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POST-RETIREMENT ADJUSTMENTS OF PENSION BENEFITS

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ABSTRACT

This paper examines why pension plans increased their liabilities by giving benefit increases to persons no longer working even though almost all of them were not required to do so by any legally enforceable contract. In our model workers and firms have implicit contracts under which post-retirement increases in benefits are purchased by workers through lower wages or initial benefits. Such arrangements permit both plans and workers to share the risk of uncertain rates of return. They also allow beneficiaries to invest at a higher net rate of return than they could obtain elsewhere because of tax advantages and, in large plans, economies of scale. We also discuss how post-retirement adjustments can be used to influence turnover.

Some empirical implications of the model are tested over a sample of beneficiaries of defined benefit plans. The major empirical findings are:

(1) There is strong evidence of compensating differentials in final salary and initial pension benefits for beneficiaries receiving post-retirement adjustments.

(2) Regardless of how the size of pension plans is measured (beneficiaries, participants, amount of benefits paid), large pension plans provide larger post-retirement benefit increases.

(3) Beneficiaries of collectively bargained plans are more likely to receive benefit increases and, among those receiving benefit increases, receive larger increases.

(4) Benefit increases are larger in percentage terms for those who have been retired the longest and for those with the most years of service.

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Introduction

What happens to an individual's private pension benefits after he retires? Until recently, economists and policy analysts have assumed they remained constant in nominal terms, even in periods of severe inflation. Feldstein (1983) and Summers (1983) even have developed models explaining why individuals prefer not to have indexed pensions. These models are consistent with the observation that very few plans provide for automatic adjustments and the assumption that ad hoc increases are uncommon and, when granted, quite small. Three recent surveys of large plans (Bankers Trust, 1980; Hay Associates 1981; and Hewitt Associates, 1981) show approximately two-thirds of large plans giving one or more ad hoc increases during the last half of the 1970s (see King, 1982). However, these surveys did not consider small or medium-sized plans, leaving open the question of how widespread the increases really were; also they did not compare the size of the increase to initial benefits of specific individuals, leaving open the question of their magnitude.

The assumption that post-retirement adjustments are relatively rare has been important in much of the recent theoretical literature on the nature of the pension contract. Barnow and Ehrenberg (1979) and Bulow (1982b) develop models of a firm's pension liability (or cost) under conditions that assume no post-retirement increases in benefits. If plans do award such benefit increases, these models underestimate pension liabilities.

The perception that pension benefits are fixed in nominal terms is also one of the primary reasons that older persons are thought to be adversely affected by inflation (Okun, 1970; White House Conference on Aging, 1982). This had led to a policy debate and research examining the desirability of requiring or encouraging automatic cost of living adjustments (papers by Clark and Spengler, Munnell, and Greenough in Clark, 1980; President's Commission on Pension Policy, 1981). Our understanding of the economic well-being of the elderly during inflationary periods and the value of such regulations will then be enhanced by studying post-retirement benefit changes.

Table 1 reports evidence from Clark, Allen, and Sumner (1983) on the magnitude of post-retirement adjustments between 1973 and 1979 for a nationally representative sample of persons in defined benefit plans who were already retired in 1973. The mean benefit rose from $2128 in 1973 to $2638 in 1979. This increase of $510 amounted to 24 percent of the 1973 benefit. These increases were very widespread, as 75 percent of all beneficiaries received at least one increase and 25 percent received an increase in every year. Among only those receiving increases, the mean 1979 benefit was 32 percent larger than the mean 1973 benefit.

Since inflation was particularly high during this period, it is interesting to compare the rate of increase of nominal benefits to the rate of increase of prices. This is not to say that these benefit increases are attributable to inflation. As we will explain in more detail below, these increases could just as easily be attributed to a risk-sharing arrangement, in which the uncertain parameter is the pension fund's nominal rate of return. This more general framework allows us to explain not only the benefit increases given in the 1970s, but also those given by some plans in periods in which the inflation rate was very low. With this proviso in mind, the Consumer Price Index (CPI) increased by 63.3 percent between 1973 and 1979.

Table 1. Mean pension benefit for persons retired in 1973[a]

Year	Mean benefit	Percent of 1973 benefit	Percent annual increase[b]	Percent change in CPI[c]	Benefit change as a percent of CPI change
1973	$2,128	100.0			
1974	2,205	103.6	3.6	11.0	32.7
1975	2,296	107.9	4.1	9.1	45.1
1976	2,384	112.0	3.8	5.8	65.5
1977	2,452	115.2	2.9	6.5	44.6
1978	2,563	120.4	4.5	7.6	59.2
1979	2,638	124.0	2.9	11.3	25.7
Change 1973-79	510	24.0	24.0	63.3	37.9

[a]Sample includes 139,316 persons who retired in 1972 or before. The observations are weighted by plan weights representing the incidence of similar plans in the pension universe.

[b]The percent increase represents the percentage change in nominal benefits from the preceding year.

[c]The percent change in the CPI is the percentage change in the average annual CPI from the preceding year.

Thus, the mean rate of increase in pension benefits for all beneficiaries equalled 38 percent of the inflation rate. Among only those receiving benefit increases, the mean rate of increase in benefits equalled 51 percent of the CPI increase.

