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11 Pension Inequality

Edward P. Lazear and Sherwin Rosen

Much attention has been given to earnings inequality in recent years. Although most agree that the variable of interest is lifetime wealth rather than current earnings,[1] there has been relatively little study of differences in nonwage and salary components of earnings. Pension inequality is interesting for a number of reasons: First, pensions are a large fraction of total nonwage compensation. Second, there have been recent changes in laws that regulate sex-based differences in pension benefits. Third, private pensions have grown in importance over time and may become even more important in the future.

What follows is an attempt to determine whether pensions exacerbate compensation inequality across groups. There are two aspects to this issue. The first is that the probability of receiving a pension may not be random across groups. For example, in Lazear (1979), Retirement History Survey data revealed that 49% of the workers in the sample had pension plan coverage, but blacks were 6.6% less likely to be covered than whites. Similarly, female coverage was 8.6% less than males. These patterns are investigated in more detail in the CPS data below. The second aspect is how the size of pensions varies with sex and race of people who are eligible to receive them. This is more difficult to determine and is the main focus of this study.

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There are two empirical tasks before us. The first is to determine the characteristics of the average retiree in each sex and race group. Especially important is the average tenure, age, and salary of the typical retiree because pension amounts in most plans depend on these variables. The second task is to estimate the pension that each group's typical retiree receives. This depends on the plan in which he is enrolled, so it is necessary to use some representative sample of plans.

The May 1979 Consumer Population Survey is used for the first task. This was chosen over the Retirement History Survey because of the emphasis in this study on male/female and black/white comparisons. The coverage of females in the Retirement History Survey is nonrepresentative, whereas the CPS has a better cross-section of the relevant population. For the second task, a data set that was constructed by Lazear (1983) was used. It is based on the Bankers' Trust *Corporate Pension Plan Study* (*1980*), covering about 200 plans.

11.1 Age, Tenure and Salary of the Typical Retiree

The 1979 CPS was used to impute the average age, tenure, and salary of the typical retiree in four race/sex categories. This task was less than straightforward because the relevant information is not reported in an appropriate form. Since the CPS is a cross-section, the date of retirement, and therefore age, tenure, and salary at the date of retirement, are not known for the group of individuals who are currently working. For the individuals who have already retired, neither tenure nor final salary on their career (or even last job) is reported. Thus, it is necessary to devise a method that estimates the requisite information from the cross-section.

The idea is to examine different cohorts and to infer from the distribution of individuals across retirement and employment-tenure classes what the retirement age and tenure must have been, using a variant of synthetic cohort analysis. The following example illustrates the basic ideas.

Suppose we are interested in the average level of tenure at retirement for some group and that only three age groups are relevant: No one retires before age 55, some retire at ages 55 and 56, and all are retired by age 57. The cross-section has workers and retirees at each age. So let us stratify the sample by age. None age 55 are retired, and their tenure on the current job is reported. Suppose that half have tenure of 20 years and half have tenure of 30 years. Although we cannot observe what happens to these individuals over the next year, we can examine the individuals who are currently 56 years old. In a steady state those individuals are identical to the current group of 55-year-olds, except that they are one year older. Suppose that half of the 56-year-olds are retired and of those who continue to work, three-fourths have tenure of 21 years, whereas only one-fourth have tenure of 31 years. That implies that three-fourths of those who retired before age 56 did so with 30 years of tenure and one-fourth did so with 20 years of tenure. Thus, $(1/2)(3/4) = 3/8$ of the population retire at age 55 with 30 years of tenure. Similarly, $(1/2)(1/4) = 1/8$ retire at age 55 with 20 years of tenure. Since all workers are retired by 57, it follows that $(1/2)(3/4) = 3/8$ of the labor force retire at age 56 with 21 years of tenure and that $(1/2)(1/4) = 1/8$ of the labor force retire at age 56 with 31 years of tenure. Given this information it is easy to calculate the expected level of tenure at retirement. In this case, it is

$$(3/8)30 + (1/8)20 + (3/8)21 + (1/8)31 = 25.5 \text{ years.}$$

The actual procedure is more complicated because there are many more age and tenure categories and because some workers take new jobs and others die. But the basic idea is the same. The procedure is applied to four groups: white males, white females, black males, and black females. The subset of the CPS sample analyzed consists of individuals who reported themselves either as retired or as currently working with valid information on job tenure, and who were from 55 to 76 years old.[2] The CPS reports whether individuals who are working are enrolled in a pension plan. We restricted the sample to those who were enrolled because there are large differences in employment status, tenure, and salary levels by pension enrollment.[3]

The next few pages begin with some definitions and describe the method used in more detail. The estimates are based on a counting algorithm and steady-state assumptions. Define marginal counts

$N(a,i)$: number of workers in the cross-section of age a who have i years of tenure.

$N(a,R)$: number age a who are retired.

and transition counts

$N_j(a,i)$: number of age a with tenure i who will have j years of tenure next year.

$N_R(a,i)$: number of age a with tenure i who retire during the year.

Ignoring unemployment, for transitions we have, for $i \geq 1$ either:

(1) $j = i + 1$: if the person remains on job

(2) $j = 1$: if the person turns over and obtains a new job

(3) $j = R$: if the person retires between years.

Finally, define

$N_D(a,i)$: number aged a and tenure i who die before age $a + 1$.

