans, giving initial impetus to earlier retirement of union workers.

[3]One reason to believe that this bias is unimportant is that there are no pronounced union-nonunion differences in the distribution of year of retirement in Appendix Table 1.

[4]The economic rationale for such contracts is explored in Allen, Clark, and Sumner (1984b).

[5]The earnings average variable is derived from the individual's Social Security earnings history. The Fox algorithm is used to estimate annual earnings for those with earnings not subject to the Social Security payroll tax.

[6]Two additional possible sources of bias are the endogeneity of pension coverage and union status. Our data set does not include retirees not receiving pension benefits or establishments that do not provide pensions, which prevents us from attempting to correct for the former type of bias. As for the latter type of bias, there is no concensus within the profession about how to deal with it.

[7]Percent union among the beneficiaries of our sample is much larger than percent union among workers. Three factors account for this. First, pension coverage is greater among union workers. Kotlikoff and Smith report from the May 1979 Current Population Survey that 76 percent of union workers are covered versus only 35 percent of nonunion workers. Although percent union among all workers is 24 percent, percent union among those covered by pension plans is 40 percent. A separate breakdown by Kotlikoff and Smith of plan participant data from EBS-1 files shows 47 percent are in union plans. The second factor is that a higher proportion of nonunion participants are in defined contribution plans. The EBS-1 files show 54 percent of the participants in defined benefit plans are in union plans. The third factor is that the average union plan has been in existence longer than the average nonunion plan. We cannot account for the magnitude of the effect of this latter factor.

[8]This adjustment is made by using the result from Allen, Clark, and Sumner (1984b) that union beneficiaries are 23 percent more likely to receive post-retirement increases. With 75 percent of all beneficiaries receiving increases and 73 percent of the beneficiaries being in collectively bargained plans, this means that 81 percent of the union beneficiaries and 58 percent of the nonunion beneficiaries receive benefit increases. These latter two figures are used to obtain estimated pension wealth for union and nonunion beneficiaries, the ratio of which is reported in the last row of Table 5.

[9]Full funding results in lower taxation because assets in the pension fund accumulate at a pre-tax rate, whereas assets held outside the fund accumulate at an after-tax rate.

References

Steven Allen, Robert Clark, and Dan Sumner, "A Comparison of Pension Benefit Increases and Inflation, 1973-79," Monthly Labor Review, May 1984a, pp. 42-46.

Steven Allen, Robert Clark, and Dan Sumner, "Post-Retirement Adjustments of Pension Benefits," NBER Working Paper No. 1364, June 1984b.

Fischer Black, "The Tax Consequences of Long-Run Pension Policy," Financial Analysts Journal, July/August 1980, pp. 3-10.

Jeremy Bulow, "What Are Corporate Pension Liabilities?" Quarterly Journal of Economics, August 1982, pp. 435-452.

Robert Clark, Steven Allen, and Dan Sumner, "Inflation and Pension Benefits," final report under U.S. Department of Labor Contract No. J-9-P-1-0074, August 1983.

Martin Feldstein and Randall Morck, "Pension Funding Decisions, Interest Rate Assumptions, and Share Prices," in Zvi Bodie and John Shoven (eds.), Financial Aspects of the United States Pension System, Chicago: University of Chicago Press, 1983.

Richard Freeman, "The Exit-Voice Tradeoff in the Labor Market: Unionism, Job Tenure, Quits, and Separations," Quarterly Journal of Economics, June 1980, pp. 643-673.

_______________, "The Effect of Unionism on Fringe Benefits," Industrial and Labor Relations Review, July 1981, pp. 489-509.

_______________, "Unions, Pensions, and Union Pension Funds," NBER Working Paper No. 1226, November 1983.

Richard Ippolito, "The Economic Function of Underfunded Pension Plans," manuscript, U.S. Department of Labor, May 1983.

Laurence Kotlikoff and Daniel Smith, Pensions in the American Economy, Chicago: University of Chicago Press, 1983.

Duane Leigh, "The Effect of Unionism on Workers' Valuation of Future Pension Benefits," Industrial and Labor Relations Review, July 1981, pp. 510-521.

H. Gregg Lewis, Union Relative Wage Effects: A Survey, forthcoming.

Olivia Mitchell, "Fringe Benefits and Labor Mobility," Journal of Human Resources, Spring 1982, pp. 286-298.

Alicia Munnell, "Who Should Manage the Assets of Collectively Bargained Pension Plans?" New England Economic Review, July/August 1983, pp. 18-30.

William Sharpe, "Corporate Pension Funding Policy," Journal of Financial Economics, June 1976, pp. 183-193.

Irwin Tepper, "Taxation and Corporate Pension Policy," Journal of Finance, March 1981, pp. 1-13.

Appendix Table 1. Post-retirement increase equations, 1950-72 retirees receiving benefit increases, by union status

Sample	Union			Nonunion		
	Mean	Coefficient	Standard error	Mean	Coefficient	Standard error
1979 recipients/10^5	.261	.193	.014	.048	.785	.054
Years of service/10^2	.258	.404	.037	.241	.195	.039
Age at retirement/10^2	.626	-.187	.072	.614	.046	.075
White	.921	-.081	.011	.959	-.072	.016
Male	.789	-.340	.008	.567	-.066	.006
Year of retirement						
1950	.001	.392	.101	.001	.270	.078
1951	.001	.421	.080	.002	.244	.068
1952	.002	.660	.060	.002	.230	.068
1953	.003	.569	.057	.036	.035	.019
1954	.004	.495	.047	.003	.226	.053
1955	.006	.458	.040	.004	.215	.048
1956	.008	.539	.034	.004	.185	.048
1957	.011	.396	.030	.051	.057	.016
1958	.015	.386	.026	.010	.163	.030
1959	.021	.384	.022	.011	.186	.029
1960	.023	.356	.021	.016	.168	.024
1961	.026	.439	.020	.061	.085	.015
1962	.034	.389	.018	.025	.106	.020
1963	.037	.363	.018	.034	.103	.018
1964	.033	.276	.018	.034	.115	.018
1965	.065	.287	.014	.108	.052	.013
1966	.069	.222	.014	.054	.149	.015
1967	.070	.186	.014	.058	.041	.015
1968	.081	.106	.013	.063	.051	.014
1969	.099	.043	.012	.128	.053	.012
1970	.102	.031	.012	.087	.041	.013
1971	.133	.039	.011	.096	.025	.013
R^2		.053			.104	
N		107352			16117	
Mean of dependent variable		.455			.240	

Note: The dependent variable is the ratio of the increase in benefits between 1973 and 1979 to 1973 benefits. The sample is restricted to those receiving benefit increases. Each equation also includes an intercept and dummy variables indicating industry (6) and whether sex is unreported.