Another interesting way to look at the magnitude of these increases is to ask how much effect they had on pension liabilities. Over this six-year period, benefits rose at an average annual rate of 4.7 percent in plans giving increases. Assuming either a 7 or 10 percent discount rate, the benefit increases raised by 14 percent the present value of benefits paid between 1973 and 1979. Since the mean age of retirement in our sample was 62, the average retiree could expect to receive benefits for 18 years. If his benefits continued to increase at a 4.7 percent rate, the plan's liability for this worker's lifetime benefits would increase by 44 percent at a 7 percent discount rate and by 39 percent at a 10 percent discount rate.

This paper examines why pension plans increase their liabilities to such a large extent when not required to do so by any explicit contract. In our model workers and firms have implicit contracts under which post-retirement increases are purchased by the worker through lower wages, lower initial benefits, or reductions in other forms of compensation. In return the worker receives (1) the prospect of a higher after-tax rate of return from reinvesting some of his pension wealth in the plan during the early years of retirement (as opposed to investing it on his own) and (2) a hedge against inflation and rate-of-return risk, the magnitude of which depends upon the composition of the plan's portfolio. The compensating differentials framework has been used by Ehrenberg (1980), Schiller and

Weiss (1980), and Smith (1981) to examine tradeoffs between wages and both the magnitude of and likelihood of receiving pension benefits. We also explore how post-retirement adjustments can be used to influence employee behavior. Some empirical implications of our model are then tested over a sample of beneficiaries collected by the Department of Labor from a representative set of defined benefit pension plans.

Rationale for Post-Retirement Adjustments

Between 1973 and 1979 most beneficiaries received increases in their nominal pension payments after they had retired. Virtually none of these increases were required by any formal contract. Why, then, are defined benefit plans voluntarily raising their pension liabilities? The explanation offered in this paper is the existence of an implicit contract between the plan and both workers and retirees, in which the plan guarantees a minimum nominal annuity plus the possibility of post-retirement adjustments in the future. The quid pro quo for future benefit increases is lower wages, lower initial benefits, or some other form of compensating differential. An unusual feature of this form of compensating differential is that the worker must pay before he benefits. Therefore, the existence of such contracts will depend in part on the confidence of the worker or retiree in the reliability of the future payment.

Consider the case of a person who is about to retire and wants to select the optimal payment schedule for receiving his pension benefits. Suppose he has two alternatives. The first is the conventional retirement annuity fixed at the same nominal level in each future period. The second is a promise of a stream of pension benefits where a minimum nominal payment per period is guaranteed to the worker but benefit increases may be granted in the future. The latter arrangement is the "implicit contract" model described in Pesando (1984). His model is motivated by the widespread tendency to finance post-retirement adjustments out of "excess" plan earnings (earnings above the level required by the rate of return used in evaluating the plan's liabilities). Here we address the question of when such contracts are preferable to workers.

The implicit contract works in the following fashion. Consider a two-period retirement framework where L_i, i = 1, 2 indicates benefits in each period. There is a minimum guaranteed benefit (L_m), which is set so the plan is fully funded for both periods based on the plan's interest rate assumption r_v. This rate will be less than the risk-free interest rate and as a result, L_m will be lower than the benefit in a fixed annuity plan (or contributions during the working period must be higher). This initial benefit or wage offset is required because benefits are assumed to be adjusted only in an upward direction, not downwardly. We also assume the minimum guaranteed benefit is paid in the first period. Letting A_v be the contribution in the pre-retirement period required to fully fund benefits of L_m in each period, we have $A_v(1+r_v)^2/(2+r_v) = L_m$. Let r be the actual rate of return earned by the plan's assets. If $r > r_v$, then the plan provides a benefit L_2 equal to the total remaining assets in the fund, i.e., $L_2 = (A_v(1+r) - L_m)(1+r)$. If $r \leq r_v$, then $L_2 = L_m$.

This contract allows the worker to receive returns to the fund's assets above the threshold indicated by r_v. It also allows the worker and the firm to share the risks of uncertain rates of return in contrast to the conventional defined benefit plan in which the firm is usually assumed to bear all of the risk of nominal rates of return and defined contribution plans in which the worker supposedly bears all of such risk. It offers the worker some insurance against unexpectedly high rates of inflation as long as the plan's rate of return has a positive covariance with inflation. The amount of inflation insurance depends on the composition of the pension fund's portfolio. Feldstein (1983) has shown that unless the worker has an infinite degree of risk aversion, full indexation is not optimal.[1]

In return for these benefits the worker must accept a lower wage, a smaller initial benefit, or both. The equilibrium condition is the sum of the present value of earnings, guaranteed pension benefits in both periods, and the expected magnitude of the post-retirement adjustment in the second period equals the present value of the marginal product while working.[2]

The choice between the fixed nominal annuity and the "implicit contract" boils down to an exercise in the optimal intertemporal allocation of income under uncertainty. There are a number of important sources of uncertainty to consider. We will focus on the plan's nominal rate of return, the nominal rate of return on assets held by the worker, and the likelihood the plan will meet its obligations. Even though workers and retirees in any plan are heterogeneous in both their preferences toward risk and their risk-bearing ability, they will be covered by the same set of rules defining the contract. A non-union plan will adopt a decision rule such as selecting the contract that maximizes the utility from pension wealth of the marginal worker. In a plan covered by collective bargaining, a different decision rule will be applied, such as maximizing the utility from pension wealth of the median union member (who may be retired). The implications of these differences for union and non-union plan behavior will be discussed later.