The following accounting identities apply in a steady state:

$$N(a,i) = N_{i+1}(a,i) + N_1(a,i) + N_R(a,i) + N_D(a,i).$$

A person must go to one of the four mutually exclusive and exhaustive classifications. Further

$$N(a + 1, i + 1) = N_{i+1}(a,i).$$

Persons found with one more year of tenure in the following year must be those who transited to that state between years. And

$$N(a + 1, 1) = \sum_i N_1(a,i) + N_1(a,R).$$

People with one year of tenure are those who changed jobs or who came out of retirement. Similarly,

$$N(a + 1, R) = \sum_i N_R(a,i) + N_R(a,R).$$

Those observed retired in the next year either transited to that state during the year or were retired earlier and remained retired. Therefore

$$(1) \quad \begin{aligned} N_R(a,i) &= N(a,i) - N_{i+1}(a,i) - N_1(a,i) - N_D(a,i) \\ &= N(a,i) - N(a + 1, i + 1) - N_1(a,i) - N_D(a,i). \end{aligned}$$

We seek to estimate $N_R(a,i)$. Both $N(a,i)$ and $N(a + 1, i + 1)$ are observed in the cross-section data. However, no data are available from a cross-section on transitions $N_1(.,.)$ or $N_D(.,.)$, so some assumptions are required to impute them.

Let $P(a,i)$ be the probability that a worker aged a with tenure i takes a new job and transits to state $i = 1$. Include R in the set $\{i\}$. Then

$$N_1(a,i) = P(a,i)\, N(a,i)$$

so

$$(2) \quad N(a + 1, 1) = \sum_i N_1(a,i) = \sum_i P(a,i)N(a,i).$$

If there are A age groups and T tenure classes (2) represents $A - 1$ equations in $(A - 1)T$ unknown $P(a,i)$. The marginal counts $N(a,i)$ are not sufficient to estimate $P(a,i)$, and therefore $N_1(a,i)$, without additional restrictions on $P(a,i)$. We know from other studies (see Mincer and Jovanovic 1981) that P is decreasing in i and probably in a as well.

To make things simple and computationally tractable we assume that $P(a,i)$ takes the form

$$P(a,i) = (\alpha_0 + \alpha_1 a + \alpha_2 a^2 + \beta_1 i + \beta_2 i^2 + \delta ia)(1 - R) + R(\delta_0 + \delta_1 a),$$

where $R = 1$ if the person is retired and $R = 0$ if not.

Define $N(a) = \Sigma_{i=R} N(a,i)$ as the working population of age a and $N(a,R)$, as before, as those retired. Then, substituting for $P(a,i)$ in (2) and summing yields

$$\text{(3)} \quad N(a + 1, 1) = \alpha_0 N(a) + \alpha_1[aN(a)] + \alpha_2[a^2 N(a)] + \beta_1[\sum_i iN(a,i)] + \beta_2[\sum_i i^2 N(a,i)] + \delta[a\sum_i iN(a,i)] + \delta_0 N(a,R) + \delta_1 aN(a,R).$$

Treat (3) as a regression equation, in which the observed counts $N(a + 1, 1)$ are regressed on observed variables $N(a)$, $aN(a)$, . . ., etc. across age groups. There are eight unknown parameters in this regression, so if $A \geq 9$, this regression can be estimated.

In our data $A = 21$, so there are only 13 degrees of freedom. Therefore, the individual parameters (α, β, λ, δ) are not estimated precisely. In addition some of the regressors are collinear. Nevertheless, we get unbiased estimates $\hat{P}(a,i)$. From these we obtain unbiased estimates of $N_1(a,i)$, from

$$\hat{N}_1(a,i) = \hat{P}(a,i)N(a,i).$$

A similar procedure works in general for estimates of $N_D(a,i)$. However, we find that the data are too thin to obtain meaningful results for the relationship between death probabilities conditional on both age and tenure. We therefore assume

$$\hat{N}_D(a,i) = P^*(a)N(a - 1, i)$$

where $P^*(a)$ is the 1979 age-specific death rate for this race-sex class. We know that there is a strong negative association between work and death so $\hat{N}_D(a,R)$ is likely to be biased from this procedure. The biases with respect to i are less clear-cut, though it is probable that $\hat{N}_D(a,i)$ for large i is upward biased, since people who are currently working and with long tenure are likely to be healthier than average.

From the identity above, $N_R(a,i)$ is estimated from

$$\text{(4)} \quad \hat{N}_R(a,i) = N(a,i) - N(a + 1, i + 1) - \hat{N}_1(a,i) - \hat{N}_D(a,i).$$

Now $\Sigma_i N_R(a,i)$ is the total number of people aged a who retire and $\Sigma_{ai}\Sigma N_R(a,i)$ is the total number who retire in the whole population at any age. Therefore,

$$\text{(5)} \quad n(a) = \Sigma_i N_R(a,i)/\Sigma\Sigma N_R(a,i)$$

is the probability of retiring at age a, given that death does not occur prior to retirement, and

(6) $$\Sigma_i\ an(a) = E(\text{age of retirement}) = Ea_R.$$

Similarly, $\Sigma_a\ N_R(a,i)$ is the number of people who retire after i years of tenure, so

(7) $$m(i) = \Sigma_a\ N_R(a,i)/\Sigma\Sigma\ N_R(a,i)$$

is the conditional probability of retiring at tenure i given that death occurs after retirement.

Therefore

(8) $$Ei_R = E(\text{tenure at retirement}) = \sum_i im(i).$$

Before turning to the estimates, some qualifications are in order.

1. If $\hat{N}_D$ is biased upward for larger i, then there is a downward bias in $m(i)$ for large i (and upward bias in $m(i)$ for small i). Therefore Ei_R is probably biased down on this account. However, this source of bias is likely to be small.

2. Even though the estimates of $N_R(a,i)$ are no doubt imprecise, the usual sampling theory suggests that $E(a)$ and $E(i)$ are better measures than any of their components, through the law of large numbers. Now if the pension formulas were linear in a and i, these means are sufficient statistics for our problem. However, these schemes are not linear. Therefore in predicting expected pensions and pension wealth from each plan, it would be preferable to take weighted averages across (a,i) pairs rather than taking the outcome for the average person. The preferred alternative is simply not feasible with these data.