Cross-sectional differences in the demand for post-retirement adjustments will be related directly to the expected difference between the after-tax rate of return earned by the plan after the worker retires and the after-tax rate of return the worker expects to receive on his own assets.[3] Although we have no direct information on the expected rates of return earned by each plan in our sample, there are a number of reasons to believe this variable is highly correlated with plan size. First, part of the cost of administering a pension plan is fixed, especially investment and actuarial services. According to Smeeding (1983), this produces economies of scale with respect to custodial (administrative overhead) fees and securities commissions for portfolio adjustments, especially the latter. Smeeding reports that these fees fall from 5.90 percent for IRA or Keogh plans to 4.43 percent for small pension plans to 3.54 percent for large pension plans. Mitchell and Andrews (1981) also show dramatic declines in administrative expenses per participant among multi-employer plans, especially among plans with fewer than 3000 participants. Second, small defined benefit plans must make more conservative actuarial assumptions because of longevity risk. As plan size decreases, the longevity distribution of its beneficiaries is less likely to approximate that of the entire population. This increases the likelihood that a sizable proportion of beneficiaries will live longer than expected and reduces the benefit that can be offered for a given contribution. Third, among plans funded through trusts, the assets of each plan must be segregated from those of all other plans. This makes it very difficult for small plans to obtain adequate diversification. Although this constraint can be avoided by funding the plan through an insurance company, the administrative fees or contingency reserves required by insurers usually fall in percentage terms with the dollar volume of annual contributions (see McGill, 1975).

An indirect test of whether some or all of these factors make rates of return an increasing function of plan size is to examine actuarial interest rate assumptions. Malca (1975) reports these assumptions for 1972 and 1973 from a Standard and Poor's survey. The results show the median asumption for the largest size category was 5.5 percent, compared to 4.8 to 5.0 percent for plans in intermediate size categories and 4.5 percent for plans in the smallest category. In summary, as long as large plans are expected to realize larger rates of return and rates of return for individual workers and retirees are randomly distributed across plans, individuals in large plans should be more willing to purchase post-retirement adjustments. In addition to rate of return differences, this coefficient will also reflect the effect of unobserved factors correlated with the size of pension plans.

Taxes are an additional factor affecting the difference between the expected rate of return for the plan and the return the worker expects to receive on his own assets after retirement. Assets can accumulate at a tax-free rate in the pension plan, whereas retirees must pay taxes on earnings from their investments. For those who intend to save part of their pension benefits in the initial years of retirement, this provides an incentive for the implicit contract outlined here. Unfortunately, our data set lacks the detail necessary to construct marginal tax rates for beneficiaries. Some of these effects will be picked up by the other variables in the empirical model outlined below.

Post-retirement adjustments are usually granted on an ad hoc basis rather than through any formal escalator clause. For instance, in our sample only seven plans had explicit contractual provisions for automatic adjustments at regular intervals. These intervals can be quite long; one plan required adjustment every three years and another required them every sixth year. If the contract for post-retirement adjustments is implicit and not legally enforceable, the plan stands to gain considerably by reneging on the agreement. The probability of cheating, in turn, determines the probability that such contracts are acceptable to workers.

The cost to the plan of cheating is reduced ability to write similar implicit contracts in the future, since workers will simultaneously lower the mean expected return and raise the expected variance from buying these contracts. It also may cause the expected real present value of the "guaranteed" minimum benefits to fall, especially in underfunded plans in which liabilities are not fully covered by insurance. These costs are likely to be especially sizable for large firms because (1) information about their activites is likely to spread more widely through the labor market and (2) assuming labor is more specialized in large firms, they are more likely to use pension plans to reduce turnover and the costs of investing in specific on-the-job training. If the likelihood of cheating is lower in large plans, this gives us an additional reason for expecting larger post-retirement adjustments in such plans.

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,[4] 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, since 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 device for investments or risk-sharing when they retire. This is especially likely to be true for older workers. Both the median voter model and recent empirical findings by Freeman (1983) on pension plan provisions indicate that the preferences of older workers receive much more weight than those of younger workers in forming union objectives. This makes it more likely that unions will act in their interests.

Another factor encouraging unions to act in this fashion is the activity of retirees in union political affairs. Retirees can sometimes vote for union officers and attend union conventions. In the United Mine Workers they even vote on contract ratification. This means distributing a portion of any rents obtained in negotiations to retirees can yield a political payoff to union officers. In contrast, retiree preferences will receive zero weight in a non-union setting, making an intergenerational transfer from workers to retirees unlikely. Thus, we expect larger post-retirement adjustments in collectively bargained plans because of (1) greater costs to the firm for reneging on the implicit contract and (2) a preference-weighting scheme tilted toward retirees and older workers.

In addition to post-retirement adjustments existing as a form of compensation for which workers must pay, firms may use these contracts to regulate employee behavior. Unvested pensions can reduce turnover by increasing the cost of taking a job at another establishment. In inflationary periods, pension benefit formulas based on salaries also penalize job changers, since the formulas are based on nominal rather than real salaries (see Clark and McDermed, 1982, and Bulow, 1982a). Post-retirement adjustments can be distributed in order to either maintain or exacerbate these effects. In Clark, Allen, and Sumner (1983), we found that 17 percent of the plans giving adjustments increased benefits by a straight percentage, whereas 27 percent made the magnitude of the increase a function of years of service. Furthermore, eligibility for increases in some of the plans was limited to workers with a minimum amount of preretirement service ranging between 10 to 20 years.

Theoretically, it is not clear that post-retirement adjustments would be a more or less effective device than the benefit formula itself for influencing employee quit decisions. Conceivably, post-retirement adjustments simultaneously increase the returns from working an additional year with the firm (via an implicit contract in which the magnitude of future benefit increases is linked to tenure) and reduce the variance of expected real pension wealth for workers by insuring against inflation risk. If these effects have a greater impact on turnover than an upward adjustment in the benefit formula costing an equal amount, such an approach may be optimal. To put the same point somewhat differently, by giving post-retirement adjustments to its retirees, the firm hopes to change the expectations of its current workers so that they value their pension contract in real rather than nominal terms.

Another possibility is that since the Employee Retirement Income Security Act (ERISA) has made it more difficult for firms to use vesting requirements to discourage turnover, basing post-retirement adjustments on years of service is now a second-best mechanism for achieving the same result. Regulation of explicit pension contracts may also explain why firms have adopted implicit contracts.[5]

Data Description

The Pension Benefit Master File (PBMF), made available to the authors by the Department of Labor, is the primary data source for this study. These data are from a stratified random sample of pension plans filing series 5500 and 5500C forms in 1975. The PBMF combines information from the Arthur Young and Company Survey of Private Pension Benefits Amounts with Social Security information from the Summary Earnings Record and from a standard summarization of the Master Beneficiary Record, known as the Survey Benefit Summary Record. The Arthur Young survey contained information on 671,000 persons receiving benefits on December 31, 1978 from 446 plans of 371 sponsors. In all but two cases where more than one plan per sponsor was reported, the second and/or third plan was not part of the formal sampling process but was provided along with the requested plan data by the responding plan. These additional plans were deleted from the analysis, since they were assigned a zero plan weight in the sampling process. Weights provided by the Department of Labor enabled us to construct weighted samples of individuals and plans that were representative of the set of pension plans that existed in 1975.

Although the PBMF included defined contribution 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 sampling procedure that made it impossible to determine potential benefit increases for many of these plans. In addition, only 50 percent of the beneficiaries in the five largest plans in the Arthur Young survey were matched with the Social Security data because of a resource constraint imposed by the Social Security Administration. The weights for individuals in these plans were doubled to compensate for the decrease in the sample size. Persons in any plan receiving a lump sum distribution were eliminated from the sample.

The PBMF contained information reported by firms on individuals who were receiving benefits in December 1978. Data were included on age, year of retirement, years of credited service, sex, race, marital status, Social Security reported earnings, and the current pension benefit. All benefit amounts were converted to an annual benefit. In addition, plan characteristics such as union status, number of beneficiaries, and industrial category also were reported. Specific questions were asked about any increases in post-retirement benefits awarded between January 1, 1973 and January 1, 1979. Plan sponsors were asked to indicate the size and method of each increase, date of each increase and types of beneficiaries eligible for each increase.

Using this information, we were able to construct annual pension benefits from 1973 to 1979 for most of the beneficiaries. This task required that individuals in each plan be examined carefully to determine if they were eligible to receive an increase. Then, using the increase formula, the magnitude of the increase for each individual was calculated. Working backwards over the six-year period, we determined the annual benefit for most of the beneficiaries in the sample. Some plans failed to report information such as years of service that was necessary to reconstruct benefit increases. In these plans we attempted to use related information in the PBMF to construct appropriate proxies for the missing data. For several plans, this was impossible and they were deleted from the sample.

The reported benefit on the PBMF is assumed to be the benefit a person would receive throughout 1979. Benefit increases reported in one year are assumed to take effect in January of the following year. For example, the 1979 benefit reported on the PBMF reflects all 1978 increases. Thus, to determine the 1978 benefit we subtract the implied increase from the 1979 benefit. This process continues until annual benefits from 1973 to 1979 are calculated. This paper reports the results of an analysis of post-retirement benefit increases for persons who were retired throughout this period. Thus, we selected individuals who were already retired at the beginning of 1973 so that they were receiving benefits during the entire period.

To estimate post-retirement benefit adjustment equations, some additional restrictions had to be imposed on the sample. In three plans the average percent increase in nominal benefits was over 190 percent. In no other plan, however, was the average increase more than 75 percent. Although the 1979 benefits in these plans were comparable to those of other plans in the sample, the derived 1973 benefits were unusually small. We attribute this to incomplete or inaccurate information provided us about how benefit increases between 1973 and 1979 were calculated. These three plans are excluded from the sample. This problem also arose for some individuals in other plans, again presumably because of errors in reporting benefit changes in complete detail. Accordingly all individuals with 1973 benefits of less than $10 are excluded from the sample. Finally, there were severe reporting errors in or missing values for three of the independent variables (years of service, year of retirement, and age at retirement; for example, average age at retirement in one plan was 85) in twelve of the plans. These plans were dropped from the sample. When only one or two of these variables were missing or implausible, sample means were substituted for the reported value. All estimates reported below are derived using the PBMF sampling weights. The weighting is necessary because large plans were intentionally overrepresented in the survey.