3. The imputation procedure assumes no cohort effects. This is dictated by a cross-section since it is well known that cohort and age effects cannot be identified in a cross-section except through arbitrary assumptions. The formulas above make the strong steady state assumption that for $a < a'$ people who attain age a' at $(a' - a)$ periods in the future will behave "as if people age a' are behaving today (1979)."

We know that the age of retirement has shown a secular decline for males in the post–World War II period. Increasing wealth, changes in tax laws and in the social security system, as well as changes in family composition and yet other factors are all contributory causes. If these trends continue, then $E(a)$ is likely to be smaller in the future than our estimate: The average age of retirement for older cohorts in our sample was surely larger than our estimate. On the other hand, those issues are reversed for females, given the large increase in female labor force participation in recent decades. Since our estimates for females are conditioned on working, it is probable that cohort bias of this sort is far less important for women than for men.

The influence of cohort effects on expected tenure is less clear-cut. There are little data on secular changes in tenure on which to base an a priori judgment. If retirement continues to occur at younger ages this is likely to reduce tenure at retirement as well. However, the relation between age and tenure is noisy, so though there may be cohort bias in $E(i)$ qualitatively similar to that of $E(a)$, it is likely to be quantitatively smaller. Changing labor force behavior of women and conditioning on labor force participants again makes these considerations less important for women; if anything, the cohort bias for women tends to go in the opposite direction than for men.

4. This procedure is based on actual counts in the CPS tape for $N(a,i)$. If all cohorts were the same size, and if sample data reflected this exactly, then, on the usual synthetic cohort assumptions, everything works out correctly. However, some adjustments are necessary if either birth cohorts vary in size (which they do) or sample sizes vary randomly with age. The following approach, which is incorporated into the calculations, corrects the problem.

Define $N(a)$ as the total number of individuals in the sample of age a. We normalize everything in terms of $N(55)$. If this were a panel, then the difference between $N(55)$ and $N(56)$ reflects only deaths during the year. But in our synthetic panel, $N(56)$ may deviate from $N(55)$ because of real differences in cohort sizes or random sampling differences across age groups. However, death rates are known with accuracy, so an estimate of the corrected age 56 sample can be easily obtained. In fact, $P^*(a)$, defined above, does exactly that. Thus, as an initial condition set

$$\hat{N}(56) = N(55)[1 - P^*(55)].$$

Then the following recursion applies for $a > 56$

$$\hat{N}(a) = \hat{N}(a - 1)[1 - P^*(a - 1)].$$

The ratio of $N(a)/\hat{N}(a) = \lambda(a)$ reflects random sampling size or cohort size differences. To correct our estimates for these factors, it is necessary only to divide all $\hat{N}_R(a,i)$ by $\lambda(a)$. Then equations (5)–(8) follow as written.

This discussion points to another possible source of bias that we have ignored, nonretirement transitions out of the labor force. This is likely to lead to relatively small error for the aged population we study here because these transitions are relatively minor among older workers.

5. In the data actually used we identify 21 age classes, $a = 55, \ldots, 75$, and 54 tenure classes, $i = 1, \ldots, 54$. Since the sample consists of some 1,600+ persons, many of the $N(a,i)$ cells are very small, and

many are empty. To deal with this problem we aggregated across tenure intervals and then interpolated tenure-specific totals by regression: In particular, define

$$x(a,I_j) = \sum_{i \epsilon I_j} \hat{N}_R(a,i).$$

After inspection of the raw cells, 11 such sums were defined for each age: $I_1 = (1)$, $I_2 = (2,6)$, $I_3 = (7,11)$, $I_4 = (12,16)$, $I_5 = (17,21)$, $I_6 = (22,26)$, $I_7 = (27,31)$, $I_8 = (32,36)$, $I_9 = (37,41)$, $I_{10} = (42,47)$, $I_{11} = (48,54)$. Define $\bar{I}_j$ as the midpoint in years of the jth I_j interval.

We fit the regression

$$\text{(9)} \quad x(a,I_j) = b_0 + b_1a + b_2a^2 + b_3\bar{I}_j + b_4\bar{I}_j^2 + b_5\bar{I}_ja + b_6B + b_7F + b_8D + b_9(aB) + b_{10}(aF) + b_{11}(D \cdot B) + b_{12}(D \cdot F)$$

to the aggregated data for purposes of smoothing and interpolation. The variables B for black, F for female, and D for $i = I_1$ are dummies. Then

$$N_R^* (a,i) = \hat{b}_0 + \hat{b}_1a + \hat{b}_2a^2 + \hat{b}_3i + \hat{b}_4 i^2 + \hat{b}_5ia$$

was used to calculate the distributions $n(a)$ and $m(i)$ used for our estimates of Ei_R and Ea_r above. Appendix A reports the regression in (9).

The estimates are contained in table 11.1.

Expected age of retirement of persons covered by private pensions is remarkably uniform across race and sex groups. Remember that these numbers are conditioned on labor force participants as well as pension eligibility. This explains the lack of appreciable differences between males and females. While older females are far less likely to participate in labor market activity than males, those that do participate show average retirement ages that are similar to those of men. In fact, Ea_R is slightly larger for women. Since estimated Ea_R is close to the early retirement age under social security, the somewhat larger value for women may reflect known smaller coverage and experience under

Table 11.1 **Estimated Age, Tenure and Salary at Retirement**

	$\hat{E}a_R$	$\hat{E}i_R$	ES
White:			
Male	62.1	22.0	\$17,970
Female	63.2	21.8	11,414
Black:			
Male	63.0	15.3	13,194
Female	65.9	16.8	10,754

social security than for men. The somewhat larger difference between black males and females may be due to these same factors, as well as to the fact that labor force participation of black women historically has exceeded that of white women. Whatever factors affect these differences in participation rates apparently also makes black women retire later in life.