Empirical Specification

The rationales for post-retirement adjustments offered above have empirical implications that are testable over the PBMF. The compensating differential implies a tradeoff between post-retirement adjustments and either wages or initial benefits. Large plans are more likely to provide post-retirement adjustments because of higher rates of return and a lower probability of reneging on the implicit contract. The latter factor, along with the political dominance of older workers and retirees, makes collectively bargained plans more likely to provide post-retirement adjustments. If these adjustments are used to discourage turnover, they should be correlated with years of service.

We will focus most of our attention on a specification in which the dependent variable is the change in benefits between 1973 and 1979 divided by 1973 benefits. This variable is not distributed normally because it is truncated from below at zero. To shed more light on the determinants of where any individual is likely to be in that distribution, we estimate OLS regressions for a binary variable indicating whether a person ever received an increase in his benefit and for the ratio of the change in benefits to 1973 benefits for those who received increases. Although OLS is not the most appropriate estimation technique for a binary dependent variable, alternatives such as probit or maximum likelihood logit are quite expensive for a sample of over 130,000 observations. Since switching from OLS to one of the other techniques generally does not radically alter estimates in large samples, we did not feel such an expense was justified.

Comparison of 1973 and 1979 benefits can produce misleading conclusions about the adjustment of benefits if there are important differences in the timing of such adjustments. To take this into account, we consider two additional dependent variables: (1) the number of increases given over this period and (2) the ratio of the present value of real benefits between 1973 and 1979 to the present value of benefits that would have been received under complete indexation.

Our choice of independent variables is restricted largely by the available data. To estimate the tradeoff between wages and post-retirement adjustments, we use the five-year salary average before retirement as the wage variable. Annual salaries were estimated from the Social Security earnings histories using Fox's (1979) method. Final five-year salary averages were not available for persons retiring before 1956. Initial benefits for this sample cannot be determined, since our knowledge of benefit increases begins with 1973. As a proxy we use 1973 benefits. This variable equals initial benefits plus any post-retirement adjustments granted before 1973. If these two variables are independent, the 1973 benefit produces a downwardly biased estimate of the tradeoff between initial benefits and post-retirement adjustments (because of positive correlation between post-retirement adjustments before and after 1973). However, since our model predicts a negative correlation between initial benefits and post-retirement adjustments, the direction of bias cannot be predicted.

Most, but not all, plans reported collective bargaining status on either the PBMF or a file of EBS-1 reports obtained from the National Bureau of Economic Research. The collective bargaining status for some plans remained unspecified. Rather than throwing out these observations, we use two union status variables. 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-non-union difference. Union status is correlated with many other variables, requiring some additional controls. A set of industry dummies at roughly the 1-digit level of aggregation is included for this reason.

Our measure of plan size is the number of beneficiaries in 1979. Since plan size can be measured in a number of additional dimensions, we also examined the number of participants and the dollar value of all benefits paid to 1979 beneficiaries. These specifications produced very similar results and are not reported. The years of service variable used is reported by the pension plan. This does not necessarily equal total years employed by a given company, depending upon rules for participation in the plan. For instance, years of credited service under the pension plan may be somewhat smaller than total years employed in the firm.

Dummy variables corresponding to the year in which the person retired are included in the model to test whether persons who have been retired the longest are treated differently from recent retirees. Although this is an important empirical question for measuring the economic well-being of the elderly, our model does not address this issue. If post-retirement adjustments result from implicit contracts to insure against inflation risk, the optimal insurance policy could conceivably be one that provides the largest payoffs to those who live the longest. Such insurance is cheaper than a policy providing equal protection at all ages and, under fairly reasonable assumptions, provides a greater reduction in wealth uncertainty during retirement. For most individuals, self insurance will be more efficient in the years immediately after retirement. They pay a "deductible" for protection in later years by receiving relatively smaller (or zero) post-retirement adjustments during first retirement years. Finally, we include age at retirement, sex and race as independent variables. These variables capture longevity risk differences or other differences not accounted for in our specification.

Empirical Results

Although the PBMF includes individuals retiring since 1950, five-year Social Security earnings data are available only for persons retiring since 1956. This forces us to examine two different samples to be able to both estimate compensating differentials and measure the distribution of post-retirement adjustments across the broadest possible number of cohorts. Columns 1-3, Table 2, show the means and coefficients of equations estimated

Table 2. Ratio of 1973-79 post-retirement adjustments to 1973 benefit equations[a]

	1956-1972 retirees[b]			1950-1972 retirees[c]	
Variable	Weighted mean	Coefficient (standard error)		Weighted mean	Coefficient (standard errors)
	1	2	3	4	5
Intercept	-	.180 (.035)	2.038 (.049)	-	.109 (.036)
1979 recipients (times 10^{-5})	.168	.320 (.011)	.481 (.011)	.170	.439 (.011)
Plan collectively bargained (yes=1)	.726	.155 (.006)	.160 (.006)	.724	.169 (.006)
Years of service (times 0.01)	.247	.328 (.025)	1.564 (.028)	.247	.360 (.026)
Age at retirement (times 0.01)	.625	-.229 (.049)	.013 (.047)	.625	-.023 (.050)
Log (final 5-year salary average)	8.501	-	-.053 (.004)		
Log (1973 benefit)	7.370	-	-.285 (.003)		
Male	.777	-.128 (.005)	-.0002 (.006)	.756	-.211 (.005)
White	.930	-.049 (.008)	-.006 (.008)	.930	-.054 (.008)
R^2	-	.051	.117		.069
N		129,057	129,057		137,038

[a]The dependent variable is the ratio of the change in benefits between 1973 and 1979 to the 1973 benefit. Binary variables indicating whether collective bargaining status is unreported, whether sex is unreported, industry of employer, and year of retirement are also included.