The most surprising result in table 11.1 refers to the sex and race differences in expected tenure at retirement. For whites we find that expected tenure on the job held at retirement is virtually the same between the sexes and is a remarkably long 22 years in length. As a check, this estimate is similar to average tenure levels for those still working in the CPS data. Recent work on job tenure patterns for males shows a characteristic pattern that most job mobility occurs at younger ages. By middle age most job mobility that will occur over a lifetime has already taken place, so it is not surprising that for the older male workers in our sample the average tenure at retirement is 22.0 years. The result for women seems surprising at first glance, but is less so when it is recalled that these calculations refer to *working women* at age 55 and older. The estimate reflects the fact that a significant number of women are permanently attached not only to the labor force but also to their place of work, either through their whole careers or certainly subsequent to reentry into the labor market after childbearing years.

These similarities between sexes are apparent among blacks as well as whites in table 11.1. However, the difference between the races is substantial. Taken at face value, these differences must reflect much greater job mobility among older blacks than among older whites. While there is some evidence that job and labor market instability is larger for blacks than for whites at younger ages, we are unaware of confirming evidence on these differences between races among older workers. It should be noted in this connection that our sample is much smaller for blacks than for whites, and the individual $N(a,i)$ cells are correspondingly thinner. Hence, the smoothness procedures used and described above may ultimately account for these differences: Certainly any confidence interval on these estimates would be much larger for blacks than for whites, based on sampling variation alone. In fact, the results for blacks are sensitive to the specification of equation (9), which causes us concern. This fact must be borne in mind when interpreting the black/white differences below. Still, there is nothing in the procedure used that would by itself produce this point estimate, and the similarity between black men and black women is not automatically implied by our method.

The last bit of information necessary to perform the simulations is the final salary at time of retirement, since many plans are geared to these figures. The estimate is based on a standard earnings regression for each race/sex group of the form

$$\text{(10)} \qquad \text{Earnings} = c_0 + c_1 a + c_2 a^2 + c_3 i + c_4 i^2 + u,$$

where the c's are regression parameters and u is regression error. After fitting this equation for each age-sex group, the estimated average salary for the average person is estimated by evaluating it at $a = Ea_R$ and $i = Ei_R$. The regressions are contained in Appendix A. The earnings estimates are shown in the third column of table 11.1, labeled ES.[4]

The salary regression does not include the usual elaborate list of controls such as education, marital status, occupation, and the like because we are not particularly interested in this study of the partial effects of such variables. Hence the coefficients on the age and tenure variables capture the effects of variations in these other variables that are correlated with age and tenure. This is conceptually appropriate for our purposes because we desire an estimate of mean final salary for each race-sex group over all education, occupation, industry, and marital status classes. A more elaborate regression would require reweighting these other effects by relative sample proportions: The regression above is self-weighting in this respect and is sufficient for the problem at hand. Also, the regression has been specified in terms of earnings levels rather than the usual log of earnings. A log transform is known to provide a better fit when all age groups are included in the sample, but there is no compelling reason for using that transform for the older people in our sample since it is well known that much of the curvature in life-cycle earnings patterns occurs at younger ages. Furthermore, it is the level and not the log of earnings that is relevant for pension determination, so we also avoid the questionable σ_u correction for error variance in transforming the log to the level by this procedure.

The percentage differences between white men and women at Ea_R and Ei_R in table 11.1 conform to the percentage differences in earnings found in the population as a whole. This is rather surprising because the women in our sample exhibit the same mean age of retirement and tenure at retirement as men do, and it is generally thought that the raw difference in earnings between men and women in the population at large is related in some way to differences in labor force activity over the life cycle. No doubt many of the women in our sample reentered the market after childbearing. Whatever the case, they never caught up with the men. This is both surprising and worthy of more detailed investigation. The same relative pattern is repeated among blacks, but at a much lower level.

One final qualification is necessary concerning these salary estimates. The salary observations are censored by the retirement decision itself.

Thus the older individuals in the sample who were working found it in their interest not to retire because their wage prospects were evidently larger than their opportunity cost of leisure. People who continue to work are generally healthier than average and many have superior earnings prospects, so the observed wages of older-than-average workers in our sample are likely to be larger than the wage prospects available to workers of these ages who chose not to work. Therefore expected salary at age of retirement calculated above probably is too large for the average worker.

11.2 Pension Values of the Typical Retirees

Given the information in table 11.1, the pension of these typical retirees can be calculated from information on pension benefit formulas. The information used comes from a data set generated by Lazear (1983). A description follows.

The data for Lazear's analysis were constructed using the Bankers' Trust *Corporate Pension Plan Study (1980)*. The study consists of a detailed verbal description of the pension plans of over 200 of the nation's largest corporations. The data set applies to approximately 10 million workers, and this comprises about one-fourth of the entire covered population. The major empirical task was to convert the verbal descriptions into machine-readable data. This required setting up a coding system that was specific enough to capture all of the essential detail associated with each plan. It was then necessary to write a program which calculates the present value of pension benefits at each age of retirement.

Pension benefit formulas are of three different types. The two most common fall under the rubric of defined benefit plans, which specifies the pension flow as a fixed payment determined by some formula. The *pattern plan* awards the recipient a flat dollar amount per year worked prior to retirement. The *conventional plan* calculates the pension benefit flow from a formula which depends upon years of service and some average or final salary. In contrast to the defined benefit plans are defined contribution plans in which the employer (or employee) contributes a specified amount each year during work life to a pension fund. The flow of pension benefits that the worker receives upon retirement is a function of the market value of that fund. The defined contribution plan is much less frequently used than is either the pattern plan or conventional plan. Only defined benefit plans are used here.

Some plans do not permit the individual to receive early retirement benefits or only permit early retirement up to a given number of years before the normal date. This means that in order to perform the necessary comparisons, some plans had to be deleted because age or tenure values in table 11.1 violated restrictions of the plan. Less than 15% of plans were deleted for this reason.

Most plans have restrictions on the maximum amount which can be accrued, and many provide for minimum benefits. Additionally, a number reduce pension benefits by some fraction of the social security benefits to which some basic class is entitled. Moreover, a number of plans provide supplements for retirement before the social security eligibility age. Sometimes these supplements relate directly to social security payments; at other times they depend upon the individual's salary or benefit level. Other restrictions have to do with vesting requirements, with the maximum age at which the individual begins employment, and with the minimum number of years served before the basic accrual or particular supplements are applicable. The accrual rate, or flat dollar amount per year to which the individual is entitled, is often a nonlinear function of tenure and salary, and these kinks had to be programmed into the calculations.

This permits computation of the flow of retirement income in each of these plans, for each of the four typical workers. To get present values of the pension flows, a 10% discount factor was used. Finally, in performing the actuarial correction, it was necessary to choose a life table. The 1979 United States Vital Statistics tables were used. The choice of table turns out to be the least crucial part of the analysis. Values do not vary greatly from year to year, and discounting makes unimportant whatever small differences there are among tables.

Each of our four typical individuals was run through 172 of the plans for which qualification criteria were met. The expected present value of retirement benefits (in date of retirement dollars) was estimated for each of those individuals in each of the plans. Table 11.2 provides some summary statistics on the results of that simulation.

Table 11.2 **Pension Present Value for Typical Retiree (All Pensions) (N = 172)**

Group	Mean	Expected Pension*	S.D.	Max	Med	Min
White males	$30,284	$18,412	$17,860	$142,111	$28,422	$862
White females	23,527	11,340	11,152	87,193	22,000	833
Black males	17,396	9,550	9,545	78,342	16,067	833
Black females	15,997	6,558	8,771	59,723	15,105	740

*Expected pension is defined as the raw probability (from table 11.6) times the mean pension.

11.3 Results

There are a number of interesting findings that come from this analysis. Let us turn first to the question that was posed at the outset, namely, does pension wealth exacerbate inequality? Recall that there are two aspects to the question. The first relates to the probability that a worker in a given demographic category has a pension; the second regards the expected pension value for pension plan participants. The first was investigated by using the CPS data to estimate a linear probability model. In table 11.6, the dependent variable is a dummy equal to one if the individual in question participates in a private pension plan. The sample consists of all working individuals between 55 and 76 years old with tenure reported.

A look at the coefficients in table 11.6 makes it appear as if blacks and females do not differ from white males in terms of their probabilities of participation in a pension plan. (Both coefficients are essentially zero.) Appearances are deceiving because earnings are held constant. Earnings have a strong positive association with pensions, and since blacks and females have lower earnings than white males, most of the difference can be accounted for by differences in earnings. While women and blacks who earn the same wages as white males are likely to enjoy the same pension participation status, women and blacks are unlikely to earn the same amount as white men.

The more important statistic for this analysis is the raw probability of participation in a pension plan. Those probabilities are reported in table 11.6 as well. White males have the highest probability of participating in a pension plan while other groups, especially black women, are substantially behind. These probabilities will play an important role in the subsequent discussion.

To examine the second question, namely, how do pensions vary among participants by race and sex, we call on the information in tables 11.2–11.5. First, compare the first and last columns of table 11.5.

The first column reports the ratio of pension value means from table 11.2 for the relevant group so that the first entry is 23,527/30,284. The fifth column reports the ratio of salary means from table 11.1 so the first entry is 11,414/17,970.

Table 11.3 **Pension Present Value for Typical Retiree Defined Benefit Pattern Plans (N = 48)**

Group	Mean	S.D.	Max	Med	Min
White males	$23,277	$6,822	$40,483	$23,724	$ 4,486
White females	22,318	6,502	39,105	22,459	4,333
Black males	15,067	4,280	26,612	15,000	13,750
Black females	15,110	4,285	26,817	15,105	3,110

Table 11.4 **Pension Present Value for Typical Retiree Defined Benefit Conventional Plans (N = 124)**

Group	Mean	S.D.	Max	Med	Min
White males	$32,991	$20,042	$142,111	$31,264	$862
White females	24,032	12,520	87,193	22,000	833
Black males	18,260	10,833	78,342	16,523	833
Black females	16,342	8,899	59,723	15,105	740

Table 11.5 **Ratios of Means**

	Pension Values and Final Salary				
Groups	All	Pattern	Conventional	Expected Pension (ALL)	Final Salary
White female/white male	.776	.958	.728	.615	.635
Black male/white male	.574	.647	.553	.518	.734
Black female/white female	.679	.677	.680	.578	.942
Black female/white male	.919	1.002	.894	.687	.815

NOTE: Black male pattern/white male conventional = .456. Black female pattern/white female conventional = .628.

First consider black males and white males. The second row of table 11.5 is relevant. Note that the ratio of the mean salary at retirement for these groups is .734 and that the ratio of pension benefits is .574. If workers were distributed randomly across the plans (which they are not), then the existence of pensions would tend to increase black/white male inequality. This is true for two reasons. First, as reported earlier, blacks are less likely to have pensions than whites. Second, given that black males do receive a pension, they receive a considerably smaller amount in pension benefits than whites. A measure that combines both aspects is the ratio of expected pension, defined as the ratio of the mean pension times the raw probabilities from table 11.6. That number is reported in the fourth column as .518 so pensions appear to exacerbate inequality. (Recall, however, that results for blacks are not robust to specification.) The magnitudes, although not astronomical, are not trivial either. For white males, the present value of pension wealth averages somewhat less than 2 years' income. For black males, the average value of pension wealth is somewhat less than 1 year of income.

Because of the significant salary differences, conventional plans, which base the pension on final salary, exacerbate the black/white male differences. Tables 11.3 and 11.4 split the sample of plans into pattern and conventional plans. The second column of table 11.5 reports the ratios of means given in table 11.3 and the third column reports the ratios of means from table 11.4.