[b]Sample includes persons who retired from 1956 to 1972 for whom a final five-year salary average was available. The weighted mean of the dependent variable is 0.285 over the 129,057 unweighted observations in this sample.

[c]Sample includes persons who retired from 1950 to 1972. The weighted mean of the dependent variable is 0.308 over the 137,038 unweighted observations in this sample.

over those retiring since 1956. A specification without the salary average and initial benefit variables is reported so valid comparisons can be made to the means and coefficients in columns 4 and 5 obtained when the sample is expanded to include persons retiring since 1950. The dependent variable is the ratio of the change in benefits between 1973 and 1979 to benefits in 1973.

The coefficients of both the salary average and the 1973 benefit variables in column 3 are negative and large in absolute value relative to their standard errors, as predicted by the compensating differentials model. Controlling for other factors, a 10 percent increase in salary average reduces the amount of post-retirement adjustment over our six-year period by about 0.5 percentage points. A 10 percent increase in 1973 pension level reduces the post-retirement adjustment over the period by 2.9 percentage points. Both of these results imply rather small losses in present value of pension flows from post-retirement adjustments for large gains in final wage or initial pensions.

The magnitude of the implied compensating wage differential indicates that a pure tradeoff between forms of compensation is not being estimated. One factor accounting for this is a tendency for all forms of compensation to be positively related across workers or firms. Only by holding constant the level of total compensation including unobservable job and worker characteristics may pure compensating differentials be calculated.[6] All the regressions in Table 2 include several controls that are positively related to level of compensation, and coefficients for each of these variables (plan size, unionization, tenure, age at retirement and male) are larger in column 3 than they are in column 2. This indicates that in column 2 they picked up part of the negative effect of salary and pension level on post-retirement adjustments. Even given the controls, however, a considerable amount of uncontrolled variation in post-retirement adjustments remains (witness the low R^2 values), so the magnitude of the compensating differential estimates is not surprising. A second reason for the small size of these coefficients is that measurement error in both the initial benefit and salary average variables downwardly biases their regression coefficients and, thus, upwardly biases the estimate of the compensating wage differential.

A final explanation for large implied compensating differentials is that the 1973 benefit variable is larger than initial benefits for all retirees who received any post-retirement adjustment prior to 1973. Using a simulation analysis, we estimated that by 1973, persons retiring between 1956 and 1960 already had received post-retirement increases of between 70 and 120 percent of their initial benefits (Clark, Allen, and Sumner, 1983). This implies that the 1973 benefit is approximately twice the size of the initial benefit for these retirees. Thus, the tradeoff implied by the coefficients reported in Table 2 for post-retirement increases as a percentage of the 1973 benefit is larger in terms of the retiree's initial pension benefit. For the pre-1960 retirees, the implied increase in benefits between 1973 and 1979 is about 6.0 percentage points (2.9 percent of the 1973 benefit is approximately 6.0 percent of initial benefit for these oldest retirees) for a 10 percent decline in initial benefits. If this 6 percentage point increase is awarded every 6 years and if a retiree lives 18 years after retirement, the benefit at death will be approximately 19 percent higher than the initial benefit.[7] Although still not an equal trade in present value terms, such a ratio may reflect more accurately tradeoffs between initial benefits and post-retirement adjustments.

Estimates of the other coefficients are generally consistent with the reasoning presented earlier. The plan size coefficient indicates a positive and significant impact of larger plan size on the magnitude of post-retirement increases. A ten thousand-person increase in the number of recipients in 1979 is associated with a 4.8 percentage point larger benefit increase. The largest plan in the sample had 67,130 recipients in 1979; the regression coefficient implies that this plan gave a 24 percentage point larger increase than the average plan. The smallest plan in the sample had one recipient. The estimated coefficient implies that it gave an 8.2 percentage point smaller increase than the average plan. A significant plan size effect was found throughout our analysis as we varied the independent variables in the equation and used alternative definitions of plan size. The magnitude of the plan size coefficient declines somewhat when salary average and 1973 benefit are deleted (column 2) or when the sample is expanded to include persons retiring in the early 1950s (column 5) but it remains significantly greater than zero.

Collectively bargained plans granted larger benefit increases than non-collectively bargained plans. The increases in union plans were estimated to be 16.0 percentage points larger than those in non-union plans. At the sample means the average union plan gave a 32.9 percent increase; the average non-union plan, a 16.9 percent increase. Thus, increases in union plans were almost twice those in the non-union plans. Table 2 indicates that the union coefficient changes only slightly when the specification and sample are changed.[8]

Years of service is strongly correlated with benefit increases. An additional year of service is associated with a 1.6 percentage point larger increase in benefits during the sample period. The age of retirement coefficient is estimated with little precision. The sign indicates the magnitudes of adjustments increase with later retirement. Use of firm-reported information about whether the retirement was early, normal, or delayed did not improve the precision of this estimate. Both the years of service and age at retirement coefficients are more sensitive to changes in the specification and sample. Age at retirement is the only variable for which the sign of the coefficient changed in response to alternative specifications.