Table 11.6 Probability of Participation in a Private Pension Plan

Variable	Coefficient	Standard Error
Constant	1.559	.121
Annual earnings (1000s)	.01856	.00098
Black	−.0019	.0298
Female	.0035	.0186
Age	−.0205	.0019

Raw probabilities of participation:

White males = .608 .
White females = .482
Black males = .549
Black females = .410

A comparison of column 3 with column 1 in table 11.5 reveals that the ratios in the third column are smaller for all groups that do not include black females, because salary levels are important for computation of conventional pension plans. Black males who have conventional plans are at even more of a disadvantage relative to white males in the same plans because their earnings are lower. Perhaps more important is that blacks and whites are unlikely to be found in similar proportions in the two plan types. Pattern plans are more typical for production workers, whereas the conventional plan is the norm for management and white-collar workers. To the extent that blacks are overrepresented among pattern plans, pension inequality is even more pronounced. At the extreme, if all black males had pattern plans and all white males had conventional plans, then the ratio of the pension value means would be .456, whereas salary ratios are .734.

The findings for black females and white females are even more striking. The salary column of table 11.5 reveals that the ratio of black to white female salary is .942, whereas the ratio of pension value is only .679. If all white females were in conventional plan occupations and all black females were in pattern plan occupations, the pension inequality would be even greater. That ratio would be .628 instead of .679. The reason for the difference is that conventional plans are generally more lucrative than pattern plans, except at very low salary levels. Similarly, the ratio of expected pension for these groups is .578, implying even greater inequality because black females are less likely to be enrolled in a pension plan at all. No matter how we measure it, pensions appear to increase black/white inequality relative to that estimated by salary measures.

The male/female comparisons are less clear-cut. Effects go in opposite directions. As reported above, female workers are less likely to be enrolled in a pension plan than male workers, but if they are enrolled, the white females do well relative to their male counterparts. The first row of table 11.5 contains the relevant information. The ratio of final salary of white females to white males is .635 whereas the ratio of pension values is .776. This implies an equalizing effect of pension benefits. Part of this results from the fact that defined benefit plans are not sex-specific, so that women, with longer life expectancies, do better than men. But this cannot account for the large difference between .776 and .635.

The reason why women do so well in pension benefits can best be understood by examining the distinction between pattern and conventional plans. Note that women are almost on par with men in terms of pension benefits received in pattern plans. This results from one factor: Pattern plans depend only on years of service, and in that respect, the women who are working at age 55 are quite similar to men. This large value of tenure maps into high pension flows in the pattern plan. (Because of the actuarial unfairness of the plan, it could actually have gone the other way. Since tenure levels are close to comparable, the longer life expectancy of females could have made their pattern plan pensions worth more than those of males.)

The equalizing effect of pensions is offset almost exactly by the fact that fewer women than men are enrolled in pension plans. From table 11.6, white men had a probability of receiving a pension of .608, whereas white women had a probability of .482. It is useful, therefore, to compare expected pensions. The ratio of expected pension for white females to white males is .615 from the last column of table 11.5. Thus, pensions leave white female/white male wealth inequality unaltered.

The same pattern is displayed for blacks. The black female's final salary is 81% that of the black male in this sample, and the mean black female's pension benefit is 92% of the mean black male's pension. But expected pension ratios tell the opposite story. Since black females are much less likely to be pension recipients, the ratio of expected pension benefits is .687. Thus, pensions increase male/female inequality substantially among blacks.

This conclusion is strengthened somewhat when it is recalled that these women are not a random sample of the overall population of women. Since a larger proportion of women will have dropped out of the labor force before reaching age 55, and since it is likely that those individuals have very small pension wealth, the numbers presented in the last paragraph tend to understate the disequalizing effect of pensions in the overall economy.

Other interesting findings are worthy of discussion. Most obvious is that there is much more variation in the benefits provided by conventional plans than in those provided by pattern plans. A comparison of tables 11.3 and 11.4 is instructive. For all four groups, the standard deviation is much larger for conventional plans. Similarly, with the exception of white males, medians are about the same across plan types, but the maximum and minimum values are much more extreme in the case of conventional plans.

Variance in pension benefits received in conventional plans depends on two factors. The first is that for a given salary, companies differ substantially more in their conventional pension formulas than in their pattern plan formulas. Second, a positive correlation between the firm's average salary and generosity of the pension formula contributes variance to benefits received. Although it is conceivable that the two types of variation will offset one another, it is unlikely. There is already some evidence of a positive correlation between average salary in the firm and the generosity of pension benefits (see Asch [1984] and the salary coefficients in table 11.6).

Before concluding, we should mention that another study addresses the same questions as we do but obtains somewhat different results. McCarthy and Turner (1984) find that blacks actually have higher pensions than whites do, both in terms of pension flow and pension wealth (see their table 1). They use the Survey of Private Pension Benefit Amounts, a data set that permits pairing of individuals with the actual pensions they receive. On the face of it, this data set is superior to those that we have used. But their findings leave some grounds for doubt on that score. In particular, it is difficult to believe that blacks have higher pensions than whites because even in the group of pension plan participants, the average final salary of a black male is only 63.5% of the white male (see our table 11.5). Since many pension plans depend on final salary, even if tenure at retirement did not differ between groups, the pension flow ratio would mirror the salary ratio. It is important to reconcile the two sets of results, but McCarthy and Turner are unable to make their data available to the public, so their results cannot be replicated.

11.4 Conclusion

The existence of pension plans appears to contribute to black/white inequality but leaves male/female inequality unchanged among whites. Even though females are less likely to receive pensions than males, those females who do receive pensions tend to receive relatively generous ones. Of course, the average pension that the typical retiring female receives is well below that received by the typical male retiree. But the difference is not as pronounced as male/female differences in salary. Among blacks, pensions exacerbate sex differences, mainly because black women are only about 75% as likely to receive pensions as black males.

Appendix
Regression Results

Var	Eq. (9)	Eq. (10)			
Dep. Var.	$= X(a,I_j)$	White Male Earnings	White Female Earnings	Black Male Earnings	Black Female Earnings
Constant	−12.80	66,077	13,777	28,493	214,280
	(8.59)	(61,992)	(40,979)	(121,344)	(351,051)
a (age)	.488	−1,199	−102	−404	−6,738
	(.262)	(2,028)	(1,377)	(3,900)	(11,913)
a^2(age^2)	−.0043	5.47	−.24	1.34	54.6
	(.0020)	(16.5)	(10.8)	(31)	(101)
$\bar{I}_j$ (tenure)	−.0570	279	305	441	327
	(.0484)	(104)	(77)	(261)	(262)
$\bar{I}_j^2$ (tenure^2)	−.00021	−1.82	−3.38	−8.17	−7.45
	(.00034)	(2.48)	(2.01)	(5.91)	(6.85)
$\bar{I}_j a$	.00092				
	(.00068)				
B	−2.96				
	(1.41)				
F	−1.69				
	(1.41)				
D	.556				
	(.445)				
aB	.038				
	(.021)				
aF	.024				
	(.021)				
$(D)(B)$	−.258				
	(.457)				
$(D)(F)$	−.613				
	(.457)				
df	911	937	509	65	42
R^2	.038	.077	.112	.074	.076
Mean of dependent variable:		18,855	10,397	13,857	10,206

Notes

1. E.g., see are Lillard (1977), Rosen (1977), Lazear (1979), and Lillard and Willis (1978).

2. We ignore those who retire earlier than 55 because it is likely that only a very small number of workers with pensions retire before age 55.

3. For example, not enrolled black women earn an average of $3,471 per year, whereas enrolled black women earn $10,206.

4. Note that although the earnings regressions are imprecise, the estimates derived from them and used in table 11.1 are close to the unconditional mean for each group.

References

Banker's Trust Company. 1980. *1980 study of corporate pension plans.* New York: Banker's Trust.

Lazear, Edward P. 1979. The narrowing of black-white wage differentials is illusory. *American Economic Review* 69:553–64.

———. 1983. Pensions as severance pay. In *Financial aspects of the United States pension system,* ed. Zvi Bodie and John Shoven. Chicago: University of Chicago Press.

Lillard, Lee A. 1977. Inequality: Earnings vs. human wealth. *American Economic Review* 67:42–53.

Lillard, Lee A., and Willis, Robert. 1978. Dynamic aspects of earnings mobility. *Econometrica* 46:985–1012.

McCarthy, David, and Turner, John A. 1984. Earnings comparisons and race-gender differences in private pension benefits. Washington: United States Department of Labor. Mimeographed.

Mincer, Jacob, and Jovanovic, Boyan. 1981. Labor mobility and wages. In *Studies in labor markets,* ed. Sherwin Rosen. Chicago: University of Chicago Press.

Rosen, Sherwin. 1977. Human capital: A survey of empirical research. In *Research in labor economics,* ed. R. Ehrenberg. Greenwich, CT: JAI Press.

Comment Sylvester J. Schieber

The question that Lazear and Rosen address here is whether private sector pensions exacerbate compensation inequality across groups—specifically by race and gender. There are two aspects to consider. First, the probability of receiving a pension is not random across groups. The second aspect is conditional on receiving a pension and whether systematic variation exists in benefit amounts received based on race or gender. The authors focus on the latter.

Organizationally, Lazear and Rosen undertake their analysis in two stages. They first determine the average age, tenure, and salary char-

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acteristics at retirement for each race-gender group on the basis of the age, tenure, and salary characteristics estimated in the first stage. They then estimate the pension that each group prototype person would receive under 172 different pension plans. They use the results of this exercise to assess the distributional effects of pensions.

The first stage of their paper is an interesting exercise in adapting data to an analytical problem for which appropriate, straightforward data generally are not available. I refrain from commenting on the mechanics of this segment of the paper because I think the generated results are inappropriate for addressing the question of the role pensions play in distributing income. I concur with the view that pensions may exacerbate wage inequality, but they are more likely to do so on a coverage than a benefit structure basis. Lazear and Rosen did not address the pension coverage, participation, and benefit receipt issue, but it is well known that pension participation rates are significantly lower at the bottom of the earnings than at mid-or-upper earnings levels. To have ignored this aspect of the question is a general limitation that weakens the remaining analysis.

In the estimates of final average salary/age/tenure characteristics for each of the race-gender groups shown in Lazear and Rosen's table 11.1, the authors include pension plan participants and those not enrolled in pension plans. Their reported estimate found that blacks' tenure in their terminal job was less than half that of whites—11 years versus 23 years. They estimated that the terminal salaries varied significantly by race and gender. For example, white males' final salaries were estimated to be about two-thirds higher than those of both white women and black men. At the same time, while estimated black males' final salaries were roughly equal to those of white women, they were about twice those of black women.

My concern is that the authors' tenure and salary estimates tend to exaggerate differences that might exist across the groups studied. The classes of workers participating in pension plans are more homogeneous at or near retirement than the authors' analysis suggests. To illuminate this point, I have looked at the same 55–75-year-old workers from the May 1979 Current Population Survey that Lazear and Rosen used in their analysis. Consider, for example, the difference in the percentage of each relevant race-gender group working full-time by whether or not the group was participating in a pension plan. Table 11.C.1 shows that, across the four groups studied, pension participants consistently were more likely to be working full time. Also, there was only one-third the variation in the percentage of pension participants, compared to nonparticipants, working full time.