As for the other coefficients, post-retirement benefit increases are insignificantly different by sex and race in the compensating differential model shown in column 3. The deletion of the salary average and 1973 benefit from the model produces estimates indicating larger post-retirement adjustments for females and nonwhites. The industry variables (not shown in Table 2) indicate that increases were larger in the mining, manufacturing and transportation sectors of the economy.

The expanded sample of 1950-72 retirees is used to examine the magnitude of benefit increases in percentage terms across cohorts. The year of retirement means and coefficients are reported in the first two columns of Table 3. These coefficients come from the benefit change equation reported in column 5 of Table 2. They indicate how the magnitude of the adjustment differs by year of retirement relative to a person retiring in 1972. The ratio of the change in benefits between 1973 and 1979 to 1973 benefits by year of retirement in the third column is calculated by adding the year of retire-

Table 3. Year-of-retirement coefficients and estimated change in benefits by year of retirement, 1950-72 retirees

Year of retirement	Weighted mean	Coefficient (standard error)	Change in nominal benefits ÷ 1973 benefits at sample means[a]	Change in nominal benefits at sample means ÷ change in CPI[b]
	1	2	3	4
1950	.001	.296 (.067)	.524	.828
1951	.001	.362 (.060)	.590	.932
1952	.002	.562 (.049)	.790	1.248
1953	.008	.231 (.026)	.459	.725
1954	.004	.356 (.035)	.584	.923
1955	.005	.348 (.031)	.576	.910
1956	.007	.397 (.026)	.625	.987
1957	.016	.229 (.018)	.457	.722
1958	.013	.284 (.019)	.512	.809
1959	.017	.303 (.017)	.531	.839
1960	.020	.271 (.016)	.499	.788
1961	.032	.255 (.013)	.483	.763
1962	.031	.265 (.013)	.493	.779

Table 3 (continued)

Year of retirement	Weighted mean	Coefficient (standard error)	Change in nominal benefits ÷ 1973 benefits at sample means[a]	Change in nominal benefits at sample means ÷ change in CPI[b]
	1	2	3	4
1963	.036	.251 (.013)	.479	.757
1964	.036	.176 (.013)	.404	.638
1965	.071	.177 (.010)	.405	.640
1966	.063	.160 (.010)	.388	.613
1967	.067	.125 (.010)	.353	.558
1968	.077	.069 (.010)	.297	.469
1969	.107	.045 (.009)	.273	.431
1970	.104	.029 (.009)	.257	.406
1971	.131	.026 (.008)	.254	.401
1972	.151	--	.228	.360

[a]The change in nominal benefits for persons retiring in a given year are calculated for a white male union manufacturing worker with the mean values of years of service, number of 1979 recipients in plan, and age at retirement.

[b]The values in this column are calculated by dividing the ratios in the preceding column by ratio of the change in the CPI between 1973 and 1979 to the CPI in 1973 (.633).

ment coefficient to the predicted benefit for a white male union manufacturing worker retiring in 1972 with the sample mean values of 1979 recipients, age of retirement, and years of service. These figures are divided by the ratio of 1973-79 CPI change to the 1973 CPI (.633) to obtain the ratios of the percentage change in nominal benefits to the percentage change in prices in the last column.

The regression coefficients show that percent increases in pension benefits are much larger for those who have been retired the longest. The average white male union manufacturing worker retiring in 1972 had a 22.8 percent larger pension in 1979 than in 1973. Workers who retired before 1964 received at least twice as large a percentage increase. This pattern of larger benefit increases is caused by firms using increase formulas that provide explicitly for larger increases to those retired for longer periods. In addition, flat dollar increases will also give larger percentage increases to long-term retirees who had lower initial benefits.

Although inflation need not have been a direct causal factor, an important policy issue is how these increases compare to the change in the cost of living over this period. We think they were surprisingly large. The conventional view has been that pension benefits did not adjust at all. Instead we find workers who retired before 1964 saw their benefits increase by more than three-quarters as much as prices. Workers who retired between 1964 and 1967 saw their benefits increase by more than 50 percent of the price rise. The workers who were most likely to have anticipated relatively little inflation during retirement were also least likely to have witnessed a severe erosion in real benefits.

To determine the sensitivity of the results to the truncation of the dependent variable at zero, the model was re-estimated using two different equations: (1) a linear probability model of whether an increase was given between 1973 and 1979 and (2) the ratio of the benefit increase to 1973 benefits, estimated over only those receiving at least one increase. These were estimated over the sample of persons retiring between 1950 and 1972 and are reported in columns 1 and 2 of Table 4. These findings are in general agreement with those reported in Table 2. Increases in plan size raise the probability of a person's having received any increase and the magnitude of increases conditional on having received at least one increase. Each 10,000 increase in the number of 1979 recipients raises the probability of receiving at least one benefit increase by 5.2 percentage points and the magnitude of the total increase between 1973 and 1979 by 2.4 percentage points. Being in a union plan raises the likelihood of receiving an increase by 22.6 percentage points and the magnitude of increases by 12.3 percentage points among those receiving increases. Years of service retains its strong positive correlation with both measures of benefit increases, whereas increases in the age of retirement lower the probability of receiving an increase but have no effect on the magnitudes of increases among those receiving them.