Though full-time employment status is not a precondition for participation in a pension plan, table 11.C.2 does indicate that a strong cor-

Table 11.C.1 Pension Participation Status of 55 to 75 Year-Old Full-time Workers

	Percent Working Full Time		
	All Workers	Pension Non Participants	Pension Participants
White Males	86.3%	76.4%	96.1%
White Females	66.7	53.9	83.3
Black Males	82.5	69.9	92.1
Black Females	55.1	37.2	84.0

SOURCE: May 1979 Current Population Survey

Table 11.C.2 Pension Participation 55 to 75 Year Old Workers

	Percent Participating in a Pension Plan	
	Part-time Workers	Full-time Workers
White Males	14.5%	56.0%
White Females	21.9	54.4
Black Males	25.9	63.5
Black Females	13.7	58.5

SOURCE: May 1979 Current Population Survey

Table 11.C.3 Estimated Final Job Tenure and Median Attained Tenure of Workers Aged 55 to 75 by Pension Participation Status

	EiR	Pension Nonparticipants*	Pension Participants*
White Males	23.1	11.1	19.5
White Females	22.9	6.3	13.8
Black Males	10.8	10.6	21.0
Black Females	10.6	7.2	16.2

*SOURCE: May 1979 Current Population Survey

relation exists between full-time employment and participation status. Table 11.C.2 also suggests that the prevalence of pension participation among blacks actually exceeds that for whites among full-time workers within the age cohorts considered here. In any event, it is clear that from a pension participation perspective, far more homogeneity exists across the four groups of full-time workers than from full- to part-time workers in any combination.

The incentives built into pension plans to discourage worker turnover also differentiate pension participants from nonparticipants. Comparison of group median tenures in table 11.C.3 indicates that among older pension participants blacks have attained longer tenures than whites and men longer tenures than women. This is a different result than that suggested by the Lazear-Rosen estimates also shown in the table 11.C.3. Analyzing pension benefit distributions based on tenures that do not reflect reasonable periods of participation under the plans biases the results. The low tenure estimates for blacks from the Lazear-Rosen model used here would have significantly reduced the estimated benefits for blacks and inflated the relative benefits estimated for white women.

The other crucial variable for determining pension benefits under most defined benefit plans is final salary. The estimated final salaries from the Lazear-Rosen model are compared in table 11.C.4 to estimated median earnings derived from the May 1979 CPS. Again the model exaggerates the group salary differences across the racial groups when the Lazear-Rosen estimates are compared with median earnings levels of pension participants from the May 1979 CPS. It is only when white males are compared to white females that the relative differences in the two sets of estimates are similar. To the extent the model systematically underestimates final salaries of blacks, it also would tend to exaggerate the authors' conclusion that "the existence of pension plans contributes to black/white inequality."

Pension plan design is regulated by the Internal Revenue Code that limits the ability of plans to discriminate among participants on the basis of salary. A pension benefit structure that provides relatively higher benefits to upper-income beneficiaries can only do so within the strict confines of the IRS integration regulations. The regulations recognize social security's redistributive nature and allow an employer to take partial advantage of the relatively higher benefits provided to lower-wage workers by social security in designing the pension benefit formula. But in no case is the progressive structure of social security to be fully offset by the pension benefit structure. The result is that the combined benefit structure, even for highly integrated plans, should still be somewhat progressive. Only 11% of the final pay plans and 23% of the career average plans included in the Bankers' Trust Survey were not integrated. The mere existence of integrated plans, and especially their prevalence included in the Lazear-Rosen analysis, suggests that social security should be included in further analytic efforts of the distributional issues addressed here.

Table 11.C.4 Estimated Final Salary and Median Earnings of Workers Aged 55 to 75 by Pension Participation Status

	ES	Pension Nonparticipants*	Pension Participants*
White Males	$17,830	$9,663	$16,300
White Females	10,680	4,816	9,524
Black Males	10,118	6,119	12,501
Black Females	5,109	3,284	8,564

*Source: Derived from the May 1979 Current Population Survey.

In their simulation of pattern plans the authors assume the same worker age, tenure, and final salary variations across the race-gender groups that they used for their simulations of conventional plans. A basic characteristic of pattern plans is that they provide almost no variation in benefit levels across groups of workers with equal tenure. For such a benefit structure to be acceptable to a participant population, there usually would have to be minimal variation in final salaries under the plan. Otherwise the plan would play an inconsistent role in maintaining the standard of living for workers retiring under the plan. Most pattern plans operate in unionized settings where the ability to discriminate on the basis of gender or race for purposes of setting salaries is quite limited. In addition what salary discrimination that previously might have existed has been obviated by the Civil Rights Act of 1964.

Furthermore, pattern plans often have only a tenure requirement as the criteria for retirement eligibility prior to the normal retirement age specified in the plan. For example, the "30 and out" provisions in the UAW plans permit retirement with unreduced benefits at any age for workers with 30 years of service. The prevalence of such criteria means that many workers retiring under "30 and out" plans tend to have the full tenure needed to fulfill this provision. This phenomenon is accentuated because many of these plans include portability provisions that grant a worker credit for tenure under similar plans with other employers. So, an auto worker shifting from one firm to another will get credit from the second employer's plan for service under the first.

The combination of limited salary variation and relatively consistent tenure patterns under pattern plans tends to suppress variation in benefits. A more careful specification of the characteristics of individuals participating under both conventional and pattern plans would likely result in different comparative results than those presented in the Lazear-Rosen analysis.

The analysis conducted here should be expanded and refined in several regards. First, any analysis of the distributional effects of retirement programs certainly should include social security. Second, much more attention should be paid to the distributive effects that might arise because some workers do not participate in pension plans. Third, a more precise specification of attained age, tenure, and final salary characteristics under the various types of plans is critical for such an analysis. Finally, I think the basic question posed here can only be adequately addressed with better data. Regrettably, such data exist at the Department of Labor but will probably never be made available for public use.

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