The timing of increases is considered by examining the number of increases given over the period (in column 3 of Table 4) and the ratio of discounted nominal benefits to fully indexed benefits (in column 4). Once again, the overall picture from Table 2 is more or less unchanged. Larger plans gave more increases and the increases were larger in present value. Collectively bargained plans, on average, gave one more increase than did non-union

Table 4. Benefit change equations, 1973-79[a]

Dependent variable	Increase given between 1973 and 1979 (yes=1)	Change in benefits 1973-79 ÷ 1973 benefit; sample restricted to those receiving increases	Number of increases 1973-79	Ratio of discounted real benefits to discounted real benefits under full indexation
Weighted mean of dependent variable	.749	.411	2.645	.902
Independent variable:		Coefficients (S. E.)		
Intercept	.855 (.016)	.172 (.043)	.470 (.061)	.825 (.012)
1979 recipients (times 10^{-5})	.520 (.005)	.241 (.012)	6.090 (.019)	.179 (.004)
Plan collectively bargained (yes=1)	.226 (.003)	.123 (.008)	1.086 (.010)	.048 (.002)
Years of service (times 0.01)	.252 (.012)	.402 (.032)	.879 (.044)	.139 (.009)
Age at retirement (times 0.01)	-.134 (.023)	-.016 (.061)	.101 (.084)	.007 (.016)
Male	-.054 (.002)	-.261 (.006)	-.069 (.009)	-.065 (.002)
White	.001 (.004)	-.077 (.010)	-.094 (.014)	-.024 (.003)
R^2	.294	.056	.650	.084
N	137,038	123,469	137,038	137,038

[a]The regression equations include year of retirement and industrial binary variables. In addition, two other binary variables indicating whether collective bargaining status or sex is unreported are included. The sample includes persons who retired between 1950 and 1972.

plans. Individuals with long job tenure were more likely to receive increases. On the whole, the basic findings seem quite robust as to the specification of the dependent variable.

Conclusion

The empirical results provide strong evidence of compensating differentials in final salary and initial benefits in firms providing post-retirement adjustments. The strong weight given to years of service is consistent with the use of benefit increases to regulate employee mobility. The results also show that post-retirement adjustments are larger in large plans and in plans covered by collective bargaining agreements. This is consistent with the notion that implicit contracts are more likely to be written when the likelihood of reneging on the contract is low. A final major finding of this paper is that benefit increases tend to be larger for those who have been retired the longest.

Regardless of how these results are interpreted in theoretical terms, they indicate strongly that the private pension system was much more responsive during the 1970s inflation than was previously believed. Further, our evidence on tradeoffs between benefit increases and both wages and initial benefits suggests regulations requiring private pension indexation will lower the welfare of those who prefer higher wages or initial benefits to future benefit increases.

These results also suggest existing models of the costs of pension benefits to employers (or equivalently, their value to employees) be re-examined. With most beneficiaries receiving sizable post-retirement adjustments, the present value of the expected stream of payments implied by the benefit formula at the time of retirement will in most cases be a downwardly biased measure of worker's pension wealth. This makes it impossible to use an "explicit contract" framework to derive either the increment to this wealth in each year of employment or the firm's total pension liability. Since the empirical results imply workers earn larger post-retirement adjustments with additional years of service, the tilt of pension accruals toward the most senior workers has been underestimated in previous studies. The "accrued benefit" approach for evaluating pension liabilities produces misleading results for the same reason. Unless the implicit contract is "unveiled," this approach will ignore an important part of the plan's liabilities.

We do not claim to have identified fully all parameters of such contracts. In future work it would be useful to explore the effects of such variables as plan financial performance and training costs on benefit increases to test directly the rationales we have offered rather than using proxies such as plan size. It is also possible that the experience in the 1970s was atypical, a conjecture that can only be tested by examining data sets from other periods.

Footnotes

[1]TIAA-CREF has developed an explicit contract similar to the one described here. TIAA-CREF Graded Payment Method allows a person to have scheduled increases in benefits in exchange for a lower initial benefit.

[2]This condition holds for a given value of the riskiness of r. The condition would have to be restated in terms of the expected utility of each party to allow the riskiness of r to vary across plans.

[3]The contract for post-retirement adjustments may be desirable as a risk-sharing device even if it offers a lower expected rate of return than the worker's own assets. This would be the case, for instance, if the variance of the plan's rate of return is lower than the worker's or if their covariance is negative.

[4]Allied Chemical & Alkali Workers v. Pittsburgh Plate Glass Co., 404 U. S. 157 (1971).

[5]Unfortunately, our time series does not allow us to test whether the use of post-retirement adjustments increased after the passage of ERISA. Simulation experiments in Clark, Allen, and Sumner (1983) indicate that benefit increases were awarded during the 1950s and 1960s.

[6]See Duncan and Holmlund (1983) for discussion of potential biases in estimates of compensating differentials.

[7]These values are derived by assuming the 1973 benefit is twice the size of the initial benefit for persons retiring during the 1950s. A 2.9 percent increase in 1973 benefits would represent approximately 6 percent of the initial benefit. The value of post-retirement adjustments would be greater if the retiree has chosen a joint survivors option where the higher benefits

would continue as long as the retiree or spouse survived. There were insufficient data on the PBMF to determine if persons selecting joint survivorship received larger or more frequent increases.

[8]For a more detailed analysis of the effect of unions on post-retirement increases as well as on the initial pension benefit, see Allen and Clark (1984).

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